Board of Governors of the Federal Reserve System

International Finance Discussion Papers

Number 652

November 1999

MONETARY POLICY'S ROLE IN EXCHANGE RATE BEHAVIOR

Jon Faust and John H. Rogers

MONETARY POLICY'S ROLE
IN EXCHANGE RATE BEHAVIOR

Jon Faust and John H. Rogers*

Abstract: While much empirical work has addressed the role of monetary policy shocks in exchange rate behavior, conclusions have been clouded by the lack of plausible identifying assumptions. We apply a recently developed inference procedure allowing us to relax dubious identifying assumptions. This work overturns some earlier results and strengthens others: i) Contrary to earlier findings of "delayed overshooting," the peak exchange rate effect of policy shocks may come nearly immediately after the shock; ii) In every *otherwise reasonable* identification, monetary policy shocks lead to large uncovered interest rate parity (UIP) deviations; iii) Monetary policy shocks may account for a smaller portion of the variance of exchange rates than found in earlier estimates. While (i) is consistent with overshooting, (ii) implies that the overshooting cannot be driven by Dornbusch's mechanism, and (iii) gives reason to doubt whether monetary policy shocks are the main source of exchange rate volatility.

Keywords: exchange rates; overshooting; forward premium bias; monetary policy; identification.

* Division of International Finance, Board of Governors of the Federal Reserve System. Email: faustj@frb.gov and john.h.rogers@frb.gov. The authors thank for useful comments Chris Sims, Eric Leeper, Lars Svensson, Mark Watson, Tao Zha, and seminar participants at Berkeley, Duke, the Federal Reserve Board, IMF, Indiana, Michigan, Michigan State, Ohio State, Penn State, Princeton, U.C. Santa Cruz, and Virginia. Thanks to Michael Sharkey and Molly Wetzel for excellent research assistance. Some computer code for the techniques used in this paper can be found at `http://patriot.net/~faustj`. The views in this paper are solely the responsibility of the authors and should not be interpreted as reflecting the views of the Board of Governors of the Federal Reserve System or of other members of its staff.

Exchange rate changes are volatile and difficult to explain. Economists have long suspected that monetary policy shocks might play an important role in accounting for this behavior, and a great deal of theoretical and empirical work has been directed at confirming this suspicion. This paper combines recent developments in international finance and econometrics to assess what firm conclusions can be drawn about the role of monetary policy shocks in exchange rate behavior.

An obvious starting point for our study is Dornbusch's (1976) overshooting model. Having received over 800 citations,[1] this work remains at the core of international finance. Dornbusch's prediction that the exchange rate should initially overshoot its long-run level in adjusting to a monetary shock owes much of its huge appeal to two factors. First, it provides hope of explaining the empirical regularity that exchange rates in the post-Bretton Woods era are more volatile than macroeconomic fundamentals such as the money supply, output, and interest rates. Second, the overshooting conclusion follows directly from three familiar components: the liquidity effect of monetary policy shocks on nominal interest rates, uncovered interest rate parity (UIP), and long-run purchasing power parity (PPP).

While overshooting is a dominant theory in international finance, its reliance on uncovered interest rate parity means that when confronted with data, the theory will be enmeshed in a dominant empirical puzzle in international finance—the tendency of the exchange rate to change in the direction opposite to that predicted by UIP. Labelled *the forward premium anomaly*, this tendency has been extensively documented [Fama (1984), Hodrick (1987), and Engel (1996)]. Because almost all work on this anomaly has been of a reduced form variety, however, it remains an open question whether UIP holds conditionally in response to monetary policy shocks, in particular.

Motivated by these facts, we focus on three questions:

1 Does the exchange rate overshoot? More specifically, at what lag horizon does the exchange rate peak after a monetary policy shock?

[1] Social Science Citations Index

2 Is the dynamic response of the exchange rate roughly consistent with uncovered interest rate parity?

3 Can monetary policy explain a large share of exchange rate variance under *any reasonable* theory?

The first two questions are about the relation between overshooting theory and the forward premium anomaly. The third addresses whether—under overshooting or any other theory of international finance—monetary policy shocks can account for a large share of exchange rate variance.

Of the papers examining the first question, one common finding is that the exchange rate overshoots its long-run value in response to policy shocks, but that the peak occurs after one to three years as opposed to happening immediately as predicted by Dornbusch.[2] A typical delayed overshooting result is shown in Figure 1, which gives the estimated dynamic response of the U.S. dollar/U.K. pound and dollar/German mark exchange rates to a stimulative U.S. monetary policy shock in our replication of work by Eichenbaum and Evans (1995). Based on such evidence, a consensus seems to be emerging that the exchange rate shows *delayed overshooting* and theorists are attempting to rationalize this fact.[3]

Few papers directly address question two. Eichenbaum and Evans (1995) find that UIP is violated during the period when the exchange rate is rising toward its delayed peak—UIP predicts that it should be falling.[4] As for question 3, concerning the share of exchange rate variability due to monetary policy shocks, papers report estimates between a few percent to over one-half [Eichenbaum and Evans (1995), Rogers (1999), Clarida and Gali (1994)].[5]

In all of this work, a crucial and highly contentious step is identifying which

[2] The results are quite consistent for bilateral rates between the U.S. and Europe and Japan: Eichenbaum and Evans (1995) and Clarida and Gali (1994) nearly uniformly find delay; Grilli and Roubini (1996) generally find delay. Cushman and Zha (1997) find no delay for U.S.-Canada rate.

[3] e.g., Gourinchas and Tornell (1996).

[4] Cushman and Zha (1997) find that the pointwise confidence intervals for the deviation from UIP generally cover zero for the U.S.-Canada exchange rate.

[5] In Clarida-Gali and Rogers this is the effect of the relative money shock in the two countries. In Eichenbaum and Evans, this share is for the monetary policy shock in the U.S. only, as we focus on in this paper.

exchange rate movements are due to monetary policy shocks. The paucity of *highly credible* identifying assumptions forces one to use questionable assumptions and limits the number of variables that can be included in the analysis—since larger models require more assumptions. This leads to questions about whether the results are robust to including other arguably relevant variables and to changes in the dubious assumptions.

Faust (1998) develops an approach to avoiding these problems. It allows one to impose any *highly credible* restrictions and then summarize all possible ways of completing identification of the model. In this paper, we apply this technique to a standard 7-variable model and a new 14-variable model for both the US-UK and US-German bilateral exchange rates. We find the following.

1 The delayed overshooting result is quite sensitive to dubious assumptions. The data are consistent with peak exchange rate effects that are very early (say, within a month after the shock) or delayed several years.

2 Monetary policy shocks seem to generate large UIP deviations. Even when applying only minimal assumptions about what constitutes a monetary policy shock, a search for a money shock that generates small UIP deviations is fruitless. Thus, if exchange rates do peak early in response to policy shocks, this overshooting is apparently not UIP-driven, Dornbusch overshooting.

3 Consistent with earlier work, we find in the 7-variable model that the U.S. policy shock might plausibly account for anything between 8 and 56 percent of the forecast error variance of the exchange rate at the 48-month horizon. In the 14-variable model, however, this range is 2 to about 30 percent. We believe that the results for the smaller model may be due to omission of important variables.

These results are developed in 5 sections. In Sections 1 and 2, we discuss relevant international finance theory and then an example of our approach to identification. Section 3 lays out the full approach, Section 4 has results, and Section 5 presents conclusions.

1 Overshooting, UIP, and the forward premium anomaly

1.1 Overshooting

The Dornbusch overshooting hypothesis predicts that *ceteris paribus* a one-time permanent increase in the money stock will cause the exchange rate to depreciate on impact beyond its long-run value and then appreciate toward the terminal value. Overshooting is a robust prediction of models exhibiting three standard building blocks: a liquidity effect of monetary policy shocks, UIP and long-run PPP. By long-run PPP, the exchange rate must ultimately settle at a depreciated value after the money expansion. In the short-run, the liquidity effect of the money expansion implies that home interest rates to fall relative to foreign rates. UIP requires that

$$Es_{t+1} - s_t = i_t - i_t^*, \tag{1}$$

where s is the logarithm of the nominal exchange rate and i and i^* are the home and foreign one-period interest rates. If i falls relative to i^*, then the exchange rate must be expected to appreciate. *Appreciation* to a *depreciated* long-run value implies an initial jump depreciation that overshoots the long-run value.[6]

Each of the three building blocks is open to question empirically. Long-run PPP could fail, but any failure present in the data is not significant enough to play a large role in our analysis.

