SOME EMPIRICAL EVIDENCE ON THE EFFECTS OF SHOCKS TO MONETARY POLICY ON EXCHANGE RATES*

MARTIN EICHENBAUM AND CHARLES L. EVANS

This paper investigates the effects of shocks to U. S. monetary policy on exchange rates. We consider three measures of these shocks: orthogonalized shocks to the federal funds rate, orthogonalized shocks to the ratio of nonborrowed to total reserves and changes in the Romer and Romer index of monetary policy. In sharp contrast to the literature, we find substantial evidence of a link between monetary policy and exchange rates. Specifically, according to our results a contractionary shock to U. S. monetary policy leads to (i) persistent, significant appreciations in U. S. nominal and real exchange rates and (ii) significant, persistent deviations from uncovered interest rate parity in favor of U. S. interest rates.

I. INTRODUCTION

This paper investigates the effects of shocks to U. S. monetary policy on exchange rates. In sharp contrast to the literature we find substantial evidence of a link between monetary policy and exchange rates. Specifically, according to our results a contractionary shock to U. S. monetary policy leads to (i) persistent, significant appreciations in U. S. nominal and real exchange rates and (ii) significant, persistent deviations from uncovered interest rate parity in favor of U. S. investments.

Our analysis builds on the literature aimed at explaining the fundamental sources of exchange rate determination and the link between alternative exchange rate regimes and international business cycles.[1] In contrast to much of this literature, we investigate how exchange rates respond to a specific impulse, namely a shock to monetary policy. We focus on conditional correlations because of the difficulty of interpreting unconditional correlations in environments where agents are subject to multiple sources of uncertainty. Consider, for example, the widely noted fact that real exchange rates have been substantially more variable after the collapse of the Bretton Woods agreements. Mussa [1986] argues that this reflects the importance of sluggish price adjustment and the increased

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1. For recent surveys of empirical research on nominal exchange rates, see Engel [1995], Frankel and Rose [1994], and Lewis [1994]. See Backus and Kehoe [1992] and the references therein for work on the links between business cycles and exchange rates.

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volatility of monetary policy. In contrast, Stockman [1988] argues that it could reflect the greater variance of real shocks in the floating exchange rate era. Evidently, the mere observation that exchange rates were more variable after the collapse of Bretton Woods cannot be viewed as decisive.

To deal with the identification problem inherent in interpreting exchange rate movements, we attempt to isolate measures of exogenous shocks to monetary policy. Our strategy is closely related to recent work on the interest rate effects of monetary policy shocks in closed economy settings.[2] To assess the robustness of our results, we consider three measures of monetary policy shocks that have been proposed in this literature: orthogonalized innovations to the federal funds rate [Bernanke and Blinder 1992], the ratio of nonborrowed to total reserves [Strongin 1992], and the index proposed by Romer and Romer [1989]. As it turns out, our qualitative results are robust across the three measures.

Our main results can be summarized as follows. First, we find that contractionary shocks to U. S. monetary policy are followed by sharp, persistent increases in U. S. interest rates, and sharp, persistent decreases in the spread between foreign and U. S. interest rates. Second, we find that the same shocks lead to sharp, persistent appreciations in U. S. nominal and real exchange rates. Taken together, these findings cast doubt on international Real Business Cycle (RBC) models in which money is introduced simply by adding cash-in-advance constraints or a transactions role for money. This is because a generic implication of these models is that negative contractionary shocks to the money supply cause domestic interest rates to fall and lead to a rise in the spread between foreign and domestic interest rates. Our findings provide support for limited participation, monetized RBC models that allow for liquidity effects (see Grilli and Roubini [1992, 1993] and Schlagenhauf and Wrace [1992a, 1992b]. They are also consistent with models that stress the importance of nominal rigidities (see, for example, Dornbusch [1976] or Frankel [1979]).

Third, we find that the maximal effect of a contractionary monetary policy shock on U. S. exchange rates is not contemporaneous; instead the dollar continues to appreciate for a substantial period of time. This finding is inconsistent with simple rational expectations overshooting models of the sort considered by Dorn-

2. For a review of this literature see Christiano and Eichenbaum [1992a] or Cochrane [1994].

busch [1976]. In conjunction with our finding that contractionary policy shocks lead to a fall in the spread between foreign and U. S. interest rates, the persistent appreciation of the dollar is also inconsistent with the hypothesis of uncovered interest rate parity. Under that hypothesis, the larger interest rate differential induced by a contractionary U. S. monetary policy shock should be offset by expected future depreciations in the dollar. Our empirical results indicate that the opposite is true: the larger return is actually magnified by expected future appreciations in the dollar. So a shock to U. S. monetary policy is associated with persistent expected "excess returns."

The finding that the U. S. dollar appreciates gradually after a contractionary monetary policy is related to the literature on the forward premium bias. That literature finds that future changes in the exchange rate tend to be *negatively* related to the forward premium (see, for example, Hodrick [1987], Engel [1995], Lewis [1994], and Frankel and Rose [1994]). This pattern is often referred to as the forward premium puzzle. What is new about our result is that we find a monetary-policy-induced forward premium puzzle. Specifically, a contractionary U. S. monetary policy shock leads to a rise in the U. S. interest rate relative to foreign interest rates. This rise is associated with a persistent appreciation of the dollar. Consequently, high interest rate differentials will be associated with an appreciating currency, thus leading to a conditional negative forward premium bias.

Finally, our results shed some light on the relationship between Romer and Romer's [1989] index of monetary policy contractions and alternative measures of shocks to monetary policy. Specifically, we find that a unit increase in the Romer and Romer index is associated with a sharp rise in the federal funds rate and a sharp decrease in the ratio of nonborrowed to total reserves. The peak response of these variables occurs with a six-month delay, and is large relative to those associated with our other policy shock measures. In effect, Romer and Romer episodes correspond to *large* monetary contractions. Nevertheless, the qualitative response of exchange rates and interest rates is very similar across the three measures of policy. The main difference is that the precision of our estimates falls sharply when we move to the Romer and Romer index. Presumably, this reflects the small number of Romer and Romer episodes.

The remainder of the paper is organized as follows. Section II discusses the measures of shocks to monetary policy that are used in our analyses. Section III presents our empirical results. Section IV relates our results to the literature on the forward premium bias. Concluding remarks are contained in Section V.

II. Measuring Shocks to Monetary Policy

To measure the effects of shocks to monetary policy, we must take a stand on an empirical measure of those shocks. Here we consider three measures: orthogonalized components of the innovation to the ratio of nonborrowed to total reserves, orthogonalized components of the innovation to the federal funds rate, and the Romer and Romer [1989] index of monetary policy contractions.

The basic strategy underlying the first two measures is to identify monetary policy shocks with the disturbance term in a regression equation of the form,

$$V_t = \zeta(\Omega_t) + \epsilon_{Vt}. \tag{1}$$

Here V_t is the time t setting of the monetary authority's policy instrument, ζ is a linear function, Ω_t is the information set available to the monetary authority when V_t is set, and ϵ_{Vt} is a serially uncorrelated shock that is orthogonal to the elements of Ω_t. To rationalize interpreting ϵ_{Vt} as an exogenous policy shock, (1) must be viewed as the monetary authority's decision rule for setting V_t. In addition, the orthogonality conditions on ϵ_{Vt} correspond to the assumption that date t policy shocks do not affect the elements of Ω_t. The first two measures of policy shocks that we use correspond to different specifications of V_t and Ω_t. Conditional on this specification, the dynamic response of a variable to a monetary policy shock corresponds to the regression coefficients of the variable on current and lagged values of the residuals to equation (1).

