Monetary policy's role in exchange rate behavior

Jon Faust
(202) 452-2328
faustj@frb.gov
http://patriot.net/~faustj

John Rogers
(202) 452-2873
john.h.rogers@frb.gov

International Finance Division
Federal Reserve Board
Washington, DC 20551

November 1998
(revised September 30, 1999)

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 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. Indeed, there is no conventional model in which the modest policy shocks we find generate large UIP deviations.

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

We thank for useful comments, Eric Leeper, Tao Zha, and seminar participants at the Federal Reserve Board, IMF, Indiana, Michigan, Michigan State, Ohio State, Penn State, and Virginia. Thanks go to Michael Sharkey and Molly Wetzel for providing excellent research assistance.

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

An obvious starting point for this work 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 opposite direction predicted by UIP. Labelled *the forward premium anomaly*, this tendency has been extensively documented [Canova and Marrinan, 1995; Engel (1996) and Hodrick (1987)]. Because almost all work on this anomaly has been of a reduced form variety, however, it remains an open question whether UIP holds in response to monetary policy shocks, in particular.

Based on this information, 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. Of course, even if the theory is true, it need not explain much of the variance of exchange rates. The third question is meant to address whether—under overshooting or any other theory of international finance—monetary policy shocks have any hope of accounting for a large share of exchange rate variance.

Of these questions, the first and third have received the greatest previous attention. 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 dynamic response of the dollar/pound and dollar/mark exchange rates to a stimulative U.S. monetary policy shock in our replication of one of the estimated models in 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 the second question. Eichenbaum and Evans [1995] find that UIP is violated during the period when the exchange rate is rising toward its peak—UIP predicts that it should be falling.[4] As for question 3, concerning the share of exchange rate variability that can be explained by monetary policy shocks, papers report a wide range of estimates estimates between a few percent to over

[2] The results are quite consistent bi-lateral 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, in particular.

[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.

one-half.[5]

In all of this work, identifying what exchange rate movements are due to monetary policy shocks is a crucial step. Of course, identifying the response to policy shocks is an extremely contentious business. The problem is that identification requires imposing identifying restrictions; the minimal number of restrictions required grows with the number of variables in the model. The profession has, however, agreed on very few *highly credible* identifying assumptions. One is forced to choose between using a few highly credible assumptions to identify a model of 2 or 3 variables and using a larger model which requires supplementing the highly credible restrictions with more dubious ones. Most analysts reject the option of using very small models: no few variables fully characterize the policy process and transmission mechanism.[6] The larger models also generate skepticism, because the results may flow from the dubious identifying assumptions.

Faust (1998) develops an approach to inference that is robust to dubious identifying assumptions. 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 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 under far weaker restrictions than have previously been examined, policy shocks that generate only small UIP deviations are highly unlikely. Thus, if exchange rates do peak early in response to policy shocks, this overshooting is not *Dornbusch overshooting* driven by UIP.

3 Consistent with earlier work, we find in the 7-variable model that the U.S.

[5] Eichenbaum and Evans [1995], Rogers [1999], Clarida and Gali [1994]. In all these papers, this share is for the monetary policy shock in one country, generally the U.S. In this paper, we focus only on the U.S. monetary policy shock.

[6] For more complete argument and demonstration, see, e.g., Sims, 1980, Leeper, et al. 1996, Leeper and Faust, 1997.

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 4 sections. In Sections 1 and 2, we discuss the relevant international finance and our econometric approach, respectively. Section 3 contains the empirical results; in Section 4 we conclude.

1 Overshooting 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. This conclusion is a robust prediction of models exhibiting three standard building blocks: UIP, long-run PPP, and 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 will cause 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.[7]

Each of the three building blocks is open to question empirically, of course. Long-run PPP could fail, but any failure present in the data are not significant enough

[7] There is a large body of theoretical work on this topic and on exchange rate behavior more generally. 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 (1986).

to play a large role in our analysis.

The assumption of a 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]. Much of the complication is due to the identifications problems discussed above: the data do not clearly supply us with experiments of unilateral exogenous changes in the money supply.[8]

1.2 UIP

The UIP element of the Dornbusch model is most problematic empirically. In a common test of UIP, people run the regression,

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

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.[9] This result is the core of the forward premium anomaly.

Under covered interest parity, $i_t - i_t^* = f_t - s_t$, where f_t is the logarithm of the forward rate; thus, the deviation from UIP 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 is not properly measuring people's expectations. Fama (1984) demonstrated that the negative β implies a negative covariance between ξ_t and the expected change in the exchange rates. He argued

[8] While the initial Dornbusch story holds for permanent shocks; the permanence is not important. So long as a monetary expansion today does not systematically lead to a significant long-run decrease in money, the basic story goes through.