The liquidity effect is more problematic. There is still great uncertainty about the size and duration of the liquidity effect [Leeper and Gordon, 1992; Pagan and Robertson, 1994; Bernanke and Mihov, 1998]. Much of the complication is due to the identification problems: the data do not clearly supply us with experiments of unilateral exogenous changes in the money supply, making identification of the effects of a monetary policy shocks controversial [e.g., Rudebusch, 1998; Sims 1998].

[6] The large body of theoretical work related to Dornbusch's hypothesis and exchange rate dynamics more generally includes Alvarez, Atkeson, and Kehoe (1999), Backus, Foresi, and Telmer (1996), Chari, Kehoe, McGrattan (1998), Eaton and Turnovsky (1983), Frenkel (1982), Gourinchas and Tornell (1996), Kollmann (1999), Mussa (1982).

1.2 UIP and the forward premium anomaly

The UIP element of the Dornbusch model is most problematic empirically. Under *covered* interest parity, $i_t - i_t^* = f_t - s_t$, where f_t is the logarithm of the forward rate. A common test of UIP considers the following regression (or its equivalent under covered interest parity),

$$s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + \varepsilon_t \tag{2}$$

If (1) holds, the population values of the coefficients are $\alpha = 0$ and $\beta = 1$. In practice, for a wide range of currencies and time periods, one finds β significantly less zero, with point estimates often below -1.[7] This result is the core of the forward premium anomaly.

The deviation from UIP, call it ξ, is the forward premium:

$$\xi_t \equiv (i_t - i_t^*) - (E[s_{t+1}] - s_t) = f_t - E[s_{t+1}]. \tag{3}$$

So long as capital markets are open and the interest rates are for nominally riskless, highly liquid bonds, then there are two primary explanations for the negative β: ξ_t is a time-varying risk premium or the regression does not account for people's expectations properly. Fama (1984) demonstrated that the negative β implies a negative covariance between ξ_t and the expected change in the exchange rates. He argued that this is problematic for the risk premium explanation, as it implies the risk premium is highest when the currency is expected to appreciate.[8]

The alternative explanation is that the forward premium regression is mismeasuring expectations. In this view, the β coefficient is a biased estimate of the population value, say, due to learning or peso effects [see Engel's (1996) survey]. More recently, Phillips and Maynard (1999) have shown that estimates of β in (2) are biased downward due to the persistence of the interest rate spread.

[7] This result is most consistent for short-horizon changes and bilateral dollar exchange rates. See Fama (1984), Hodrick (1987), Canova and Marrinan (1993), and Engel (1996).

[8] Backus, Foresi and Telmer (1998) further characterize the Fama puzzle, by characterizing what the negative β implies within a standard class of asset pricing models.

These results are about unconditional UIP—the response of the exchange rate to all shocks on average. The results shed little light on whether monetary policy shocks systematically generate deviations from UIP. Indeed, of all the sources of uncertainty we often speak of, one might suppose that money shocks are least likely to generate large short-run fluctuations in risk premia. To investigate this question, we must identify the response to a monetary policy shock.

2 Conventional identification

A linear reduced form model is consistent with infinitely many causal structures of the model. The problem of identification is to choose among these causal structures. Take the reduced form dynamic model,

$$B(L)Y_t = u_t, \tag{4}$$

where Y_t is an $(n \times 1)$ vector of data, $B(L) = \sum_{i=0}^{p} B_i L^i$, $B_0 = I$, $LY_t = Y_{t-1}$, and u_t is a vector of shocks. One can premultiply both sides of (4) by any full rank matrix A_0 to arrive at a system $A_0 B(L) Y_t = A_0 u_t$, which can be written[9]

$$A(L)Y_t = w_t.$$

For any A_0, the system has reduced form, (4), and has moving average representation, $Y_t = A(L)^{-1} w_t$, or

$$Y_t \equiv C(L) w_t \tag{5}$$

The dynamic response (impulse response) of, say, the i^{th} variable to an impulse to the j^{th} shock, w_t, is given by the coefficients of $C_{ij}(L)$. The practical problem is that while each A_0 gives a system with the same reduced form, each gives rise to a different impulse response function. As Koopmans and the Cowles commission emphasized [1953], one can only choose among these different causal interpretations by bringing to bear *a priori* identifying restrictions.

[9] Since $B_0 = I$, it makes sense to name $A(L) \equiv A_0 B(L)$: the coefficient of L^0 in $A(L)$ is A_0.

The standard practice in the recent VAR literature is to identify only the dynamic response to a shock of particular interest, in our case the monetary policy shock. The causal structure of the remainder of the system is left uninterpreted. Conventional VAR identification begins with the assumption that the underlying structural shocks are orthogonal. We too will maintain this assumption throughout.[10]

After assuming orthogonality of the shocks, identifying the response to the money shock requires $N-1$ additional assumptions in an N variable system. The identification is usually completed using restrictions on contemporaneous interactions: output does not respond to a policy shock within the month, or foreign policy does not respond to home policy within the month.[11] One can typically construct plausible arguments for such restrictions [e.g., Leeper, Sims, and Zha (1996)].

2.1 Example: delayed overshooting in a 7-variable model

Take the the seven-variable model of Eichenbaum and Evans (1995). The model contains U.S. and foreign industrial production (Y and Y^*), the U.S. consumer price index (P), U.S. and foreign short-term interest rates (i and i^*), the ratio of U.S. non-borrowed reserves to total reserves ($NBRX$), and the exchange rate in dollars per foreign currency (S) (for details, see appendix A). All variables except interest rates are in logs. Data are monthly from 1974:1 to 1997:12. The reduced form is estimated with 6 lags of each variable and a constant. In this example, we focus on the US-UK case.

In a preferred identification approach, Eichenbaum and Evans identify the response to a policy shock by imposing the recursive ordering $[Y, P, Y^*, i^*, NBRX, i, S]$. The shock to $NBRX$ is interpreted as the money shock. This recursive ordering implies 6 substantive assumptions: Y, P, Y^*, and i^* do not respond to *U.S.* policy shocks within the month that they occur, and policy does not respond to shocks to

[10] In data measured at sufficiently high frequency, this assumption is not highly controversial. Even with monthly data, interactions within the month could cause problems.

[11] Some papers impose restrictions implied by, e.g., long-run monetary neutrality. See Blanchard and Quah (1989) and Faust and Leeper (1997) for a critique.

i and s within the month.

This basic identification scheme has been used in closed-economy settings by Strongin (1995) and others. The impulse responses look familiar from closed-economy applications (See Figure 2, solid lines).[12] The rise in NBRX is associated with a decline in nominal interest rates, a hump-shaped response of output that peaks around 12 to 18 months after the shock, and an initial negative response of prices that eventually turns positive. These sorts of effects are generally taken as reasonable in the literature. The exchange rate response to money peaks at about three years in these estimates, supporting the delayed overshooting conclusion.

While the identified money shock passes the duck test,[13] at least 3 of the 6 identifying restrictions are questionable. Fed policymakers are aware of data for exchange rates and domestic interest rates up to the minute when their policy decisions are taken; it is unlikely that surprising movements in those variables are ignored by policymakers. The assumption that the foreign short-term interest rate does not respond to policy within the month is also questionable. The domestic short-term rate and the exchange rate can (and do in the VAR) react contemporaneously to policy. These two variables are tied to the foreign short-term interest rate and the forward exchange rate by covered interest arbitrage. It is difficult to imagine why the foreign short-term rate would not do some of the adjusting to make covered interest parity hold.

The use of questionable restrictions is no secret, and the standard response is to present results for a few sets of identifying assumptions. Eichenbaum and Evans assess other recursive orderings of these variables and find that key results such as delayed overshooting arise systematically.

Of course, the arguments against the preferred recursive ordering hold for *any* recursive ordering. Indeed, identifications showing simultaneity among money mar-

[12] In Figures 2 and 3 we provide error bands around the OLS point estimates. These are created using the the Bayesian simulation method under the natural conjugate prior described in the RATS manual and Sims and Zha (1999). We used 1000 draws; the 68 percent coverage bands are the 16^{th} and 84^{th} percentile points from the simulation.