Feedback rule (1) can be thought of as emerging from an infinite horizon optimal control problem in which the monetary authority maximizes the expected value of a criterion function subject to the constraints of technology and private agents' decision rules.[3] Under this interpretation, the shock ϵ_{Vt} might reflect exogenous shocks to the preferences of the monetary authority, perhaps due to shifts in the relative weight given to unemployment and inflation. More generally, ϵ_{Vt} could reflect a variety of random

3. The optimal decision rule will be linear if the monetary authority has a quadratic criterion function and linear constraints. Alternatively, (1) can be viewed as a linear approximation to the true decision rule.

factors that affect policy decisions. These include the personalities and views of the members of the Federal Open Market Committee (FOMC), political factors, as well as technical factors like measurement error in the data available to the FOMC when they decide on policy actions.

Less favorable to our procedure, the shock ϵ_{Vt} could reflect error in the way we have specified the monetary authority's decision rule. For example, the Fed's decision rule could have changed during the sample period that we consider. One way to deal with this problem is to investigate the robustness of inference to splits in the sample. Evans [1994] and Lewis [1993] look at weekly exchange rate data for two subsamples of the 1974–1990 period. Using a VAR-based identification scheme slightly different from ours, they obtain results similar to ours. A different possibility is that the Fed's decision rule is nonlinear. In the extreme case V_t would be an exact nonlinear function of Ω_t. Under these circumstances, the estimated time series ϵ_{Vt} would entirely reflect the error involved in approximating a nonlinear function with a linear function. A different form of nonlinearity might arise if the actual decision rule of the Fed involves moving V_t by discrete amounts. For example, each period the Fed chooses between not changing the federal funds rate at all or moving it by 25, 50, or 75 basis points. Since decision rule (1) assumes that V_t has continuous support, the estimated time series on ϵ_{Vt} would in part reflect specification error. In general, these types of specification errors imply that our procedure for isolating shocks to monetary policy is not valid. But absent taking a stand on the precise nonlinearities in the Fed's decision rule, it is hard to say whether these sources of error would substantively affect inference.

Conditional on these caveats, the procedure that we use to estimate the effects of exogenous shocks to policy is asymptotically equivalent to computing the impulse response function of a variable to a particular shock in an appropriately identified Vector Autoregression (VAR). Denote the set of variables in the VAR by Z_t. Assume that Ω_t includes the lagged values of Z_t as well as the time t values of a subset of the variables in Z_t, which we denote by X_t. The identifying assumptions in (1) correspond to a Wold ordering in which X_t is (causally) prior to V_t. This corresponds to the assumptions that (i) the monetary authority sets V_t seeing lagged values of all the components of Z_t as well as the current values of X_t, and (ii) the current values of X_t do not respond contemporaneously to movements in V_t. The "shock" to monetary policy is the component of the innovation to V_t which is orthogonal to innovations in X_t.

Our first measure of the policy instrument, V_t, is the ratio of the log of nonborrowed reserves to the log of total reserves. Our decision to work with a nonborrowed reserves (*NBR*)-based measure of money rather than one based on broader monetary aggregates is motivated by arguments in Christiano and Eichenbaum [1992a, 1992b] and Strongin [1992]. The basic idea is that innovations to nonborrowed reserves primarily reflect exogenous shocks to monetary policy, while innovations to broader monetary aggregates primarily reflect shocks to money demand. While Christiano and Eichenbaum [1992a, 1992b] use *NBR* as the monetary aggregate in their analysis, Strongin [1992] argues that an even sharper measure of exogenous shocks to the money supply can be obtained using the ratio of *NBR* to total reserves. We denote this ratio by *NBRX*.[4] In our context, working with *NBR* or *NBRX* leads to qualitatively similar results.

Our second measure of shocks to monetary policy is motivated by arguments in McCallum [1983], Bernanke and Blinder [1992], and Sims [1992] that, at least relative to high-order monetary aggregates like M1 and M2, orthogonalized shocks to the federal funds rate are a better measure of shocks to monetary policy than orthogonalized shocks to the stock of money. Finally, our third measure of monetary policy shocks is motivated by results in Romer and Romer [1989], who use historical methods to identify specific periods in which the FOMC initiated contractionary changes in monetary policy.

III. Empirical Results

In reporting our empirical results, we display results using a benchmark specification, a broader *NBRX*-based measure of monetary policy shocks and a federal funds using policy shock measure. We then discuss results based on the Romer and Romer [1989] index of monetary policy contractions. To facilitate comparisons across the monetary policy measures, we normalize the policy shocks to be contractionary. Consequently, the *NBRX* innovation is negative. All results were generated using monthly data covering the sample period 1974:1–1990:5. The Appendix contains a descrip-

4. Strongin actually measures V_t as $NBR_t/(\text{Total Reserves})_{t-1}$ while we use $NBR_t/(\text{Total Reserves})_t$. This has virtually no impact on our results.

tion of our data. All VARs were estimated using six lags of all variables.[5]

We consider five nominal (spot) exchange rates, s_t^{For}, For = {Yen, Deutschmark (DM), Lira, French Franc (FF), U. K. Pound (PD)}. Here s_t^{For} denotes the logarithm of the number of U. S. dollars needed to buy one unit of the foreign currency at time t. Defined in this way, an increase in s_t^{For} corresponds to a depreciation of the U. S. dollar. In addition, we consider the logarithms of five real exchange rates, s_{Rt}^{For}, For = {Yen,DM, Lira,FF,PD}, defined as

$$s_{Rt}^{For} = s_t^{For} + P_t^{For} - P_t. \tag{2}$$

The variables P_t and P_t^{For} denote the time t U. S. and foreign price levels, respectively. Given this definition, s_{Rt}^{For} is the relative price of the foreign good in terms of the U. S. good. An increase in s_{Rt}^{For} denotes a depreciation of the U. S. real exchange rate.

We begin by reporting results from a benchmark five-variable VAR that includes U. S. industrial production (Y), the U. S. Consumer Price Level (P), the ratio of nonborrowed to total reserves ($NBRX$), a measure of the difference between U. S. and foreign short-term interest rates ($R^{For} - R^{US}$), and the real exchange rate (s_R^{For}).[6] All variables are in logarithms except for R^{For} and R^{US}. Dynamic response functions were calculated assuming a Wold ordering of $\{Y,P,NBRX,R^{For} - R^{US},s_R^{For}\}$. So here a contractionary monetary policy shock is measured as the component of a negative innovation to $NBRX_t$ that is orthogonal to P_t and Y_t.[7] Among other things, this corresponds to the assumption that the U. S. monetary authority looks at the contemporaneous values of P_t and Y_t when setting $NBRX_t$ but not $R_t^{For} - R_t^{US}$ or s_{Rt}^{For}. Notice that it is the *difference* between foreign and U. S. short-term nominal interest rates that enters into the analysis. Imposing this restriction is of interest for two reasons. First, a variety of authors like Meese and Rogoff [1983] consider theoretical and empirical models where it is the difference between foreign and U. S. interest rates that is relevant for exchange rate determination. Second, this

5. Our lag length was selected based on robustness of inference to higher order lags.

6. The short-term foreign interest rate, R^{For}, was measured using a short-term interest rate taken from the *International Financial Statistics* tape. The short-term U. S. interest rate, R^{US}, was measured using the three-month Treasury bill rate.