[9] This result is most consistent for bi-lateral dollar exchange rates. See Canova and Marrinan, 1995; Engel (1996) and Hodrick (1987).

that this is problematic for the risk premium explanation, as it implies the risk premium is highest when the currency is expected to appreciate.[10]

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.

Perhaps the most important lesson for the questions of this paper is that this previous work has been about unconditional UIP—the response of the exchange rate to all shocks on average. These results shed no light on whether monetary policy shocks systematically generate deviations from UIP. Of all the sources of uncertainty we often speak of, one might suppose that money shocks are least likely to generate large short-run changes in risk premia. As Eichenbaum and Evans demonstrate, one can use identified VAR techniques to study this question.

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, $B(L) = \sum_{i=0}^{p} B_i L^i$, $LY_t = Y_{t-1}$, and u_t is a vector of shocks. Since this is the reduced form, $B_0 = I$. 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[11]

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

[10] 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.

[11] 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.

This system has the same reduced form, (4), and has moving average representation,

$$\begin{aligned} Y_t &= A(L)^{-1} w_t \\ Y_t &\equiv C(L) w_t \end{aligned} \tag{5}$$

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 each A_0 gives rise to a different impulse response function, C_{ij}. As Koopmans and the Cowles commission emphasized [1953], one can only among these different causal interpretation over another by bringing to bear *a priori* identifying restrictions.

2.1 Identification in VARs

The standard practice in the 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.

The conventional VAR identification begins with the assumption that the underlying structural shocks are orthogonal. We too will maintain this assumption throughout.[12]

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.[13] One can typically construct plausible arguments for such restrictions [e.g., Leeper, Sims, and Zha, 1996].

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

[13] Some papers impose long-run monetary neutrality restrictions on the response of the real economy to policy shocks. See Blanchard and Quah [1989] and Faust and Leeper [1997] for a critique.

2.2 Example: delayed overshooting in a 7-variable model

Take the the seven-variable model in Eichenbaum and Evan's (1995) work on overshooting. The model contains U.S. and foreign industrial production (Y and Y^*), the U.S. consumer price index (P), U.S. and foreign 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). 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 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 i and s within the month.[14]

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 Fig. 1a and 2 for the US-UK results). 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. The exchange rate response to money peaks at about three years in these estimates, supporting the delayed overshooting conclusion.

2.3 The problem: Questionable identifying assumptions

In models of more than 2 or 3 variables, it has generally proven difficult to find enough uncontroversial assumptions to identify a policy shock, and this 7-variable

[14] More accurately, policy in month t does not reflect changes in month t in the domestic interest rate or exchange rate that cannot be predicted by lagged values of all the variables plus contemporaneous values of the first 5 variables.

model is no exception. While the identified money shock passes the duck test,[15] at least 3 of the 6 restrictions are quite questionable.[16]

Fed policymakers are aware of data for domestic exchange and interest rates up to the minute when their policy decisions are taken; it is not entirely plausible 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 rate and forward 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 this relation 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 variables like i, i^*, and s 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 similarly plausible formulations allowing simultaneity.

2.4 Example continued

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.

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

[16] The other assumptions can be argued as well (of course).

In the US-UK example, an *ad hoc* search 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. 2 (dashed lines). The money 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. 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 10 basis points is restricted to have no effect on i^*, but such a shock lowers i^* 3 basis points in the alternative model.[17] We conclude that the data are consistent with either early or late peak exchange rate effects. The next section presents a method to more systematically do structural inference when identifying assumptions are questionable.

3 Inference robust to questionable identifying assumptions

Faust [1999] develops a approach to when one is lacking sufficient restrictions to identify the items of interest. The method relies on three facts.

3.1 Searching the 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. The first useful fact is that every money shock in every possible identification (that maintains

[17] Neither Eichenbaum-Evans nor we report the response of the variables to the 6 uninterpreted shocks in the system. In any case, 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 difference is not of much use in distinguishing the credibility of these two formulations.

orthogonal, unit variance shocks) can be written, $\alpha' w_t$ for some α satisfying $\alpha'\alpha = 1$. Thus, we can cast our search as a search of the unit vectors α.

We can further limit the search by imposing some identifying restrictions we find credible. The second 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. The same is true of restrictions on linear combinations of impulse responses. Thus, the restriction that a stimulative money shock raises money growth on impact can be written $R\alpha \geq 0$, where the elements of R depend only on $C(L)$.[18] 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.

Of course, the method is motivated by the fact that the credible restrictions are not sufficient to identify the quantity of interest: after imposing such restrictions, there remain arbitrarily many ways to identify the model. The third important fact is that for some properly structured questions, we can cast the search of these identifications as a straightforward optimization.