[13] If it walks like a duck and quacks like a duck, it might actually be a duck.

ket variables are surely *at least as plausible* as any recursive ordering. As a result, it is natural to wonder whether results like delayed overshooting are somehow special to recursive formulations or also hold for other plausible formulations.

2.2 Example continued: early peaks in the exchange rate

Lacking agreement on a set of credible identifying assumptions, one option is to search all possible identifications allowing simultaneity among $[i^*, NBRX, i, s]$. If all the credible identifications show delayed overshooting, the issue is settled. Otherwise, one must admit to uncertainty about the peak timing until sharper identifying restrictions emerge.

In the US-UK example, the method described below turns up many identifications of the 7-variable model that show no delayed overshooting. The exchange rate response to the policy shock in one such identification is shown on Fig. 2a (dashed lines). The response to the policy shock is strikingly similar to the recursive identification in all respects except that the exchange rate effect peaks in the first month after the shock. Indeed, the solid and dashed point estimates typically lie almost entirely within each other's error bands, except for the first half-year for the exchange rate and UIP deviation.[14]

The dashed line identification involves the same recursivity with respect to Y, Y^* and P as the fully recursive identification. Indeed, the only notable difference is that in the recursive system a policy shock that lowers i (by around 15 to 20 basis points) is restricted to have no impact effect on i^*, but such a shock lowers i^* (by around 5 to 15 basis points) in the dashed lines.[15] This sort of evidence suggests that

[14] Note that the fact that the pointwise error bands do not overlap for the first few months of the exchange rate response does not mean that the peak timings are statistically significantly different. Changing the peak timing on a response involves simultaneously changing several of the point responses and, thus, inference on peak timing requires consideration of both variances *and* covariances of the responses. As Kilian and Chang (1998) show, one can simulate the peak timing itself and calculate coverage intervals. We take up such simulations below.

[15] Neither Eichenbaum-Evans nor we report the response of the variables to the 6 uninterpreted shocks in the system. There are, however, no differences in the two systems with respect to the response to the first 3 uninterpreted shocks (the orthogonalized shocks in the y, p, and y^* equations). In the simultaneous system, i^*, i, $NBRX$ and S shocks each respond to the orthogonalized shocks to i^*, i and S. Since there is a presumption in favor of simultaneity among these variables, this

the delayed overshooting result may be sensitive to dubious identifying assumptions. The next section presents a method to more systematically do this sort of structural inference when identifying assumptions are questionable.

3 Setting aside dubious identifying assumptions

Faust (1998) develops an approach to inference when one is lacking sufficient restrictions to identify the items of interest. One imposes any credible assumptions, but there will generally remain a range of possible answers to questions of interest when the assumptions do not fully identify policy shock. This method allows one to do inference about this range of answers by systematically searching all identifications consistent with the restrictions.[16] For a more complete description of the approach, see Faust (1998).

3.1 Searching all reasonable identifications

Given the reduced form (4) we can always choose an A_0 that transforms the model to have orthogonal errors with unit variance (any recursive ordering will do this):

$$Y_t = C(L)w_t$$

where $Ew_tw_t' = I$. The choice of unit variance is merely a normalization. Every money shock in every possible identification (that maintains orthogonal, unit variance shocks) can be written, $\alpha' w_t$ for some α satisfying $\alpha'\alpha = 1$. Thus, we can cast our search of reasonable identifications as a search of the unit vectors α, with each α defining a shock $\alpha' w_t$.

We can limit the search by imposing some identifying restrictions we find credible. A second useful fact is that for the shock defined by α, zero restrictions and sign restrictions on the impulse response to a money shock imply linear restrictions on α. Thus, the restriction that a stimulative money shock raises money growth

difference is not of much use in distinguishing the credibility of these two formulations.

[16] This approach can be seen as a generalization of the approach in King and Watson (1992) and Bernanke and Mihov (1998).

on impact can be written $R\alpha \geq 0$, where the elements of the row vector R depend only on $C(L)$. Each added restriction adds a row to R.[17] Restrictions on linear combinations of impulse responses are also of this form, so one can restrict whether the impulse response function is rising or falling between two points.

Once we impose all highly credible assumptions, the problem is that the policy response is still not identified. We would still like to have a way to see what range of answers to our questions is possible after only imposing highly credible assumptions. For some properly structured questions, we can cast the search for this range of answers as a straightforward optimization.

Take question 2 in the introduction. One measure of UIP deviations after a policy shock is the root mean square UIP deviation over the first, say, 4 years after a policy shock. The expected UIP deviation at $t+l$ of a shock at t is given by,[18]

$$c(i,l) - c(i^*,l) - 400[c(s,l+3) - c(s,l)].$$

where $c(x,l)$ is the response of variable x at lag l to the shock defined by α. The mean square expected UIP deviation (hereafter, UIPD) comes from summing the squared deviations over some horizon.[19] In interpreting the results for UIPD, it is useful to remember that the mean square UIP deviation can be written as the squared mean deviation plus the variance over the chosen horizon. Thus, a large UIPD implies either large absolute deviations or highly variable deviations, or both.

Some simple algebra shows that the root mean square expected UIP deviation can be written, $(\alpha' M\alpha)^{1/2}$, where the elements of M are functions only of $C(L)$.

[17] The restriction that the money shock has a positive effect on the j^{th} variable at lag k requires putting the j^{th} row of C_k as a row of R.

[18] This is annualized, presumes monthly data, and three-month interest rates in annual percentage rate units.

[19] Some tricky timing and definition questions arise. We use monthly average data for exchange rates and interest rates. If the identification is correct, then the calculated UIP deviations should be interpreted as the expected path of the monthly-average UIP deviation in response to a money shock. The VAR treats the monthly average of the money market variables such as foreign interest rates and the exchange rate as available when policy is made for the month. This is not strictly correct and could contaminate the identification. This problem is no different from other VARs and is not solved, say, by methods that assume that the policymaker sees none of the monthly average. Indeed, the approach of this paper is meant specifically to shed light on this sort of problem by assessing whether results are sensitive to different assumptions in this regard.

To find whether there is any shock satisfying restrictions and leading to small UIP deviations, we can do the optimization,

$$\min_{\alpha} \alpha' M \alpha$$

subject to $\alpha'\alpha = 1$, $R_s\alpha \geq 0$, $R_z\alpha = 0$, where R_s and R_z reflect the credible sign and zero restrictions, respectively. Faust (1998) shows how to do this optimization.

If the minimum UIPD is large, then we have a robust conclusion that money shocks generate large UIP deviations: small UIP deviations are not mutually consistent with the restrictions. If the minimum UIPD is small, but the analogous maximum UIPD is large, we conclude that UIPD is not sharply identified. All three questions can be handled in this manner.[20]

3.2 Inference

Up to this point, the discussion has focussed only on point estimates, and thus we have not taken account of the fact that the reduced form parameters must be estimated. We propose two inference methods.

The first method is an extension of the conventional simulation method used to produce the error bands on the impulse responses in Figure 2.[21] For any particular value of the reduced form parameters of the VAR, we can calculate the minimum and maximum for the parameter of interest, say, θ, under the chosen restrictions. For example, θ might be the UIPD, and we calculate θ_{min} and θ_{max}. Using the standard simulation method we can obtain coverage intervals for θ_{min} and θ_{max} just as we would for impulse responses. We treat the 5^{th} percentile of θ_{min} and the 95^{th} percentile of the θ_{max} as a *robust* 90 percent coverage band.

[20] Question 3 can be handled in the same manner, interpreting the question as asking whether the policy shock accounts for a large share of the forecast error variance of the exchange rate. The forecast error variance share due to the shock defined by α can also be written as a quadratic form in α. Question 1 is somewhat different. For question 1, we can impose that the exchange rate peak in, say, the first or second period after the shock and then use the optimization algorithm to see if there is *any* shock that satisfies the money restrictions and the early peak restriction.

[21] As noted above, these are based method described in the RATS manual and studied recently in Sims and Zha (1999).

The intervals are robust in the following sense. Remember that the few restrictions we impose will not be sufficient to identify θ. The coverage interval we calculate will be robust in that it will contain the coverage interval that would be obtained under any additional restrictions (so long as the restrictions are not overidentifying).[22] Thus, any value outside the coverage interval under the minimal set of restrictions would also be outside under additional restrictions.