7. We found that our results were very robust to adopting different recursive orderings, such as putting $NBRX_t$ ahead of P_t and Y_t in the Wold ordering and putting P_t, Y_t, and s_{Rt}^{For} ahead of $NBRX_t$ in the Wold ordering.

system captures, in a parsimonious way, a subset of the key results that emerge from VARs where this restriction is not imposed (see below).

The first two rows of Figure I display the dynamic response functions of $R_t^{For} - R_t^{US}$ and s_{Rt}^{For} to a contractionary monetary policy shock. Solid lines represent our point estimates while dashed lines denote plus- and minus-one-standard-deviation bands.[8] We also conducted our analysis replacing the real exchange rate with the nominal exchange rate. The resulting dynamic response functions of $R_t^{For} - R_t^{US}$ are virtually identical to those reported in row 1, and are not reproduced here. Row 3 reports the dynamic response functions of s_t^{For} to the policy shock.

A number of important results emerge from Figure I. First, a contractionary shock to U. S. monetary policy leads to a persistent, significant decrease in the spread between foreign and U. S. nominal interest rates. For example, the initial impact of a one-standard-deviation negative shock to $NBRX_t$ is a {28,38,27,22,44} basis point decline in $\{R_t^{For} - R_t^{US}: For = \text{Yen,DM,Lira,FF,PD}\}$, respectively.[9] Second, the estimated impulse response functions of nominal and real exchange rates are very similar. This is consistent with the fact that movements in real and nominal exchange rates are highly correlated with each other (see, for example, Mussa [1986]). Third, a contractionary shock to U. S. monetary policy leads to persistent appreciations in nominal and real U. S. exchange rates. For example, the initial impact of a one-standard-deviation negative shock to $NBRX_t$ is a {0.28,0.50,0.42,0.36,0.28} percent fall in $\{s_{Rt}^{Yen}, s_{Rt}^{Dm}, s_{Rt}^{Lira}, s_{Rt}^{FF}, s_{Rt}^{PD}\}$, respectively, which represents an appreciation.

The maximal impact of the monetary shock on s_{Rt}^{For} and s_t^{For} does not occur contemporaneously. For example, the maximal impact on $\{s_t^{Yen}, s_t^{Dm}, s_t^{Lira}, s_t^{FF}, s_t^{PD}\}$, which equals {−1.91, −2.96, −2.95, −3.00, −1.86} percent, occurs {24,35,38,37,39} months after the monetary policy shock. This response pattern is inconsistent with simple overshooting models of the sort considered by Dornbusch [1976], since, in those models, a contractionary monetary policy shock generates a large initial appreciation in nominal (and real) exchange rates followed by subsequent depreciations. However, our results could be viewed as supporting a broader view

8. These were computed using the method described in Doan [1990], example 10.1, using 500 draws from the estimated asymptotic distribution of the vector autoregressive coefficients and covariance matrix of the innovations.

9. The shock to $NBRX_t$ equals −1.16 percent, −1.21 percent, −1.18 percent, −1.19 percent and −1.18 percent for the case in which Japan, Germany, Italy, France, and the United Kingdom are the foreign country included in the VAR.

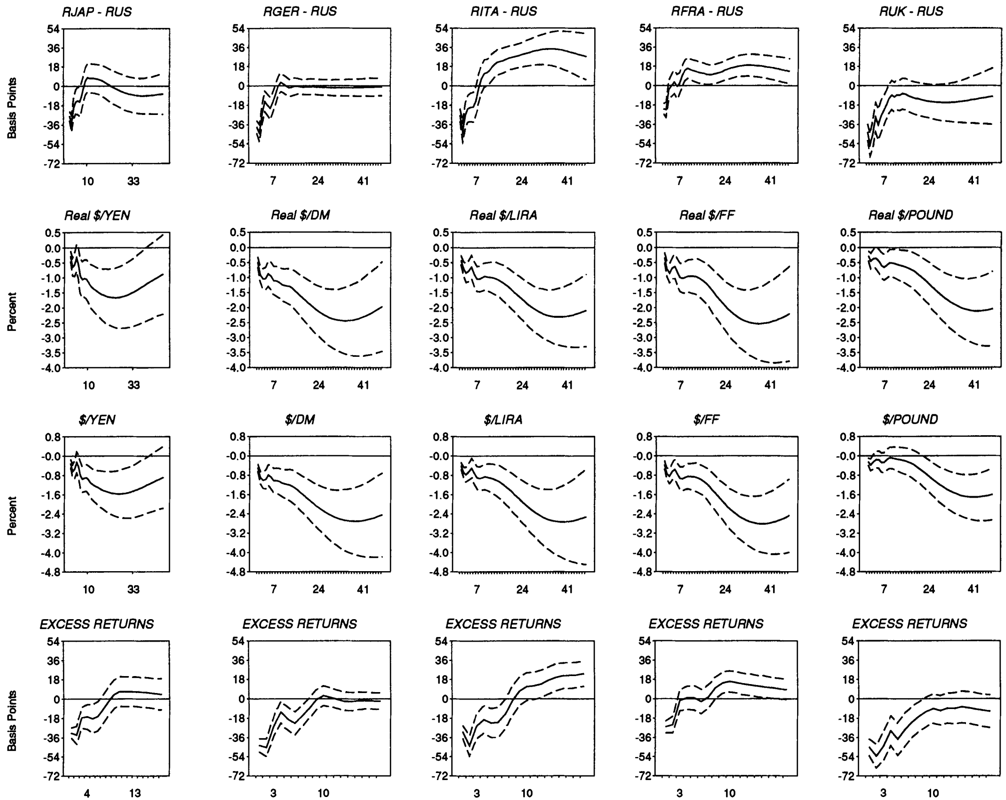

FIGURE I

Dynamic Response Functions: Benchmark Specification

Column 1 displays the dynamic effect of an orthogonalized negative innovation in *NBRX* on the difference between Japanese and U. S. interest rates (RJAP-RUS), the real U. S.–Japan exchange rate, the nominal U. S.–Japan exchange rate, and an uncovered interest parity of overshooting in which exchange rates eventually depreciate after appreciating for a period of time. In Section IV we discuss some recent work aimed at generating this type of response function.

Regardless of one's interpretation of the overshooting hypothesis, the estimated response path of s_t^{For} is inconsistent with uncovered interest rate parity. This is because uncovered interest rate parity implies that the fall in $R_t^{For} - R_t^{US}$ induced by contractionary monetary shock is offset by an expected depreciation of the dollar between time t and time $t + 1$. This prediction is at variance with the impulse response functions reported in Figure I. There we see that, in response to a contractionary monetary shock, $R_t^{For} - R_t^{US}$ falls while s_t^{For} *declines* between time t and time $t + 1$. So the time t expected one-period return on holding the foreign rather than the U. S. asset is lower for two reasons: (i) $R_t^{For} - R_t^{US}$ is lower, and (ii) the dollar is expected to *appreciate* between time t and time $t + 1$. The finding that uncovered interest rate parity does not hold is not new (see, for example, Hodrick [1987]). What is new is the finding that a monetary policy shock induces a systematic departure from uncovered interest rate parity.