Take question 2 in the introduction. One measure of UIP deviations after a policy shock is the root mean square expected UIP deviation over the first, say, 4 years after a policy shock. The expected UIP deviation at $t+l$, seen from lag t, following the shock defined by α is given by,[19]

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

where $c(x,l)$ is the response of variable x at lag l to the monetary policy shock. Some simple algebra shows that the root mean square UIP deviation (hereafter, UIPD) over any horizon can be written, $\alpha' M \alpha$, where the elements of M are functions only of $C(L)$.

To answer whether there is any *potential* money shock leading to small UIP

[18] For example, the restriction that the money shock has a positive effect on the j^{th} variable at lag k, would require putting the j^{th} row of C_k as a row of R.

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

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 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. If the minimum UIPD is small, but the analogous maximum UIPD is large, we conclude that UIPD is not sharply identified.

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.[20] 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. If so, we can use the algorithm to find the early peaking shock that implies the smallest UIP deviations or accounts for the largest share of exchange rate variance.

3.2 Inference

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

Perhaps the most common approach to inference regarding impulse responses is the Bayesian bootstrap. Sims and Zha [1999] have argued that the coverage properties of this method are, from a classical perspective, better than, say, the delta-method and similar to more sophisticated methods. To implement the procedure, one repeatedly draws from the posterior for the reduced form coefficients,[21] applies the just-identifying restrictions, and calculates the impulse response func-

[20] The forecast error variance share due to the shock defined by α can also be written as a quadratic form in α.

[21] Under the natural conjugate prior.

tion. The 5^{th} and 95^{th} percentiles of the empirical distribution is treated as a 90 percent coverage interval.

In a straightforward extension extension of this approach, we draw from the posterior for the reduced form, and for each draw find the minimum and maximum for the parameter of interest, say, θ. For example, θ might be the UIPD, and we calculate θ_{min} and θ_{max}. The 5^{th} percentile of θ_{min} and the 95^{th} percentile of the θ_{max} is a 90 percent confidence interval.

Unfortunately, these confidence intervals are likely to be excessively large because in doing the optimization on each draw, one cannot impose everything one believes about policy shock. Indeed, computationally, one can impose very few sign restrictions.[22] This causes no conceptual problem—in much inference we impose less structure than we may believe. The cost, however, is an increase in the size of the confidence interval.[23] For this reason, we believe one can with confidence reject points outside these confidence intervals, but many points inside might be rejected by considering a more complete set of restrictions.

Procedure 2 is partial remedy for this problem. We take the maximum likelihood estimate[24] of the reduced form parameters and find the range, $[\theta_{min}, \theta_{max}]$ for the parameter of interest. We assert that this range is contained in a reasonable confidence interval 95 percent for that parameter; we remain silent about the outer limits of the confidence interval. Thus, this procedure only teaches us about some points that probably should not be rejected. 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 that valid confidence intervals satisfying the assumption can be formed; confidence intervals implied, say,

[22] Perhaps 18 restrictions can practically be imposed in the 14 variable model.

[23] On each draw, the calculated minimum must rise and maximum must fall when additional restrictions are imposed. This is so long as the added restrictions are not locally over-identifying—that is, so long as there is a no shock consistent with the restrictions. Otherwise the minimum and maximum are undefined. Thus, the confidence limits will be conservative whenever one does not believe, but fail to impose, locally over-identifying restrictions. The method will often be conservative when this condition is not met. See Faust [1998] for more discussion of this issue.

[24] The maximum likelihood estimate conditional on the initial conditions.

by inverting a likelihood ratio test for θ would share this property.

Importantly, under procedure 2, we can fully inspect the impulse responses giving rise to the limiting values θ_{min} and θ_{max}. We can impose any restrictions we like and avoid the problem of procedure 1.[25] Indeed, in procedure 2 we are free to impose sufficient conditions for a reasonable shock, not worrying about whether the conditions are necessary.[26] Relaxing sufficient conditions can only expand $[\theta_{min}, \theta_{max}]$, expanding the nonrejection set.

Overall, in procedure 1, we impose less than we may believe and reliably learn only about points that should probably be rejected; in procedure 2 we can impose more than is strictly necessary, but we reliably learn only about points that probably should not be rejected.

4 Empirical results

In this section we address the three questions posed in the introduction, providing evidence on (i) timing of the peak exchange rate effect, (ii) size of UIP deviations following policy shocks, and (iii) the maximum share of exchange rate variation that can be explained by money shocks. We first present evidence about the *nonrejection* region $[\theta_{min}, \theta_{max}]$ associated with the maximum likelihood estimate under various sets of restrictions. We argue that the impulse responses giving rise to the limiting values are reasonable. When the nonrejection region is small, we move on to the bootstrap results to see if minimal restrictions will allow us to confidently reject any values. Our results are for 7 and 14 variables models of US-UK and US-Germany.[27]

As noted above, lack of sufficient identifying restrictions often forces analysts to use small models that leave out seemingly relevant variables. The 7-variable model

[25] One could, in principle, do this for every draw in procedure 1, but this would be quite slow and difficult to document and reproduce.