While robust, these coverage intervals may in practice turn out to be quite large. In computing the minimum and maximum of θ in the simulation, one cannot computationally impose everything one believes about the policy shock, tending to increase the size of the coverage interval.[23]

Procedure 2 is partial remedy for this problem. We take the maximum likelihood estimate of the reduced form parameters and find the range, $[\hat{\theta}_{min}, \hat{\theta}_{max}]$, for the parameter of interest. We call this a "nonrejection region" in that it provides a set of points that probably should not be rejected without further evidence. This approach has intuitive appeal, since it basically rests on the assumption that we should not reject any value for θ consistent with the maximum likelihood estimate. We know, for example, that valid classical confidence intervals that always contain $[\hat{\theta}_{min}, \hat{\theta}_{max}]$ can be formed.[24] Under procedure 2, we can fully inspect the impulse responses giving rise to the limiting values $\hat{\theta}_{min}$ and $\hat{\theta}_{max}$. Thus, we can avoid the problem of procedure 1 by verifying that the validity of any restrictions we believe.

Overall, in procedure 1, we impose less than we may believe and reliably learn only about parameter values that are unlikely; in procedure 2 we can impose sufficient conditions for a money shock that may be more than is strictly necessary, but

[22] That is, so long as the support for the reduced form parameters under the additional identifying assumptions is the same as that under the original restrictions. The robustness result follows directly from the fact that on each draw θ_{max} must be weakly greater than and θ_{min} weakly less than the θs that would be obtained under more restrictions.

[23] Perhaps 20 restrictions can practically be imposed in the 14 variable model. On each draw, the calculated minimum must rise and maximum must fall when additional restrictions are imposed—so long as there is a shock consistent with the restrictions. In the simulation, draws inconsistent with the restrictions are thrown out. This is discussed further in Faust (1998).

[24] Confidence intervals implied, say, by inverting a likelihood ratio test for θ would share this property. While the likelihood ratio test need not have optimality properties in the current case, optimal inference in this case is an open question.

we reliably learn only about points that probably should not be rejected without further information.

4 Empirical results

In this section we address the three questions posed in the introduction, providing evidence on the (i) timing of the peak exchange rate effect, (ii) size of UIP deviations following policy shocks, and (iii) maximum share of exchange rate variation that can be explained by money shocks. For each questions, we first present evidence about the *nonrejection* region for the parameter of interest—the minimum and maximum values for the parameter consistent with the chosen restrictions at the OLS point estimate of the VAR. We attempt to demonstrate that the impulse responses associated with these minima and maxima are reasonable responses to a money shock. When the nonrejection region is small, we move on to the simulated coverage interval results to see if we can confidently reject any values.

We present results for 4 models: 7-variable and 14-variable models for a US-UK system and US-Germany system. The 7-variable model was discussed above. The 14-variable model consists of home and foreign output (Y and Y^*), prices (P and P^*), money supplies (M and M^*), short-term nominal interest rates (i and i^*), and long-term nominal interest rates (r and r^*). We also include commodity prices (CP), and U.S. non-borrowed reserves (NBR), and total reserves (TR). All variables are in logarithms except the interest rates; the sample period and number of lags are as in the 7-variable model.

It is worth noting that the 7-variable model is large by the standards of the VAR literature, but contains no long-term interest rates and has only two foreign variables, Y^* and i^*. Clearly, these 7 variables may not contain all variables relevant to sorting out the transmission of monetary shocks at home and abroad. A significant benefit of the approach of this paper is the ability to study larger models without having to rely on increasingly questionable identifying assumptions.

4.1 When does the exchange rate peak after a monetary policy shock?

Return to the 7-variable model discussed in the example above. For the US-UK case, we have already presented a nonrejection region of 1 to 35 months for the peak exchange rate effect in the example and argued that both values are associated with reasonable shocks. For Germany we find a range of 1 to 28 months (Table 1). The US-GE impulse responses are in Figure 2b; once again, the solid line is the recursive identification and the dashed line shows an identification involving simultaneity among the money market variables. The responses of output, prices, non-borrowed reserves and interest rates in the alternative are remarkably similar to those in the recursive identification. This suggests that the alternative is reasonable—at least from the perspective of recent VAR applications.[25]

The 14-variable models give very similar nonrejection regions (Table 1). The impulse responses associated with these ranges are reasonable by conventional standards (Figure 3a and 3b). Overall, the results for the 7 and 14 variable models are quite consistent and lead us to conclude that for the peak exchange rate response, a range of one-month to roughly three-years is consistent with the data.

4.2 Monetary policy shocks and UIP

We now turn to the question of whether the economy approximately satisfies UIP conditionally in response to monetary policy shocks. We know that there are large unconditional UIP deviations in the data. For both countries and both the 7 and 14-variable models, the unconditional UIPD is about 200 basis points. For the 7-variable model, we find a nonrejection region for the conditional UIPD of about 30

[25] Although how we found these alternative identifications does not matter for the point, it may be of interest. The dashed lines on Figures 2 were generated by imposing: (1) the impact effects on P, Y, and Y^* are zero on impact; (2) the impact effect on i is negative, and on $NBRX$ and S positive; (3) the response of P at lag 80 is no larger than at lag 36; (4) (U.K. only) the response of S at lag 23 is no larger than at lag 12. Figures 2a and 2b represent the identification consistent with these restrictions that explains the largest share of the forecast error variance of output at a horizon of 48 months. It happens that this gives an early peak even when that is not imposed. The approach in the 14-variable model is very similar.

to 90 basis points in the US-UK and US-GE models (Table 2). The lower bounds are associated with the recursive ordering and the upper bounds are associated with the alternative identifications in Figure 2, which have already been argued to be reasonable. Even the lower bounds are surprisingly large, since they result from short-term interest rate declines that are brief and do not exceed 25 basis points at any time. From Figure 2, it is clear that the UIP deviations are both large at times and quite variable. The peak deviation is much larger than the changes in short-term rates and interest rate differentials.

The non-rejection regions for the UIPD in the 14-variable model are similar (Table 2); impulse responses associated with the minima are shown in the dashed lines on figure 3; maxima are associated with the solid lines.

These results suggest that we cannot reject the existence of large and variable UIP deviations, but shed no light on whether there are other reasonable money shocks that produce small deviations. The simulated coverage intervals can help answer this question.

Table 2 presents one-sided (left-tail) bounds for the simulated coverage intervals on UIPD.[26] These are the 5^{th} and 10^{th} percentiles of the minimum UIPD from the simulation.[27] On each draw, we calculate minimum UIPD—root mean square UIP deviation at horizon 48—subject to certain restrictions. We consider 2 sets of restrictions. First are money restrictions (MR) meant to be necessary for a reasonable monetary policy shock. In the 7-variable model, these are that the responses of: (1) P, Y, Y^*, $NBRX$, and S are greater than or equal to zero on impact; (2) i and i^* are less than or equal to zero on impact; and (3) P at horizon 80 is no larger than at horizon 36, Y^* is no more than one-half of that of Y on impact, and the decline in i^* is no larger than one-half of the decline in i on impact.

The second type of restrictions are shape restrictions (SR) on the path of the

[26] The results are from 1000 simulation draws.

[27] Results for the full distribution are more informative than just these two points. Such results are omitted for brevity, but are available from the authors. In every case presented in tables 2 and 3, the simulated distributions appear to be single peaked with the peak to the right of the 10^{th} percentile point.

exchange rate. Specifically, we impose that the exchange rate response falls between lags 1–2, 2–3, 3–4, 4–6, 6–12, 12–18, 18–36, and 18–80.

In the 14-variable model we use slightly different restrictions. This is because the computational burden goes up with the number of sign restrictions we use. In the 14-variable model we use for MR: (1) P, P^*, and Y^* are zero on impact; (2) Y, CP, NBR, M, M^*, S are greater than or equal to zero on impact, as is Y at lag 8; (3) i and i^* on impact, and i at horizon 4, are less than or equal to zero; (4) P at horizon 80 is no larger than at horizon 36; (5) on impact, the drop in i^* is no more than one-half of the drop in i; and (6) on impact, the rise in M^* is no more than one-half of the rise in M. Our shape restrictions on the exchange rate require that it fall between periods 1–2, 2–6, 6-12, 12–18, 18–36, 18–80.

We report results under three combinations of these restrictions, (i) MR only; (ii) MR and SR, and (iii) neither MR nor SR, under the columns labelled "none".[28] For purposes of discussion we focus on the 10^{th} percentile values associated with a 90 percent confidence bound.