To explore this issue in greater detail, it is useful to define Ψ_t^{For} as the ex post difference in the return between investing \$1 in one-period foreign bonds and investing \$1 in one-period U. S. bonds. Measured in U. S. dollars, this excess return, Ψ_t^{For}, is given by

$$\Psi_t^{For} = R_t^{For} - R_t^{US} + (s_{t+1}^{For} - s_t^{For}). \tag{3}$$

One implication of uncovered interest rate parity is that

$$E_t \Psi_{t+j}^{For} = 0 \tag{4}$$

for all $j \geq 0$, where E_t denotes the time t conditional expectation operator.

Given our estimated VARs, we can compute the dynamic response function of $E_t \Psi_{t+j}$ to a monetary policy shock. Under uncovered interest rate parity, this response function ought to be identically equal to zero. Row 4 of Figure I reports the point estimates (and standard errors) of the dynamic response function of $E_t \Psi_t^{For}$ that emerge from the unconstrained VAR underlying rows 1 and 2.[10] The key result here is that, for all cases, $E_t \Psi_t^{For}$ falls

10. Standard errors were computed using the method described in footnote 2, with one modification. For each Monte Carlo draw we computed the dynamic response function of $E_t \Psi_t^{For}$ to a monetary policy shock. The dotted lines in row 4 of Figure I correspond to a one-standard-deviation band for each coefficient in the dynamic response functions across the 500 Monte Carlo draws.

for a prolonged period of time after a contractionary monetary shock. And for each case we can easily reject the hypothesis that the individual coefficients in the response functions equal zero. We conclude that, after a contractionary monetary policy shock, the expected returns from investing in foreign short-term bonds falls relative to the returns from investing in short-term U. S. Treasury bills. Moreover, these excess returns are persistent. This persistence is consistent with the fact that future changes in the exchange rate tend to be *negatively* related to the forward premium.

In principle, one could construct a variety of statistics to summarize the "shape" of the impulse response functions as a way of characterizing the dynamic response of exchange rates to policy shocks. For example, we could ask whether various impulse response functions are identically equal to zero. We find it more revealing to consider the *average* response of s_{Rt}^{For} and s_t^{For} to a time t monetary shock over various time horizons, say from time $t + i$ to time $t + j$. We denote these responses by $\mu_{For,R}(i,j)$, and $\mu_{For}(i,j)$, respectively. In population these are equal to the average value of coefficients i through j of the corresponding impulse response functions.[11]

Results for s_{Rt}^{For} are reported in Table Ia. Row (1) reports the estimated correlation between the *innovation* to s_{Rt}^{For} and $NBRX_t$. Notice that in every case the estimated correlation is positive and significantly different from zero. Rows (2) through (7) report the estimated values of $\mu_{For,R}(i,j)$, $\{(i,j) = (1,6),(7,12). . . ,(31,36)\}$, respectively. For each country there exist a number of horizons for which we can reject, at conventional significance levels, the hypothesis that $\mu_{For,R}(i,j) = 0$. Indeed for Germany, France, and Italy, this hypothesis can be rejected for every specification of (i,j) at the 5 percent significance level. Consistent with Figure I, these rejections are not the strongest for the early periods.

Row (8) reports the maximal impact of a negative monetary policy shock on s_t^{For}. In every case the point estimate of this statistic

11. We cannot use the standard deviation bands about the estimated impulse response functions in Figure I to formally test hypotheses about $\mu_{For,R}(i,j)$ and $\mu_{For}(i,j)$. This is because each element in these bands summarizes the sampling uncertainty in the corresponding element of the estimated impulse response function, *not* taking into account the covariance between the different coefficients. To deal with this problem, we calculated standard errors for these statistics using the method described in footnote 2, with one modification. For each Monte Carlo draw we computed the values of $\mu_{For,R}(i,j)$ and $\mu_{For}(i,j)$. We then calculated the standard deviation of these statistics across the 500 Monte Carlo draws. Alternatively, inference could be based on the empirical distribution function of these statistics. In practice, we found that inference was very robust across the two procedures.

TABLE Ia
BENCHMARK SPECIFICATION REAL EXCHANGE RATES

	Dynamic response functions				
	Japan	Germany	Italy	France	United Kingdom
(1) Corr(*NBRX,EXCH*)	0.155	0.266	0.222	0.204	0.169
Std. error	0.068	0.068	0.068	0.073	0.070
Significance	0.011	0.000	0.001	0.003	0.008
(2) 1–6 months	−0.552	−0.888	−0.714	−0.679	−0.424
Std. error	0.363	0.301	0.274	0.290	0.274
Significance	0.064	0.002	0.005	0.010	0.061
(3) 7–12 months	−1.140	−1.195	−1.003	−0.980	−0.565
Std. error	0.648	0.504	0.438	0.512	0.438
Significance	0.039	0.009	0.011	0.028	0.098
(4) 13–18 months	−1.490	−1.485	−1.111	−1.120	−0.673
Std. error	0.807	0.648	0.558	0.634	0.550
Significance	0.032	0.011	0.023	0.039	0.110
(5) 19–24 months	−1.647	−1.957	−1.520	−1.693	−1.024
Std. error	0.945	0.780	0.651	0.759	0.662
Significance	0.041	0.006	0.010	0.013	0.061
(6) 25–30 months	−1.667	−2.131	−1.720	−1.953	−1.224
Std. error	0.997	0.841	0.690	0.821	0.713
Significance	0.047	0.006	0.006	0.009	0.043
(7) 31–36 months	−1.460	−2.437	−2.248	−2.499	−1.890
Std. error	1.133	1.087	0.851	1.104	0.916
Significance	0.099	0.012	0.004	0.012	0.020
(8) Max impact	−2.032	−2.679	−2.474	−2.748	−2.283
Std. error	1.033	1.226	1.031	1.361	1.159
Significance	0.025	0.014	0.008	0.022	0.024
(9) Max month	23.650	32.070	36.498	36.162	39.754
Std. error	11.818	10.851	9.480	8.209	9.704
Significance	0.023	0.002	0.000	0.000	0.000
	Variance decompositions				
(10) 31–36 months	23.016	42.917	38.122	37.520	26.153
Std. error	13.640	15.713	15.481	14.877	15.034
Significance	0.092	0.006	0.014	0.012	0.082

is negative and substantially larger (in absolute value) than $\mu_{For,R}(1,6)$. Also notice that in every case we strongly reject the hypothesis that the maximal impact of a contractionary monetary policy shock is equal to zero. Row (9) reports the time to the maximal appreciation in the real exchange rate following a policy shock. While there is substantial uncertainty about the exact time period when the maximal appreciation occurs, for every country, we can easily reject the hypothesis that it occurs contemporaneously. Table Ib is the exact analog to Table Ia except that it is based on VARs that include s_t^{For} rather than s_{Rt}^{For}. As before, using nominal rather than real exchange rates has very little impact on inference. Table V reports the average response of Ψ_t^{For} in the first and second half year's horizons after a shock to monetary policy. Consistent with the failure of uncovered interest rate parity, for every country, we can easily reject the hypothesis that the average response of $E_t \Psi_t^{For}$ in the first six months after a monetary shock is zero. The extent of the excess returns ranges from nine basis points (France) to 40 basis points (United Kingdom).