[26] Thus, we could impose that the policy shock's effect on output in the first period is larger than the effect on prices. While we can imagine policy shocks that do not do this, shocks satisfying this property are surely reasonable.

[27] We did some work with the U.S. dollar against the Japanese yen, Italian lira, French Franc, and Canadian dollar, and generally found similar results.

contains no long-term interest rate, which is arguably important in the transmission mechanism. The model comprises only two foreign variables, Y^* and i^*, and, for example, surely does not contain all the variables determining the foreign policy response to US policy. The model also excludes commodity prices, which Sims (1992) argues belongs in monetary VARs because it proxies for supply shock information.

The method of this paper allows us to analyze larger models. We present results for a fourteen-variable model consisting 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.

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 UK case, we have already presented a nonrejection region of 1 to 35 months 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-Germany 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.[28]

[28] 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

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 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 policy shocks. We find substantial UIP deviations and then take up the question of whether these deviations contribute to or tend to offset the negative β that characterizes the unconditional forward premium anomaly.

For the 7-variable model, we find a nonrejection region of about 30 to 90 basis points in the US-UK and US-Germany 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 at the low end of the range, these deviations are very large since they result from short-term interest rate declines that are brief and do not exceed 25 basis points at any time. If these deviations are risk premia, they are larger premia than the changes in short-term rates that precipitate them. 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 solid lines on figure 3; maxima are associated with the dashed lines.

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

Table 2 presents one-sided (left-tail) confidence bounds on UIPD. We show the 5^{th} and 10^{th} percentiles of the minimum UIPD from the bootstrap.[29] On each draw, we calculate minimum UIPD—root mean square UIP deviation at horizon

approach in the 14-variable model is very similar.

[29] We did 2000 and 1000 draws for the 7-variable and 14-variable models, respectively.

48—subject to 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 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) neither MR nor SR, under the columns labelled "none"; (ii) MR only; or (iii) MR and SR. 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. For both countries and models, 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—requiring the exchange rate to peak in the first month— raises the 10^{th} percentile about 10 more basis points.

Recall that the UIPD statistics are root mean square deviations over the four-year horizon. 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. 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 have the correct correlation patterns to help generate the negative β in the forward premium anomaly. We can decompose the unconditional β into the contribution due to the response to each shock in the economy. Each structural shock either contributes to the anomaly—pushing β downward—or tends to offset it some. From the impulse response to the money shock, we can calculate the contribution of the identified money shocks to β.[30] One can view the conditional β as the value β would take if the economy were only subject to money shocks, *ceteris paribus*. For both countries in the 7-variable models, the nonrejection region for β spans -1.5 to 1.5.[31] Thus, while all money shocks generate fairly substantial UIP deviations, it is an open question whether they contribute to or offset the negative β phenomenon.

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].

It is informative to examine the shocks that produce the variance shares at the

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

[31] For example, the solid line on Figure 2a implies a β of 0.65 and the dashed line implies -1.44.

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 no less than -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. It is not clear that *explaining* exchange rates with such a shock is very useful: while it pinpoints the source of the exchange rate variation as a policy shock, the exchange rate variation that flows from the shock is rather bizarre.

The bootstrap results provide additional evidence on this question. We are interested in ruling out large values in this case, and 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 all the 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 variance, so long as one is agnostic about the response of the exchange rate.

In the 14-variable model, the non-rejection region for the maximum share of exchange rate variance explained by money shocks 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 estimate – about 30 percent for both countries. This raises a serious question about whether the shares of over one-half shown for the 7-variable model are due to the omission of important variables, as discussed above.

We think that these results should give serious pause to those who hope 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 odd. In the 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 on which earlier work rested, there are ongoing debates about possible problems with VAR work. Rudebusch [1998] raises many of these arguments; Sims [1996], on the other hand, argues that these problems are not so serious. Continued progress on such issues as seasonal adjustment, structural stability, and use of revised data will undoubtedly shed additional light on the questions of this paper.

5 Conclusions

A great deal of theoretical and empirical work in international finance has addressed the role of monetary policy shocks in explaining exchange rate behavior. Empirical work on this subject is impeded by the lack of fully credible identifying assumptions. Even in models of only 7 variables, identifying the policy shock requires supplementing our few solid identifying assumptions with more dubious ones. Of course, 7 variables may be too few to properly sort out the interactions between policy effects in two countries.

This paper applies an inference approach that allows one to proceed based only on solid identifying assumptions. This allows us to check the robustness of conclusions in a standard 7-variable model and to present results for a new 14-variable model. The results sharpen some conclusions in the literature, and overturn some earlier conclusions.