For both models and countries, we find that when no restrictions are imposed one cannot rule out UIPDs of less than 10 basis points. Requiring the shock to satisfy the money restrictions raises this total to about 20 basis points. Further requiring that the shock satisfy the shape restrictions—so that the exchange rate must peak in the first month—raises the 10^{th} percentile about 10 more basis points.

Thus, money shocks that generate the sort of modest and short-lived effects on interest rates seen in the earlier figures, seem to be associated with UIP deviations that are at least 20 basis points (in the root mean square sense) over 4 years.[29] This result is largely unaffected by whether one restricts the exchange rate to peak early or not. Thus, even when the exchange rate peaks early, it is not driven by UIP as it would be under Dornbusch overshooting.

Given that there are large UIP deviations, we can ask whether these deviations

[28] By "none" we mean no impulse response restrictions. In all cases, we impose that the structural shocks are orthogonal with unit variance.

[29] We also estimated UIPDs at the 18-month horizon. The results were so similar to those at the 48-month horizon that we omit them for brevity.

have the correct correlation patterns to help generate the negative β (in (2)) in the forward premium anomaly. We can decompose this unconditional β into a weighted average of the β associated with each shock:

$$\beta = \sum_{j=1}^{n} \omega_j \beta_j$$

where β_j is the β that would emerge if all shocks but shock j were eliminated from the system and ω_j is the variance share of shock j in the total variance of the interest rate differential (See Appendix B). Thus, each structural shock either contributes to the anomaly—pushing β downward—or tends to offset it some.

In our data, the unconditional OLS estimates of βs are -1.71 (t=-3.2) for the UK and -0.54 (t=-0.53) for Germany. For all models except the UK 14-variable model, the nonrejection region for the conditional β is greater than $[-1.5, 1.5]$; for the UK 14-variable model all the conditional βs were above 2.[30] Further, the ωs for the policy shocks are always quite small. Across both countries and both 7-variable and 14-variable models, the non-rejection region is [.02,.07].[31] Overall, while money shocks generate fairly substantial UIP deviations, these shocks recieve relatively little weight in determining the unconditional β, and the sign on the contribution is not clearly identified. Thus, the evidence does not clearly support the view that monetary policy shocks are the source of the forward premium bias.

4.3 How much exchange rate variation is due to monetary policy shocks?

Table 3 provides a nonrejection range for the forecast error variance share of the exchange rate explained by the money shock at horizon 48. In the 7-variable model the nonrejection region runs from about 10 to 50 percent for both countries, consistent with earlier estimates for recursive identifications [Clarida and Gali, 1994; Eichenbaum and Evans, 1995; Rogers, 1999].

[30] Actually, we calculate the contribution of the money shock at the 48-month horizon. This is very close to the full contribution.

[31] In Figure 2a, for example, the implied βs are 0.65 for the solid line and -1.44 for the dashed line, and the corresponding values for ω are .04 and .06.

It is informative to examine the shocks that produce the variance shares at the upper end of the range. These shocks produce the largest deviations from UIP (the UIPDs are over 100 basis points), and are associated with the largest negative values of β—values like -11. Further, these shocks explain almost none of the variance of output. Thus, while one can find money shocks that account for a large part of the variance of the exchange rate, they do so by producing very extreme and, perhaps implausible, exchange rate behavior.

The simulated coverage intervals again provide additional evidence on this question. We are interested in whether or not large variance shares are consistent with the evidence, and hence focus on the 90^{th} percentile. In the 7-variable model, for the cases where only money restrictions are imposed, the *upper-bound* estimate of the exchange rate variance share is over 55 percent for the U.K. and Germany, consistent with earlier results. Once again, adding the shape restriction that the exchange rate peak early does not change the picture much. Overall, in the 7-variable model, it would be difficult to reject that policy shocks account for over half the variance of exchange rate changes, so long as one is relatively agnostic about the response of the exchange rate.

In the 14-variable model, the non-rejection region for the variance share is much smaller than in the 7-variable model: 2 to 6 percent in the case of the U.K. and 2 to 13 percent for Germany. The simulation results for the 14-variable model using only the money restrictions also produce much smaller upper-bound estimates—about 30 percent for both countries.

These results should give pause to those who believe that monetary policy shocks are the primary culprit leading to exchange rate variance. In the 7-variable model, policy shocks can account for a large share, but the exchange rate response to the shocks is very odd. In the broader 14-variable model, large shares are far less likely.

4.4 Caveats

As with all work in this area, these results should be read with caution. They are for US-UK and US-Germany only and only deal with the U.S. monetary policy shock. While the conclusions are meant to be robust to implausible identifying assumptions, we have imposed some assumptions. If the monetary shocks consistent with these assumptions are viewed as too peculiar, perhaps these identifying assumptions should be questioned. Further, while our conclusions are robust to some identification criticisms, there are ongoing debates about many possible problems with VAR work. For example, Rudebusch [1998] raises many of these arguments; Sims [1998] responds that these problems are not so serious. Continued progress on such issues as seasonal adjustment, structural stability, variable selection, and use of revised data will undoubtedly shed additional light on the questions of this paper.

5 Conclusions

Empirical work on the role of monetary policy shocks in explaining exchange rate behavior is impeded by the lack of fully credible identifying assumptions. This paper applies an inference approach to test the robustness of conclusions to the relaxation of dubious assumptions and changes in the number of variables in the model.

We find that the delayed overshooting result is sensitive to dubious assumptions. This conclusion comes from loosening the standard assumption of recursiveness in money market variables to allow plausible simultaneity. We also find that monetary policy shocks generate large expected root mean square UIP deviations. Even when imposing very little on the behavior of the money shock, we are unable to find policy shocks that generate interest rate and exchange rate responses roughly consistent with UIP. There is little evidence, however, that these large UIP deviations are the main source of the forward premium anomaly, and, indeed, monetary policy shocks may tend to offset what would otherwise be a larger anomaly.

Finally, the results suggest that monetary policy shocks may explain less exchange rate variance than previously believed. In our 7-variable model, policy shocks that account for much of the variance of the exchange rate also seem to generate very odd exchange rate behavior. In the 14-variable model, we find it highly unlikely that U.S. policy shocks account for more than one-third of exchange rate variance.

These results have important implications for what *stylized facts* theorists should be attempting to explain and they present a mixed bag for theorists hoping that relatively conventional theories will do the trick. The results allow for an early peak in the exchange rate, which might give a role for the conventional overshooting model. Unfortunately, the bulk of the variance of the exchange rate after policy shocks is due to large deviations from UIP. This is inconsistent with Dornbusch overshooting, and indeed, no conventional models we are aware of generate large variance in foreign exchange risk premia in response to policy shocks that have the modest effects on output and interest rates that we find. Perhaps models in which large *ex post* UIP deviations arise from information problems offer greater hope.

Appendix A: Data

The data were acquired through the Federal Reserve Board's database and the IMFs International Financial Statistics database. All series are expressed in natural logarithms except interest rates, which are expressed in percentage points. The series definitions and sources are listed as follows:

Source: Federal Reserve Board

Y (Y^*) = index of U.S. (foreign) industrial production - total, 1992 base;
P = U.S. CPI - all urban, all items;
NBR = non-borrowed reserves plus extended credit, seasonally adjusted, monthly average;
TR = total reserves, seasonally adjusted, monthly average;
$NBRX = NBR/TR$;
S = spot exchange rate; monthly average; US$/foreign currency;
CP = commodity prices - materials component of the U.S. producer price index.
M (M^*) = U.S. (foreign) money supply, seasonally adjusted; M1 for U.S. and Germany, M0 for the U.K.
r (r^*) = U.S. (foreign) ten-year Treasury bond rate.

Source: IMFs International Financial Statistics

i^* = foreign t-bill rate, percent per annum (line 60c);
i = U.S. t-bill rate, percent per annum (line 60c);
P^* = foreign consumer price index, (line 64).