We conclude this subsection by discussing the overall contribution of monetary shocks to the variability of exchange rates. To this end, we computed the percentage of the variance of the k step ahead forecast error that is attributable to monetary shocks. As k goes to infinity, this corresponds to the percentage of the variance of exchange rates that is due to monetary shocks. Row (10) of Tables Ia and Ib reports the average of this percentage over the 31- to 36-month horizon for real and nominal exchange rates, respectively. The estimated percentages range from a low of 18 percent (United Kingdom, nominal exchange rates) to a high of 43 percent (Germany, real). While there is substantial sampling uncertainty associated with these point estimates, in the case of Germany, Italy, and France, we can easily reject the hypothesis that the percentage is zero, for either real or nominal exchange rates. The rejections are more marginal for Japan and the United Kingdom.

An important restriction of our benchmark specification is the assumption that only the difference between foreign and U. S. interest rates is relevant for exchange rate determination. While this restriction is quite natural from the perspective of various theoretical models, it is desirable to assess the impact of relaxing it. To this end, we now discuss the results of considering a specification in which foreign and U. S. interest rates enter separately. There are two additional advantages to doing this. First, we can explicitly assess the impact of policy shocks on the level of domestic and foreign interest rates. Second, we can more easily compare results obtained with *NBRX*-based policy shock measures with those obtained using interest-rate-based policy shock measures.

In expanding the benchmark specification, we must deal with the issue of just how many variables to include in the analysis. This

TABLE Ib
BENCHMARK SPECIFICATION, NOMINAL EXCHANGE RATES

	Japan	Germany	Italy	France	United Kingdom
	Dynamic response functions				
(1) Corr(*NBRX,EXCH*)	0.150	0.260	0.225	0.207	0.139
Std. error	0.069	0.066	0.063	0.071	0.068
Significance	0.015	0.000	0.000	0.002	0.021
(2) 1–6 months	−0.499	−0.843	−0.622	−0.656	−0.244
Std. error	0.340	0.310	0.279	0.287	0.262
Significance	0.071	0.003	0.013	0.011	0.176
(3) 7–12 months	−1.005	−1.097	−0.897	−0.885	−0.145
Std. error	0.561	0.542	0.507	0.508	0.434
Significance	0.037	0.021	0.039	0.041	0.370
(4) 13–18 months	−1.333	−1.393	−1.096	−1.049	−0.239
Std. error	0.718	0.694	0.661	0.616	0.549
Significance	0.032	0.022	0.049	0.044	0.332
(5) 19–24 months	−1.525	−1.926	−1.639	−1.697	−0.632
Std. error	0.878	0.840	0.802	0.732	0.632
Significance	0.041	0.011	0.021	0.010	0.159
(6) 25–30 months	−1.566	−2.145	−1.894	−2.001	−0.854
Std. error	0.942	0.914	0.871	0.791	0.671
Significance	0.048	0.009	0.015	0.006	0.101
(7) 31–36 months	−1.437	−2.650	−2.613	−2.720	−1.535
Std. error	1.093	1.236	1.227	1.052	0.819
Significance	0.094	0.016	0.017	0.005	0.030
(8) Max impact	−1.913	−2.961	−2.950	−3.000	−1.859
Std. error	0.952	1.532	1.815	1.330	0.983
Significance	0.022	0.027	0.052	0.012	0.029
(9) Max month	24.654	35.304	37.990	37.478	38.872
Std. error	11.818	10.412	7.360	6.918	9.553
Significance	0.018	0.000	0.000	0.000	0.000
	Variance decompositions				
(10) 31–36 months	22.084	41.021	38.767	38.474	18.752
Std. error	13.901	16.271	15.135	15.879	12.428
Significance	0.112	0.012	0.010	0.015	0.131

decision involves the following trade-off. To minimize omitted variable bias, we would like to include as many variables as possible in the analysis. But we cannot ignore the problem of parameter profligacy. If we include k lags of n variables in the analysis, we have to estimate $(k \times n^2)$ free parameters. As n expands, our degrees of freedom rapidly disappear, and inference becomes impossible. To deal with this problem, we decided to treat the United States and foreign countries in an asymmetric manner. Specifically, while we included a narrow U. S. monetary aggregate in the analysis, we did not include a narrow foreign monetary aggregate. This decision was based on a number of considerations. First very narrow monetary aggregates like *NBRX* are not available for countries like the United States. Second, including a broad monetary aggregate seemed to have little added value given our objective of identifying shocks to U. S. monetary policy. Moreover, Sims [1992] argues that shocks to foreign monetary policy are better captured by orthogonalized shocks to foreign interest rates than by orthogonalized shocks to broad foreign monetary aggregates. Since we include foreign interest rates in VARs, the foreign monetary authority's reaction function is, in principle, included in the analysis. Because the VARs are unconstrained, the foreign monetary policy reaction can vary across the countries. Of course, our results could be sensitive to including broad foreign monetary aggregates. Fortunately, they are not, at least from a qualitative point of view. On the same basis we also did not include a measure of the foreign price level in our VARs.

The first three rows of Figure II display results from a seven-variable VAR that includes U. S. industrial production (Y), the U. S. Consumer Price Level (P), foreign output (Y^{For}), the foreign interest rate (R^{For}), the ratio of *NBR* to *TR* (*NBRX*), the three-month U. S. Treasury bill rate (R^{US}), and the real exchange rate (s_R^{For}). All variables are in logarithms except for R^{For} and R^{US}. Impulse response functions were calculated assuming a Wold ordering of $\{Y,P,Y^{For},R^{For},NBRX,R^{US},s_R^{For}\}$. Among other things, this corresponds to the assumption that the contemporaneous portion of the feedback rule for setting $NBRX_t$ involves $(Y_t,P_t,Y_t^{For}, R_t^{For})$ but not R_t^{US} or s_{Rt}^{For}. Rows 1, 2, and 3 report the estimated dynamic response functions of R_t^{US}, R_t^{For}, and s_{Rt}^{For}, respectively, to a one-standard-deviation negative monetary policy shock. We also conducted our analysis replacing the real exchange rate with the corresponding nominal exchange rate. The resulting dynamic response functions of R_t^{US} and R_t^{For} are virtually identical to those reported in rows 1 and 2. Rows 4 and 5 report the dynamic response functions of s_t^{For} and $E_t\Psi_t^{For}$ *to the policy shock.*

Comparing Figures I and II, we see that our key results are robust to departing from the benchmark specification. Specifically, according to Figure II, a contractionary shock to U. S. monetary