We find that the delayed overshooting conclusion is sensitive to dubious assumptions. This conclusion comes from loosening the standard assumption of recursiveness in money market variables to allow some simultaneity. Such simultaneity is probably at least as plausible as any recursiveness assumption. We also find that monetary policy shocks generate deviations from UIP on average. 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. The UIP deviations tend to be larger than the largest absolute change in interest rates after the policy shock. The UIP deviations following a money shock may contribute to the negative regression β that characterizes the unconditional forward premium anomaly, but there are also reasonable identifications in which the response to money shocks tends to offset what would otherwise be a larger anomaly.

Finally, the results provide reason to question earlier estimates that the policy shocks might explain most of the variance of the exchange rate. 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 policy shocks account for more than 1/3 of exchange rate variance.

These results have important implications for what *stylized facts* theorists should be attempting to explain; 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, it will probably prove difficult to find any model in which rather modest monetary policy shocks generate large variance in foreign exchange risk premia. Many currently popular multi-country general equilibrium models are solved by approximation under certainty equivalence which assumes away an important role for risk premia. These models have essentially no hope of generating the results found in the data. Perhaps models in which large *ex post* UIP deviations arise from information problems offer greater hope.

Data Appendix

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).

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. [1996]. "Comment on 'What does monetary policy do'," *Brookings Papers on Economic Activity,* 2, 69-73.

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.

Cooley, T. and S. LeRoy (1985) "Atheoretical macroeconomics: A critique," *Journal of Monetary Economics*, 16, 283–308.

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.

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.

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*, forthcoming.

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?"' manuscript, Yale University.

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

Sim, C. and T. Zha (1996a). "Does monetary policy generate recessions?" manuscript, Yale University.

Sims, C. and T. Zha (1996b). "Bayesian Methods for Dynamic Multivariate Models," manuscript, Yale University.

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

Todd, Richard M. (1991). "Vector autoregression evidence on monetarism: another look at the robustness debate," *Quarterly Review of the Federal Reserve Bank of Minneapolis*, Spring.

Uhlig, H. (1997). "What are the effects of monetary policy? Results from an agnostic identification procedure," manuscript, Tilburg University.

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 solid line; the maximum is from the dashed 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 Figure 1. In each case the minimum is from the solid line; the maximum is from the dashed line. From top left to bottom right, the posterior odds in favor of the restrictions are 61.5, 4.1, ∞, 999, 2.9,∞, 25.3, ∞,infty,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 Figure 1 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.

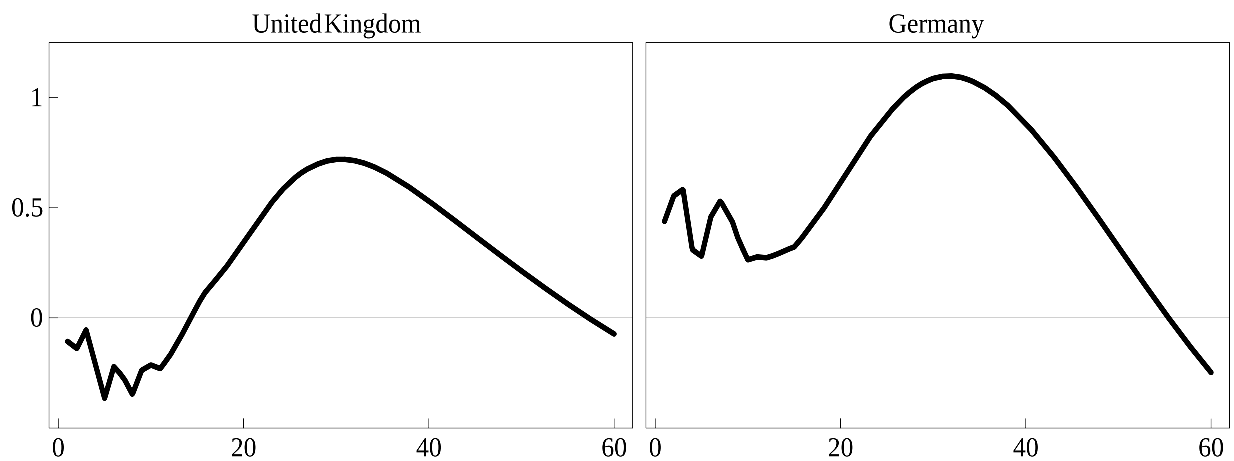

Figure 1: Responses of the \$/pound and \$/DM nominal exchange rates to a positive shock to U.S. non-borrowed reserves in the 7-variable (recursive) model of Eichenbaum and Evans (1995). The response horizon is given on the horizontal axis. The scales on the vertical axis are in percent, and are the same for each panel.