Appendix B: Decomposing β

The population β is,

$$\begin{aligned}
\beta &= \frac{\mathrm{cov}(s_{t+3} - s_t, i_t - i_t^*)}{\mathrm{var}(i_t - i_t^*)} \\
&= \sum_{j=1}^{n} \frac{\mathrm{cov}_j(s_{t+3} - s_t, i_t - i_t^*)}{\mathrm{var}(i_t - i_t^*)} \\
&= \sum_{j=1}^{n} \frac{\mathrm{var}_j(i_t - i_t^*)}{\mathrm{var}(i_t - i_t^*)} \frac{\mathrm{cov}_j(s_{t+3} - s_t, i_t - i_t^*)}{\mathrm{var}_j(i_t - i_t^*)} \\
&= \sum_{j=1}^{n} \omega_j \beta_j,
\end{aligned}$$

where the sums are across all shocks j and cov_j and var_j are the covariances and variances of the argument variables if shock j were the only shock with positive variance. The weight ω_j is the variance share of shock j in interest differential and β_j is the β that would result if shock j were the only shock.

References

Alvarez, A., Atkeson, A., and Kehoe, P. (1999) "Volatile Exchange Rates and the Forward Premium Anomaly: A Segmented Markets View," manuscript, Federal Reserve Bank of Minneapolis.

Backus, D., Foresi, S., and Telmer, C. (1998) "Affine models of currency pricing: Accounting for the forward premium anomaly," manuscript, Stern Business School.

Beaudry, P., Devereux, M. (1995). "Money and the Real Exchange Rate with Sticky Prices and Increasing Returns," *Carnegie-Rochester Conference Series on Public Policy* 43, 55-102.

Bernanke, B. and I. Mihov (1998). "The liquidity effect and long-run neutrality," *Carnegie-Rochester Conference Series on Public Policy*, 49, 149–94.

Blanchard, O., and D. Quah (1989). "The dynamic effects of aggregate demand and supply disturbances," *American Economic Review*, 79, 655–73.

Canova, F. and J. Marrinan (1993), "Profits, risk, and uncertainty in foreign exchange markets," *Journal of Monetary Economics*, 32, 259–286.

Chari, V.V., Kehoe, P., and McGrattan, E. (1998). "Monetary Shocks and Real Exchange Rates in Sticky Price Models of International Business Cycles" manuscript, Federal Reserve Bank of Minneapolis.

Christiano, Lawrence J., Martin Eichenbaum, and Charles Evans (1997). "Monetary policy shocks: What have we learned and to what end?" manuscript, Federal Reserve Bank of Chicago.

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

Cushman, D., and T. Zha (1997) "Identifying monetary policy in a small open economy under flexible exchange rates," *Journal of Monetary Economics*, 39, 433–448.

Devereux, M., Engel, C. (1998) Fixed vs. Floating Exchange Rates: How the Pricing Decsion Affects the Optimal Choice of Exchange Rate Regimes," mimeo, University of Washington.

Dornbusch, R. (1976) "Expectations and Exchange Dynamics," *Journal of Political Economy* 84, 1161-1176.

Eaton, J., Turnovsky, S. (1983) "Covered Interest Parity, Uncovered Interest Parity, and Exchange Rate Dynamics," *Economic Journal* 93, 555-575.

Eichenbaum, M., Evans, C. (1995) "Some Empirical Evidence on the Effects of Shocks to Monetary Policy on Exchange Rates," Quarterly Journal of Economics 110, 975-1010

Engel, C. (1996) "The Forward Discount Anomaly and the Risk Premium: A Survey of Recent Evidence," *Journal of Empirical Finance.*

Fama, E. (1984) "Spot and Forward Exchange Rates," *Journal of Monetary Economics* 14, 319-338.

Faust, J. (1998) "The Robustness of Identified VAR Conclusions About Money," *Carnegie-Rochester Conference Series on Public Policy*, forthcoming.

Faust, J., and E. Leeper (1997). "When do long-run identifying restrictions give reliable results?" *Journal of Business Economics and Statistics*, 345–354.

Frenkel, J., Rodriguez, C.A. (1982) "Exchange Rate Dynamics and the Overshooting Hypothesis," *IMF Staff Papers* 29, 1-29.

Gourinchas, P.O., Tornell, A. (1996) "Exchange Rate Dynamics and Learning," NBER working paper 5530.

Grilli, V., Roubini, N. (1996) "Liquidity models in open economies: theory and empirical evidence," *European Economic Review* 40, 847-859.

Hodrick, R. (1987) *The empirical evidence on the efficiency of forward and futures foreign exchange markets*, New York: Harwood Academic Publishers.

Kilian, L. and Chang, P. (1998). "How reliable are VAR estimates of responses to monetary policy shocks?" manuscript, University of Michigan.

King, R. and M. Watson (1992). "Testing long-run neutrality," manuscript, Northwestern University.

Koopmans, T. (1953). "Identification problems in economic model construction," Chapter 2 in W.C. Hood and T.C.Koopmans, eds., *Studies in Econometric Method.* New York: Wiley.

Kollmann (1999). "Explaining International Comovements of Output and Asset Returns: The Role of Money and Nominal Rigidities," working paper, University of Bonn.

Leeper, E., C. Sims, and T. Zha (1996). "What does monetary policy do?," *Brookings Papers on Economic Activity,* 2:1996, 1–78.

Leeper, E. and D. Gordon, (1992). "In search of the liquidity effect," *Journal of Monetary Economics* 29, 341-69.

Mussa, M. (1982) "A Model of Exchange Rate Dynamics," *Journal of Political Economy* 90, 74-104.

Mussa, M. (1986) "Nominal Exchange Rate Regimes and the Behavior of Real Exchange Rates: Evidence and Implications," *Carnegie Rochester Conference Series on Public Policy* 25, 117-214.

Obstfeld, M., Rogoff, K. (1995) "Exchange Rate Dynamics Redux," *Journal of Political Economy* 103, 624-660.

Pagan, A., and J. Robertson (1994) "Resolving the liquidity effect," manuscript, University of Rochester.

Phillips, P.C.B., and A. Maynard (1999) "Rethinking an old empirical puzzle: econometric evidence on the forward discount anomaly," manuscript, Yale University.

Rudebusch, G. (1998). "Do measures of monetary policy in a VAR make sense?" *International Economic Review* 39, 907-931.

Rogers, J. (1999) "Monetary Shocks and Real Exchange Rates," *Journal of International Economics* 48, 269-288.

Sims, C. (1980). "Macroeconomics and reality," *Econometrica* 48, 1–48.

Sims, C. (1992). "Interpreting the macroeconomic time series facts: the effects of monetary policy," *European Economic Review,* 36, 975–1011.

Sims, C. (1996). "Comment on Glenn Rudebusch's 'Do measures of monetary policy in a VAR make sense?"' *International Economic Review* 39, 933–41.

Sims, C. and T. Zha (1999). "Error bands for impulse responses," *Econometrica*, 67, 1113–1157.

Strongin, S. (1995) "The Identification of Monetary Policy Disturbances: Explaining the Liduidity Puzzle," *Journal of Monetary Economics* 35, 463–498.

Table 1: Nonrejection ranges for timing of peak exchange rate effect in months

Country	nvar	Min.	Max.
US-UK	7	1	35
US-GE	7	1	28
US-UK	14	0	47
US-GE	14	0	30

Notes: Reading from the top to bottom row, the impulse response functions associated with these peaks are shown in Figures 2a, 2b, 3a, and 3b. In each case the minimum is from the dashed line; the maximum is from the solid line.

Table 2: Nonrejection range and one-sided confidence interval for UIPD (root mean square UIP deviation in percent)

				Rejection					
		Nonrejection		MR		MR+SR		none	
country	nvar	min.	max.	5^{th}	10^{th}	5^{th}	10^{th}	5^{th}	10^{th}
US-UK	7	0.37	0.82	0.19	0.21	0.30	0.34	0.08	0.09
US-GE	7	0.31	0.92	0.16	0.18	0.24	0.27	0.08	0.09
US-UK	14	0.28	0.70	0.20	0.23	0.23	0.27	0.07	0.07
US-GE	14	0.40	0.92	0.19	0.22	0.22	0.24	0.07	0.07

Notes: The impulse responses giving the nonrejection ranges are as in Table 1. In the 7-variable models (Figs. 2a and 2b), the minimum is from the solid line and the maximum is from the dashed line; in the 14-variable models (Figs. 3a and 3b), this is reversed. From top left to bottom right, the posterior odds in favor of the restrictions are 61.5, 4.1, ∞, 999, 2.9,∞,∞, 25.3, ∞,∞,199,∞.