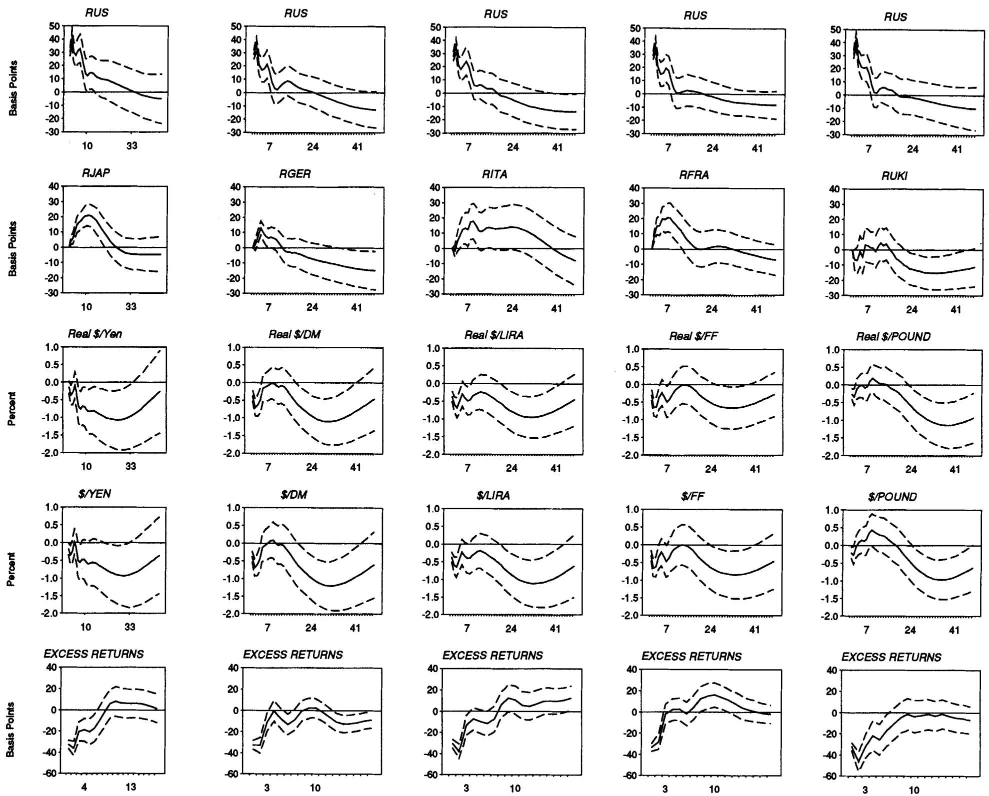

FIGURE II

Dynamic Response Functions: Orthogonalized Shocks in *NBRX*, Seven-Variable System

Column 1 displays the dynamic effect of an orthogonalized, negative innovation in *NBRX* on the U. S. interest rate (RUS), the Japanese interest rate (RJAP), the real U. S.–Japan exchange rate, the nominal U. S.–Japan exchange rate, and an uncovered interest parity policy leads to a sharp, persistent increase in the U. S. interest rate as well as a persistent rise in all of the foreign interest rates (except the United Kingdom). In all cases, the increase in R_t^{US} exceeds the corresponding increase in R_t^{For}. So, consistent with Figure I, the shock leads to a fall in $R_t^{For} - R_t^{US}$. From the perspective of the arguments in Sims [1992], this can be interpreted as reflecting a policy in which foreign monetary authorities initially only partially accommodate the increase in U. S. interest rates. Also note that, as before, a negative monetary shock leads to pronounced, persistent appreciations in real and nominal U. S. exchange rates. Not surprisingly, given the large number of variables in the VAR (and the correspondingly large number of parameters that must be estimated), the impulse response functions of s_{Rt}^{For} and s_t^{For} are less precisely estimated than in the benchmark specification. Tables IIa and IIb, which are the exact analogs to Tables Ia and Ib, confirm this impression. In particular, we find substantially less evidence against the hypotheses that $\mu_{For,R}(i,j)$ and $\mu_{For}(i,j)$ are equal to zero in population. Still for each country there exists at least one specification of (i,j) for which we can reject, at the 10 percent significance level (or better), these hypotheses. Moreover, for every country we can reject, at the 10 percent significance level (or better), these hypotheses. Moreover, for every country we can reject, at the 5 percent significance level or better, the hypothesis that the maximal fall in s_t^{For} after a contractionary monetary policy shock is zero. Finally, for all countries except Japan, we can reject, at the 5 percent significance level, the hypothesis that the correlation between the innovations to *NBRX* and s_{Rt}^{For} (or s_t^{For}) is equal to zero. For Japan this hypothesis can be rejected at the 8 percent significance level.

As before, our results indicate that $E_t\Psi_t^{For}$ falls for a substantial amount of time (see row 5 of Figure II). Interestingly, despite the large dimensionality of the VAR, the dynamic response functions of $E_t\Psi_t^{For}$ are estimated quite accurately. This is confirmed by the formal tests reported in Table V. We conclude that the failure of the strict overshooting hypothesis and the emergence of expected excess returns is robust to allowing R^{For} and R^{US} to enter the VARs separately.

Finally, row (10) of Tables IIa and IIb reports the average percentage of the forecast error variance over the 31- to 36-month horizon for real and nominal exchange rates that is attributable to monetary shocks. Notice that the estimated percentages are lower than those emerging from the five-variable VAR and now range

TABLE IIa
NBRX-BASED MEASURES OF POLICY SHOCKS, SEVEN-VARIABLE SYSTEM, REAL EXCHANGE RATES

	Dynamic response functions				
	Japan	Germany	Italy	France	United Kingdom
(1) Corr(*NBRX,EXCH*)	0.104	0.254	0.226	0.219	0.156
Std. error	0.072	0.064	0.069	0.068	0.068
Significance	0.075	0.000	0.001	0.001	0.011
(2) 1–6 months	−0.317	−0.416	−0.483	−0.432	−0.140
Std. error	0.309	0.259	0.246	0.256	0.248
Significance	0.153	0.054	0.025	0.046	0.286
(3) 7–12 months	−0.764	−0.071	−0.351	−0.258	0.080
Std. error	0.550	0.446	0.427	0.473	0.387
Significance	0.083	0.437	0.206	0.292	0.582
(4) 13–18 months	−0.871	−0.350	−0.358	−0.072	−0.066
Std. error	0.705	0.527	0.523	0.528	0.475
Significance	0.108	0.253	0.247	0.446	0.445
(5) 19–24 months	−1.027	−0.815	−0.667	−0.406	−0.417
Std. error	0.786	0.564	0.547	0.557	0.519
Significance	0.096	0.074	0.111	0.233	0.211
(6) 25–30 months	−1.062	−0.952	−0.797	−0.532	−0.613
Std. error	0.815	0.587	0.557	0.578	0.536
Significance	0.096	0.052	0.076	0.178	0.126
(7) 31–36 months	−0.944	−1.056	−0.932	−0.650	−1.085
Std. error	0.922	0.689	0.603	0.591	0.603
Significance	0.153	0.063	0.061	0.136	0.036
(8) Max impact	−1.450	−1.319	−1.190	−1.008	−1.271
Std. error	0.889	0.623	0.516	0.395	0.610
Significance	0.051	0.017	0.011	0.005	0.019
(9) Max month	21.528	24.192	23.882	18.692	34.508
Std. error	11.404	13.522	14.101	15.303	11.796
Significance	0.030	0.037	0.045	0.111	0.002
	Variance decompositions				
(10) 31–36 months	13.263	12.983	13.535	8.372	10.687
Std. error	10.677	8.830	10.324	6.448	7.814
Significance	0.214	0.144	0.190	0.194	0.171