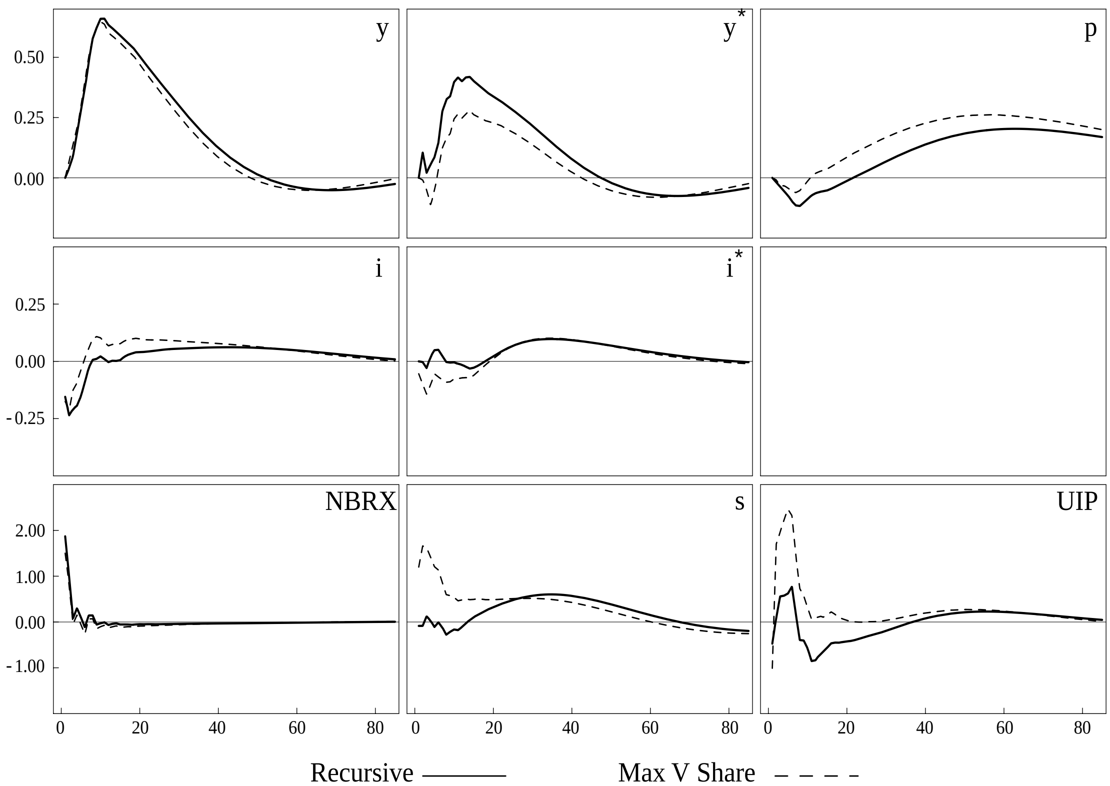

Figure 2a: U.K. results. Response to the policy shock in the 7-variable model for the recursive and Faust-Rogers identifications. In the latter, the shock is the one that maximizes the forecast error variance share of output, subject to the restrictions on the impulse responses listed in the text. The scale on the vertical axis is the same for each panel in the particular row. The units are approximate percent (in annual percentage rates for i, i*, and the deviations from UIP).