Table 3: Nonrejection range and one-sided confidence interval for exchange rate forecast error variance share

				Rejection					
		Nonrejection		MR		MR+SR		none	
country	nvar	min.	max.	90^{th}	95^{th}	90^{th}	95^{th}	90^{th}	95^{th}
US-UK	7	0.08	0.52	0.57	0.63	0.54	0.59	0.78	0.83
US-GE	7	0.10	0.56	0.56	0.61	0.48	0.53	0.83	0.87
US-UK	14	0.02	0.06	0.28	0.32	0.24	0.27	0.74	0.80
US-GE	14	0.02	0.13	0.32	0.35	0.25	0.28	0.77	0.81

Notes: The impulse responses giving the nonrejection ranges are as in Table 2 with the exception of the 7-variable maxima. These are common values in the literature and are omitted for brevity. The posterior odds are as in Table 2.

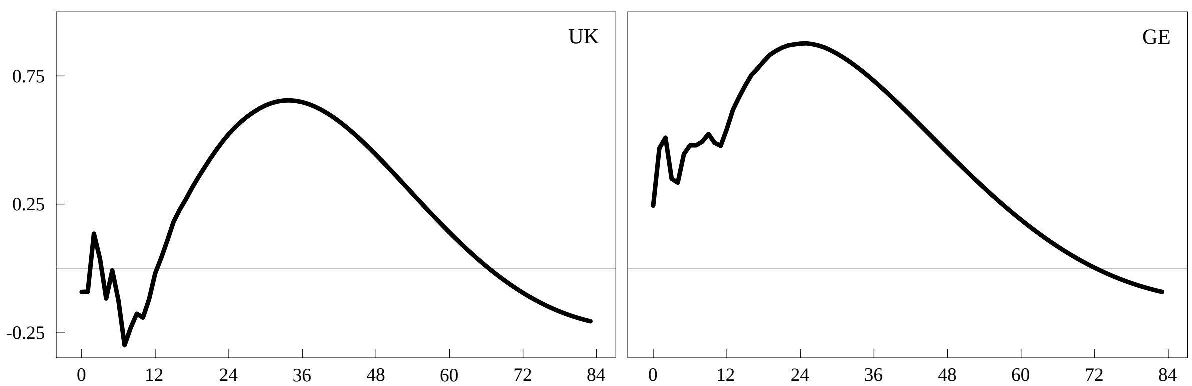

Figure 1: Responses of the \$/pound (UK) and \$/DM (GE) nominal exchange rates to a stimulative U.S. monetary policy shock in our replication of a 7-variable model of Eichenbaum and Evans (1995). The response horizon in months is given on the horizontal axis. The units on the vertical axis are in approximate percent and are the same for each panel.

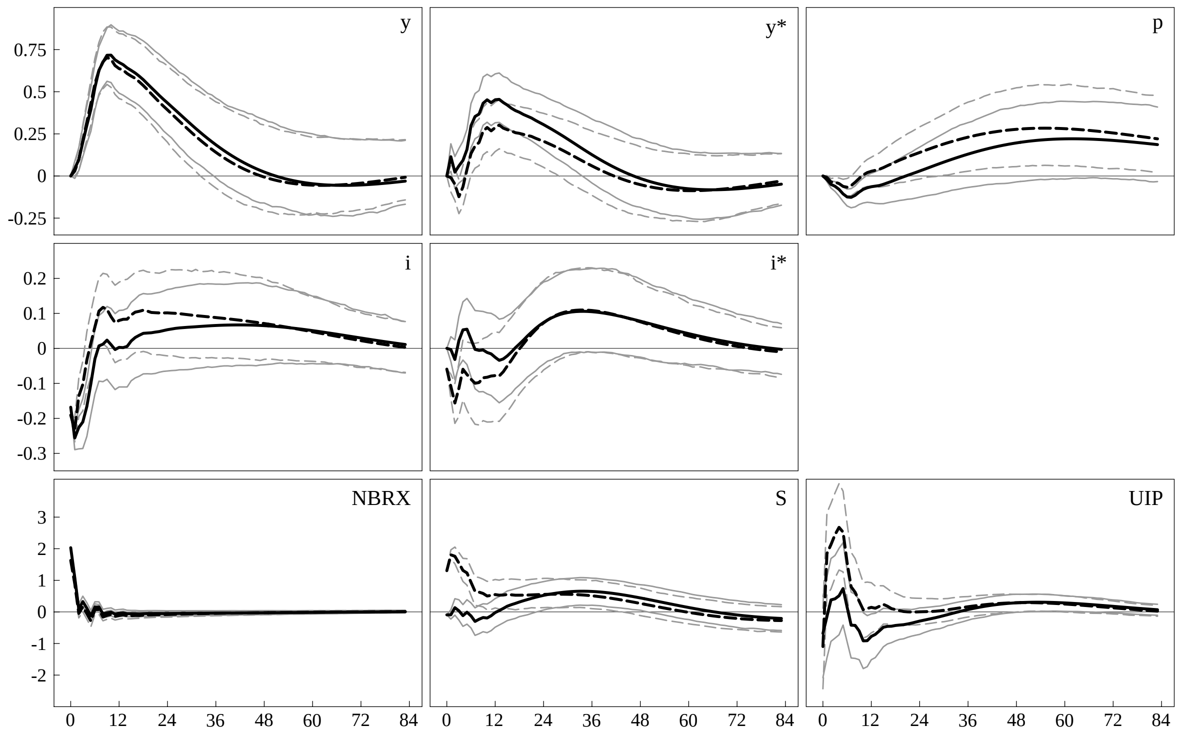

Figure 2a: Response to monetary policy shock in two identifications of the 7-variable US-UK model. The scale on the vertical axis is the same for each panel in any row. The units are approximate percent (in annual percentage rates for i, i^*, and the deviations from UIP). The horizontal axis is lag horizon in months. Error bands (as defined in the text) for the solid and dashed lines are provided by the gray solid and dashed lines, respectively.

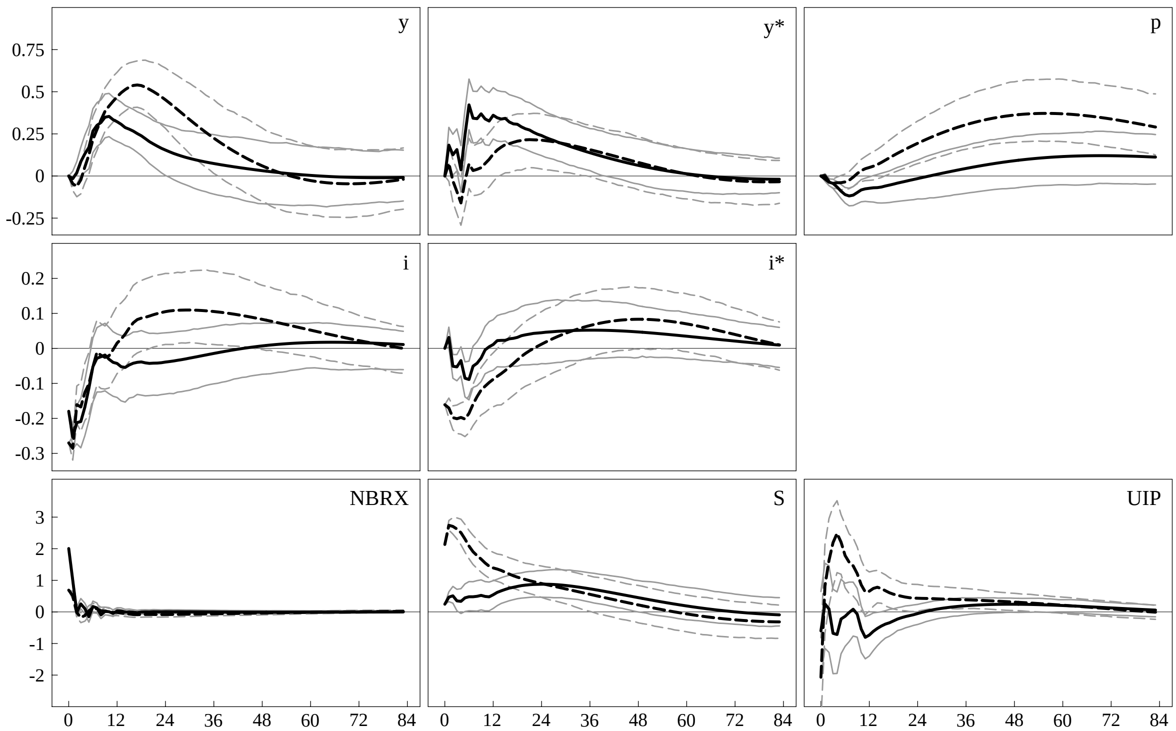

Figure 2b: Response to monetary policy shock in two identifications of the 7-variable US-GE model. See notes to Figure 2a.