TABLE IIb
NBRX-BASED MEASURES OF POLICY SHOCKS, SEVEN-VARIABLE SYSTEM, NOMINAL EXCHANGE RATES

		Dynamic response functions				
		Japan	Germany	Italy	France	United Kingdom
(1)	Corr(*NBRX,EXCH*)	0.102	0.249	0.238	0.213	0.120
	Std. error	0.071	0.065	0.069	0.066	0.073
	Significance	0.076	0.000	0.000	0.001	0.049
(2)	1–6 months	−0.233	−0.380	−0.441	−0.415	−0.029
	Std. error	0.291	0.276	0.252	0.267	0.259
	Significance	0.211	0.084	0.040	0.060	0.455
(3)	7–12 months	−0.539	0.010	−0.302	−0.214	0.329
	Std. error	0.559	0.494	0.442	0.499	0.434
	Significance	0.168	0.508	0.247	0.334	0.775
(4)	13–18 months	−0.608	−0.305	−0.345	−0.090	0.154
	Std. error	0.674	0.593	0.518	0.578	0.497
	Significance	0.184	0.303	0.253	0.438	0.622
(5)	19–24 months	−0.797	−0.822	−0.707	−0.490	−0.268
	Std. error	0.752	0.612	0.571	0.607	0.515
	Significance	0.145	0.090	0.108	0.209	0.301
(6)	25–30 months	−0.864	−0.984	−0.870	−0.640	−0.481
	Std. error	0.791	0.625	0.593	0.622	0.522
	Significance	0.137	0.058	0.071	0.152	0.178
(7)	31–36 months	−0.886	−1.175	−1.104	−0.834	−0.942
	Std. error	0.921	0.718	0.676	0.672	0.543
	Significance	0.168	0.051	0.051	0.107	0.042
(8)	Max impact	−1.315	−1.390	−1.327	−1.148	−1.109
	Std. error	0.862	0.676	0.610	0.567	0.515
	Significance	0.063	0.020	0.015	0.022	0.016
(9)	Max month	22.690	26.236	27.454	23.054	32.974
	Std. error	12.589	13.373	13.340	15.630	11.495
	Significance	0.036	0.025	0.020	0.070	0.002
		Variance decompositions				
(10)	31–36 months	11.179	13.271	13.743	8.634	9.406
	Std. error	9.497	9.329	9.601	7.004	6.134
	Significance	0.239	0.155	0.152	0.218	0.125

from a low of 8 percent (France, real exchange rates) to a high of 14 percent (Italy, nominal exchange rates). In addition, the standard errors of these statistics are substantially larger than before.

We now consider results obtained measuring monetary policy shocks as an orthogonalized component of the innovation to the federal funds rate. Figure III reports results from a seven-variable VAR that includes data on U. S. industrial production (Y), the U. S. Consumer Price Level (P), foreign output (Y^{For}), the foreign interest rate (R^{For}), the federal funds rate (FF), the ratio of NBR to TR ($NBRX$), and the real exchange rate (s_{Rt}^{For}). All variables are in logarithms except R^{For} and FF. Impulse response functions were calculated assuming a Wold ordering of $\{Y,P,Y^{For}, R^{For},FF,NBRX, s_R^{For}\}$. A monetary policy shock is identified as the component of the innovation in FF_t that is orthogonal to Y_t, P, Y^{For}, and R^{For}. Among other things, this corresponds to the assumption that the contemporaneous portion of the feedback rule for setting FF_t involves $(Y_t,P_t,Y_t^{For},R_t^{For})$ but not $NBRX_t$ or s_{Rt}^{For}. We also conducted our analysis using nominal exchange rates rather than real exchange rates. The resulting dynamic response functions of R_t^{US} and R_t^{For} are virtually identical to those reported in rows 1 and 2.

Our results are qualitatively very similar to those obtained with the $NBRX$-based measures of policy shocks. First, although not reported, we find that, consistent with the presence of a strong liquidity effect, a positive shock to the federal funds rate generates sharp, persistent declines in $NBRX$. Second, from Figure III we see that a contractionary monetary policy shock (i.e., a *positive* shock to the federal funds rate) is associated with persistent appreciations in nominal and real U. S. exchange rates. For example, the initial impact of an approximately 60-basis-point positive shock to the federal funds rate is a $\{0.31,0.46,0.40,0.38,0.15\}$ percent decline in $\{s_{Rt}^{Yen},s_{Rt}^{DM},s_{Rt}^{Lira},s_{Rt}^{FF},s_{Rt}^{PD}\}$, respectively. Third, the maximal impact of the monetary shocks on s_{Rt}^{For} and s_t^{For} does not occur contemporaneously. For example, the maximal impact on $\{s_t^{Yen},s_t^{DM}, s_t^{Lira},s_t^{FF},s_t^{PD}\}$ of an approximately 60 basis-point shock to FF_t is a $\{1.71,2.00,1.81,1.96,1.15\}$ percent fall that occurs $\{22,31,33,32,30\}$ months later. Fourth, the dynamic response functions of real and nominal exchange rates to monetary shocks are very similar. And, consistent with results of the previous subsections, a contractionary monetary policy shock is associated with persistent, significant increases in the returns to investing in short-term U. S. bills versus foreign bills ($E_t\Psi_t^{For}$).

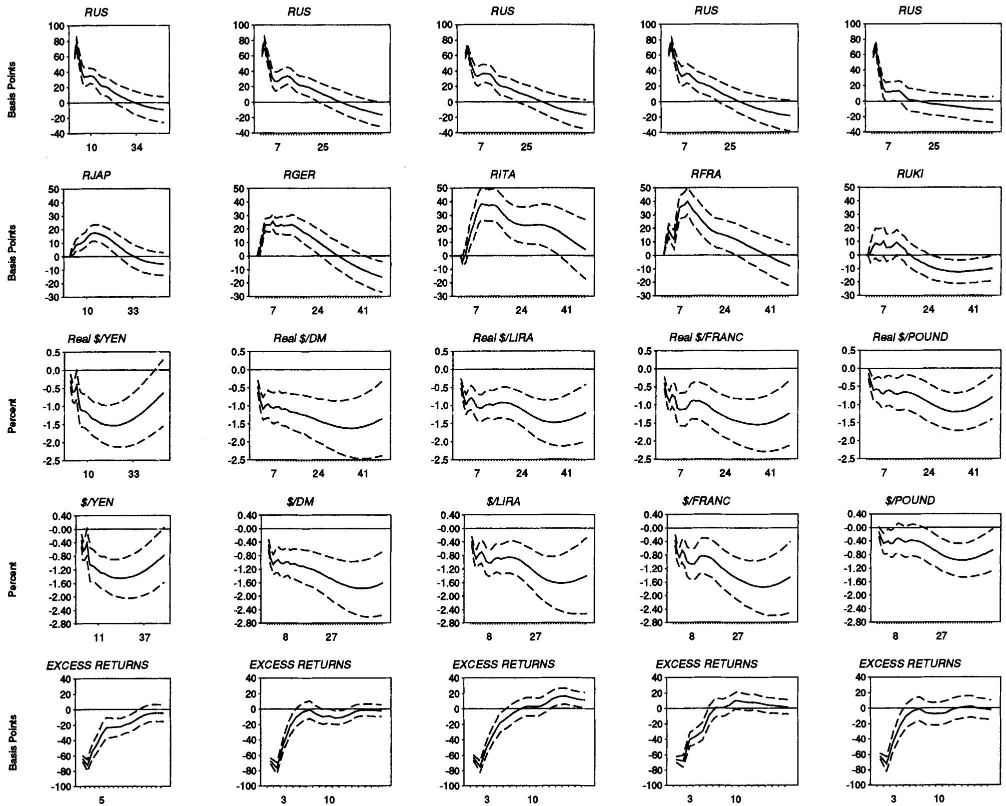

FIGURE III

Dynamic Response Functions: Orthogonalized Shock in Federal Funds Rate, Seven-Variable System

Column 1 displays the dynamic effect of an orthogonalized innovation in the Federal Funds rate on the Federal Funds rate (RUS), the

Interestingly, the dynamic response functions of s_{Rt}^{For} and s_t^{For} are estimated more precisely than they were with the *NBRX*-based policy shock measures. This can be seen informally by comparing the relevant standard deviations bands in Figures II and III. This impression is confirmed by Tables IIIa and IIIb, which are the exact analogs to Tables IIa and IIb.