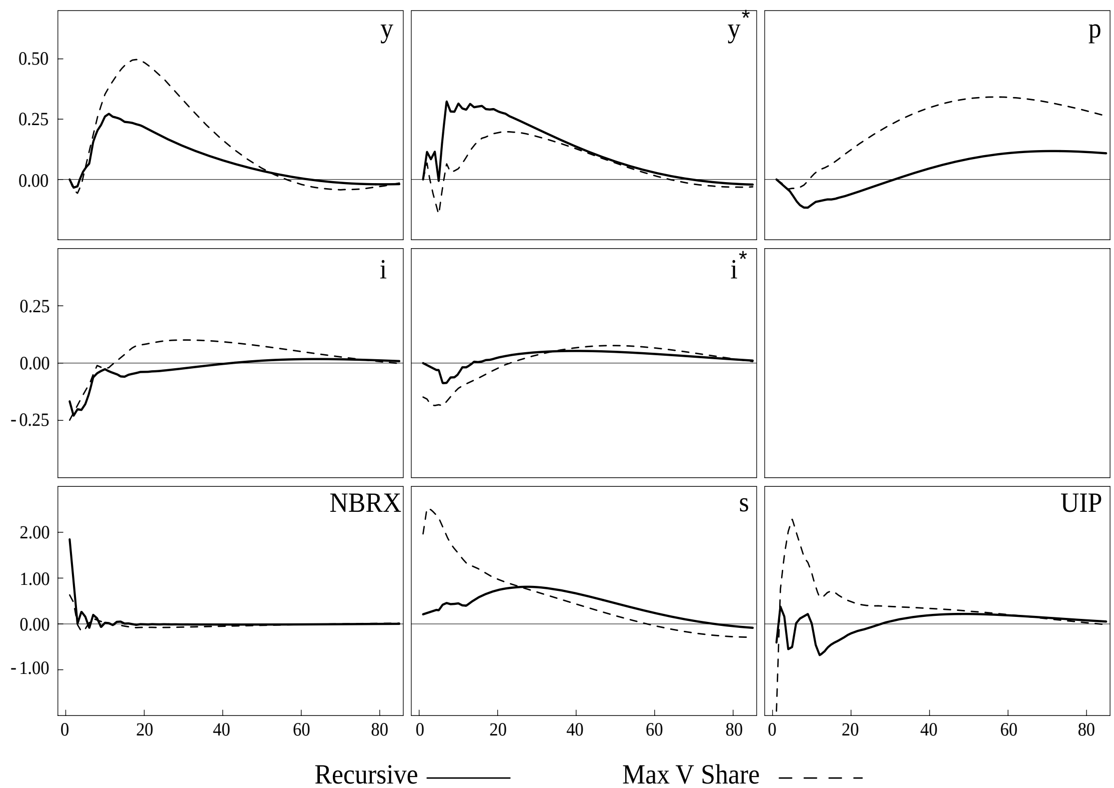

Figure 2b: German results. See notes to Figure 2a.

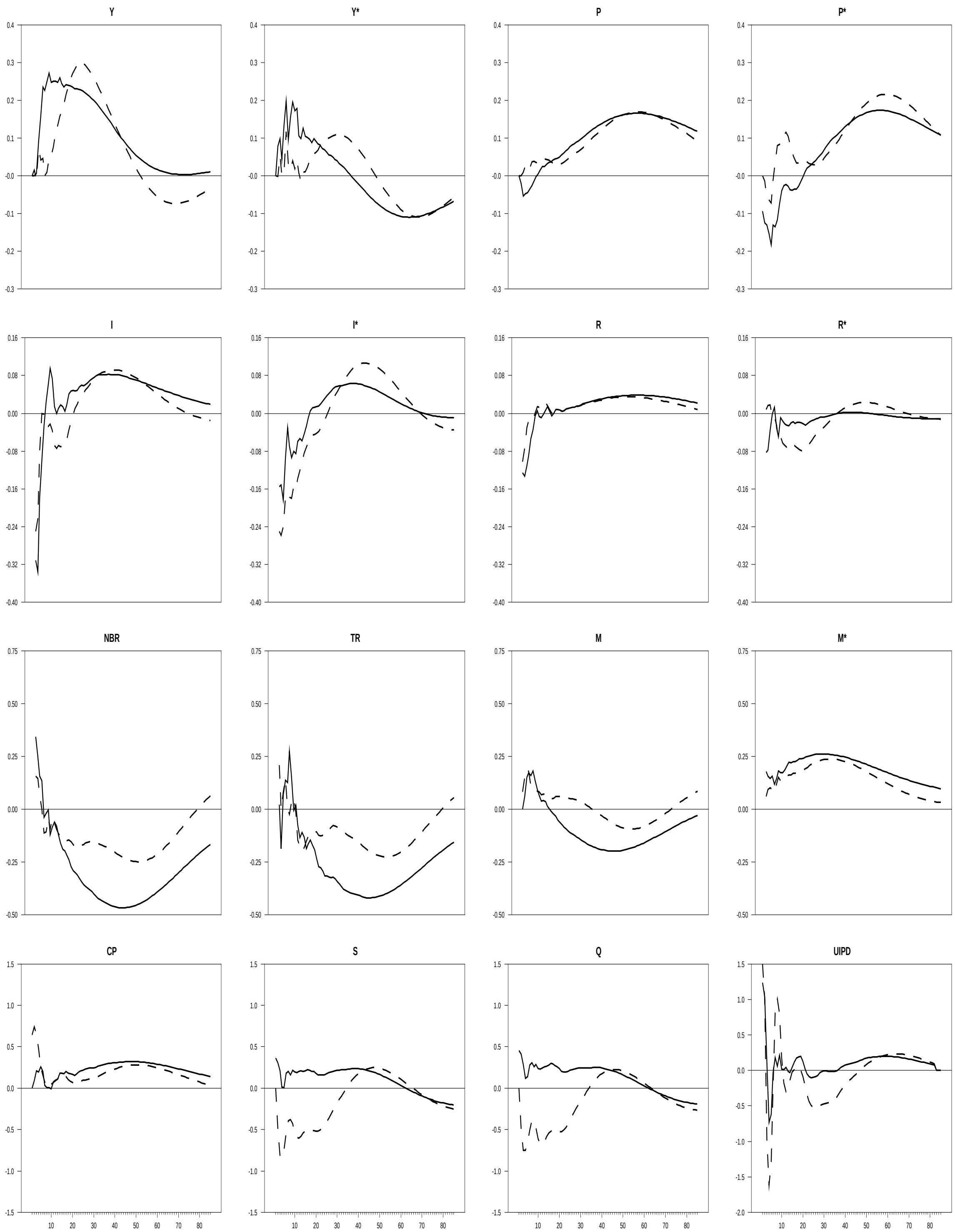

Figure 3a: 14-variable model results, U.K. Response to the policy shock in the 14-variable model under Faust-Rogers identification. The shock depicted by the solid (dashed) lines is the one that maximizes the forecast error variance share of output (exchange rate), subject to the restrictions on the impulse responses listed in the text. The units are approximate percent (in annual percentage rates for i, i*, and the deviations from UIP).