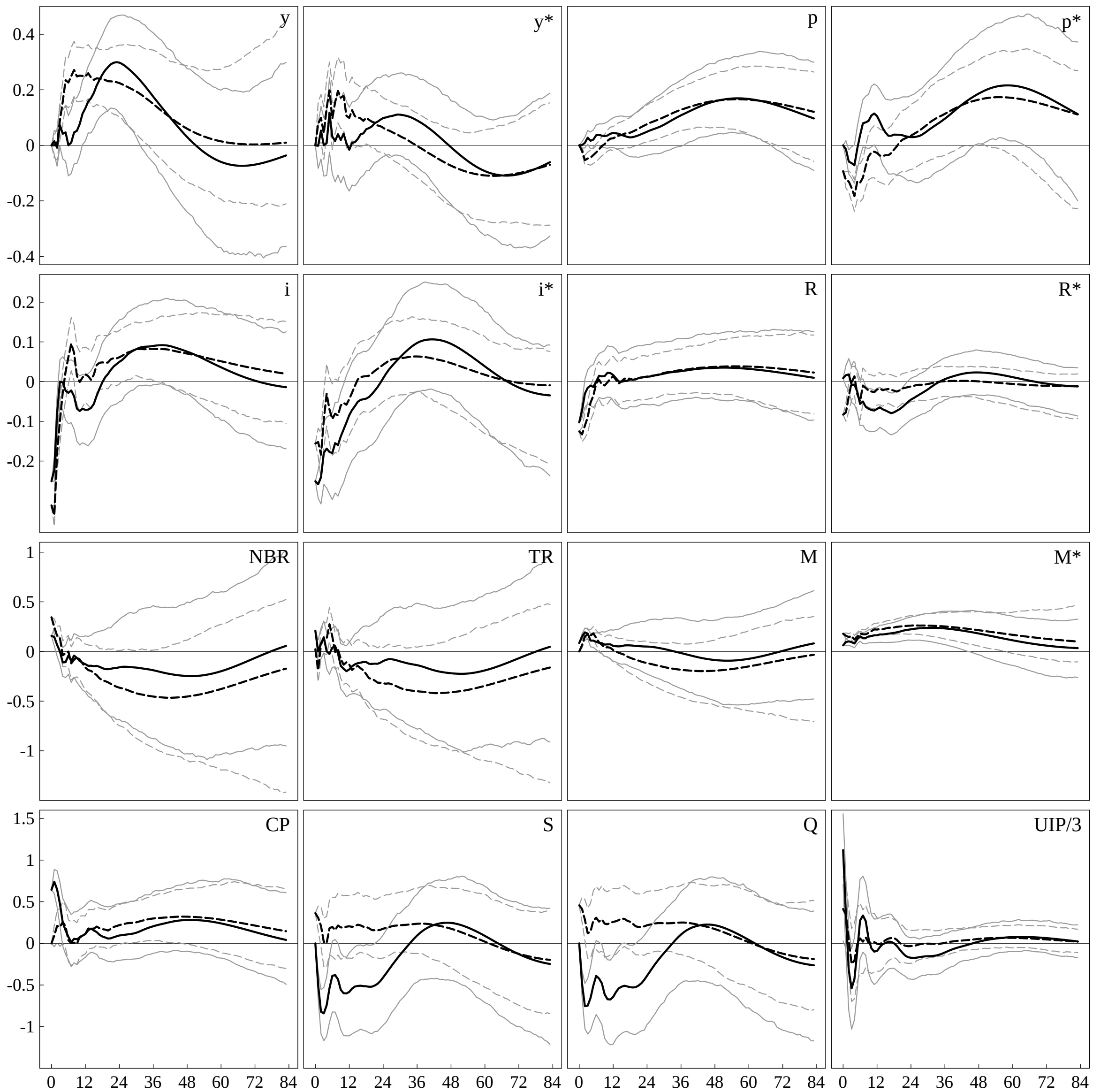

Figure 3a: Response to monetary policy shock in two identifications of the 14-variable US-UK model. In this figure only, UIP is scaled by 1/3 for readability. See the notes to Figure 2a.

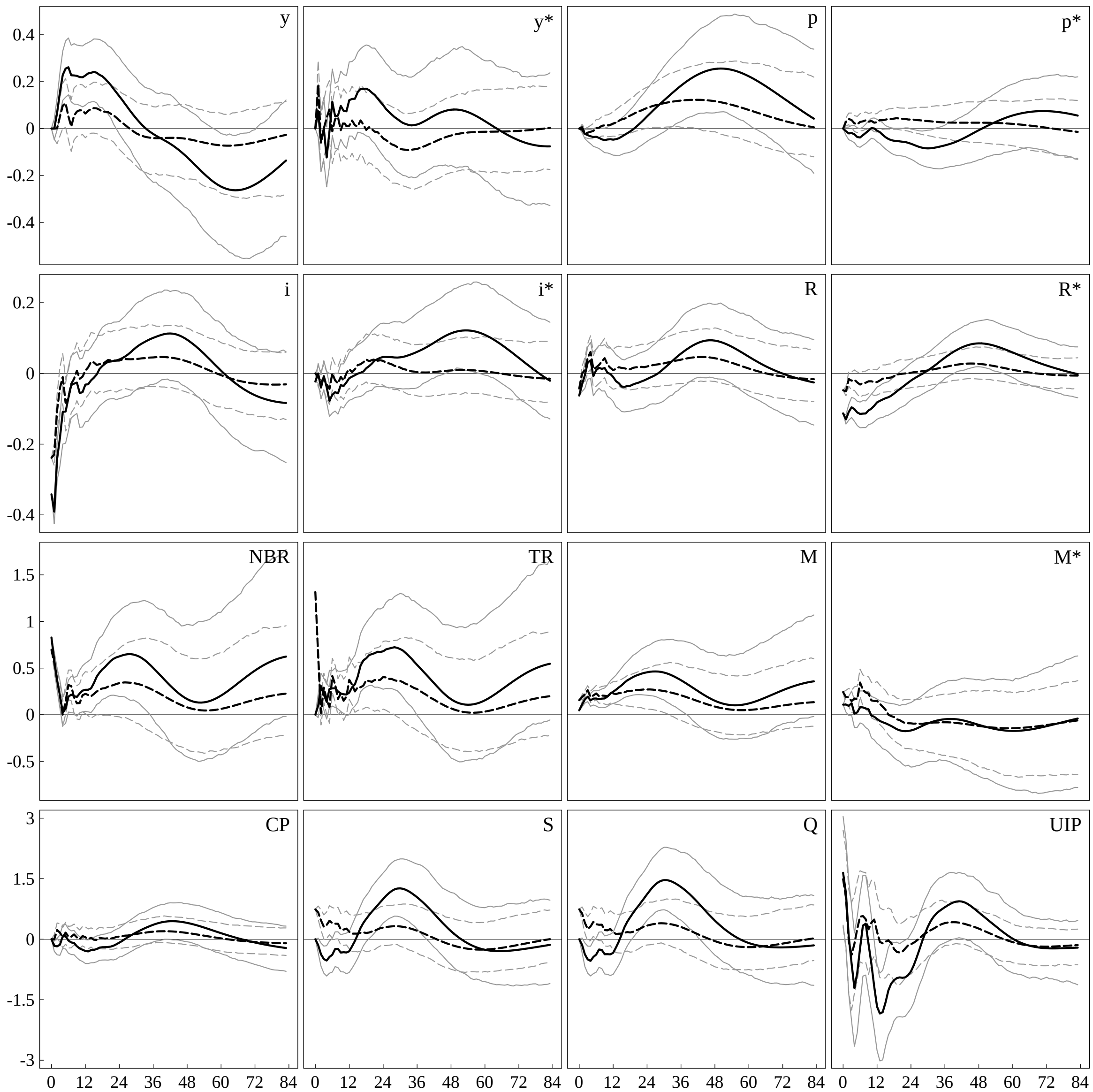

Figure 3b: Response to monetary policy shock in two identifications of the 14-variable US-GE model. See the notes to Figures 2a.