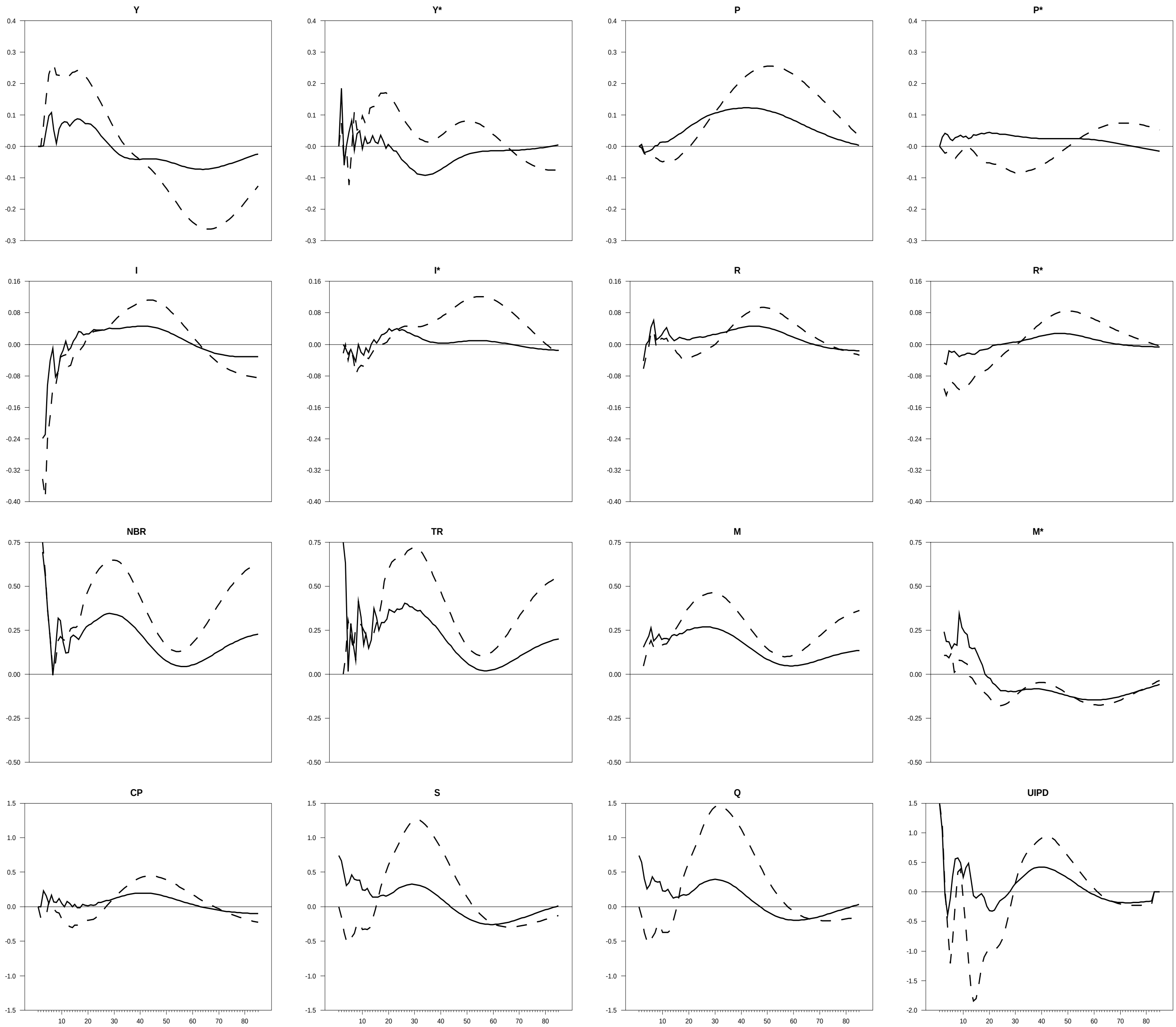

Figure 3b: 14-variable model results, Germany. Response to the policy shock in the 14-variable model under Faust-Rogers identification. The shock depicted by the solid (dashed) lines is the one that minimizes the sum of squared deviations from UIP (maximizes the forecast error variance share of the exchange rate), subject to the restrictions on the impulse responses listed in the text. The units are approximate percent (in annual percentage rates for i, i*, and the deviations from UIP).