A number of key results emerge from these tables. First, innovations to the federal funds rate are *negatively* correlated with innovations to nominal and real exchange rates (see Row (1)). The hypothesis that these correlations equal zero in population can be easily rejected for the Japanese, German, Italian, and French cases. The rejection is more marginal for the United Kingdom. Second, there is very strong statistical evidence that monetary policy shocks affect real and nominal exchange rates. For example, except for the United Kingdom, the hypothesis that $\mu_{For,R}(i,j)$ equals zero can be rejected, at the 4 percent significance level or better, for all six specifications of (i,j). In the U. K. case we can reject this hypothesis at the 5 percent significance level in four out of six specifications (i,j). Third, the hypothesis that the maximal impact of a monetary policy shock on s_{Rt}^{For} and s_t^{For} equals zero can be strongly rejected (see row (8)). Fourth, we find substantial evidence that the maximal effect of a policy shock does not occur contemporaneously (see row (9)). Finally, Table V indicates that we can easily reject the hypothesis that the average response of $E_t\Psi_t^{For}$ for the first half year after a policy shock is equal to zero.

Tables IIIa and IIIb reports the average percentage of the forecast error variance over the 31-to-36-month horizon for real and nominal exchange rates that is attributable to moneary shocks. For all countries, except the United Kingdom, monetary shocks are estimated to account for over 20 percent of the variance of real and nominal exchange rates. Also notice that there is less sampling uncertainty with this measure of monetary shocks than with *NBRX*-based measures. So once we move to federal funds-based measures of policy shocks, we find substantial evidence that an important percentage of the variability of exchange rates can be attributed to policy shocks.

We now report results obtained using the Romer and Romer [1989] index of monetary policy. Figure IV reports results obtained from a VAR that includes U. S. industrial production (Y), the U. S. Consumer Price Level (P), foreign output (Y^{For}), the foreign interest rate (R^{For}), the ratio of *NBR* to *TR* (*NBRX*), the real exchange rate (s_R^{For}), and the federal fund rate (*FF*). All variables

TABLE IIIa
FED-FUNDS-BASED MEASURES OF POLICY SHOCKS, SEVEN-VARIABLE SYSTEM, REAL EXCHANGE RATES

	Dynamic response functions				
	Japan	Germany	Italy	France	United Kingdom
(1) Corr(*FF,EXCH*)	−0.151	−0.269	−0.231	−0.228	−0.105
Std. error	0.072	0.070	0.065	0.069	0.069
Significance	0.018	0.000	0.000	0.001	0.063
(2) 1–6 months	−0.627	−0.901	−0.779	−0.773	−0.498
Std. error	0.309	0.274	0.238	0.268	0.254
Significance	0.021	0.001	0.001	0.002	0.025
(3) 7–12 months	−1.211	−1.071	−1.026	−1.034	−0.663
Std. error	0.448	0.424	0.358	0.452	0.404
Significance	0.003	0.006	0.002	0.011	0.051
(4) 13–18 months	−1.440	−1.203	−0.950	−0.949	−0.677
Std. error	0.504	0.505	0.412	0.515	0.481
Significance	0.002	0.009	0.011	0.033	0.080
(5) 19–24 months	−1.532	−1.338	−1.025	−1.227	−0.844
Std. error	0.565	0.570	0.448	0.558	0.513
Significance	0.003	0.009	0.011	0.014	0.050
(6) 25–30 months	−1.535	−1.413	−1.121	−1.330	−0.951
Std. error	0.597	0.605	0.471	0.587	0.523
Significance	0.005	0.010	0.009	0.012	0.034
(7) 31–36 months	−1.328	−1.611	−1.453	−1.532	−1.200
Std. error	0.722	0.754	0.596	0.686	0.511
Significance	0.033	0.016	0.007	0.013	0.009
(8) Max impact	−1.795	−1.902	−1.675	−1.791	−1.371
Std. error	0.608	0.803	0.627	0.717	0.497
Significance	0.002	0.009	0.004	0.006	0.003
(9) Max month	21.632	28.228	28.590	28.192	29.040
Std. error	10.521	15.195	14.569	14.138	12.373
Significance	0.020	0.032	0.025	0.023	0.009
	Variance decompositions				
(10) 31–36 months	21.642	26.542	25.399	24.730	16.957
Std. error	10.456	11.456	10.093	11.733	10.052
Significance	0.039	0.021	0.012	0.035	0.092

TABLE IIIb
FED-FUNDS-BASED MEASURES OF POLICY SHOCKS, SEVEN-VARIABLE SYSTEM, NOMINAL EXCHANGE RATES

	Dynamic response functions				
	Japan	Germany	Italy	France	United Kingdom
(1) Corr(*FF,EXCH*)	−0.155	−0.269	−0.238	−0.217	−0.094
Std. error	0.075	0.066	0.065	0.071	0.073
Significance	0.019	0.000	0.000	0.001	0.100
(2) 1–6 months	−0.625	−0.887	−0.718	−0.736	−0.380
Std. error	0.319	0.263	0.241	0.278	0.259
Significance	0.025	0.000	0.001	0.004	0.071
(3) 7–12 months	−1.172	−1.027	−0.955	−0.974	−0.384
Std. error	0.450	0.399	0.387	0.460	0.422
Significance	0.005	0.005	0.007	0.017	0.181
(4) 13–18 months	−1.360	−1.132	−0.879	−0.913	−0.396
Std. error	0.499	0.482	0.456	0.541	0.464
Significance	0.003	0.009	0.027	0.046	0.196
(5) 19–24 months	−1.448	−1.291	−0.996	−1.254	−0.570
Std. error	0.551	0.547	0.519	0.576	0.474
Significance	0.004	0.009	0.027	0.015	0.114
(6) 25–30 months	−1.455	−1.382	−1.125	−1.392	−0.681
Std. error	0.577	0.584	0.556	0.599	0.478
Significance	0.006	0.009	0.022	0.010	0.077
(7) 31–36 months	−1.327	−1.699	−1.572	−1.710	−0.960
Std. error	0.679	0.728	0.735	0.730	0.485
Significance	0.025	0.010	0.016	0.010	0.024
(8) Max impact	−1.707	−2.004	−1.807	−1.956	−1.154
Std. error	0.635	0.765	0.985	0.870	0.486
Significance	0.004	0.004	0.033	0.012	0.009
(9) Max month	22.560	31.716	32.604	31.978	29.616
Std. error	10.936	14.907	13.085	12.871	13.436
Significance	0.020	0.017	0.006	0.006	0.014
	Variance decompositions				
(10) 31–36 months	22.908	25.966	23.155	26.749	11.571
Std. error	10.853	11.208	10.250	12.145	7.933
Significance	0.035	0.021	0.024	0.028	0.145

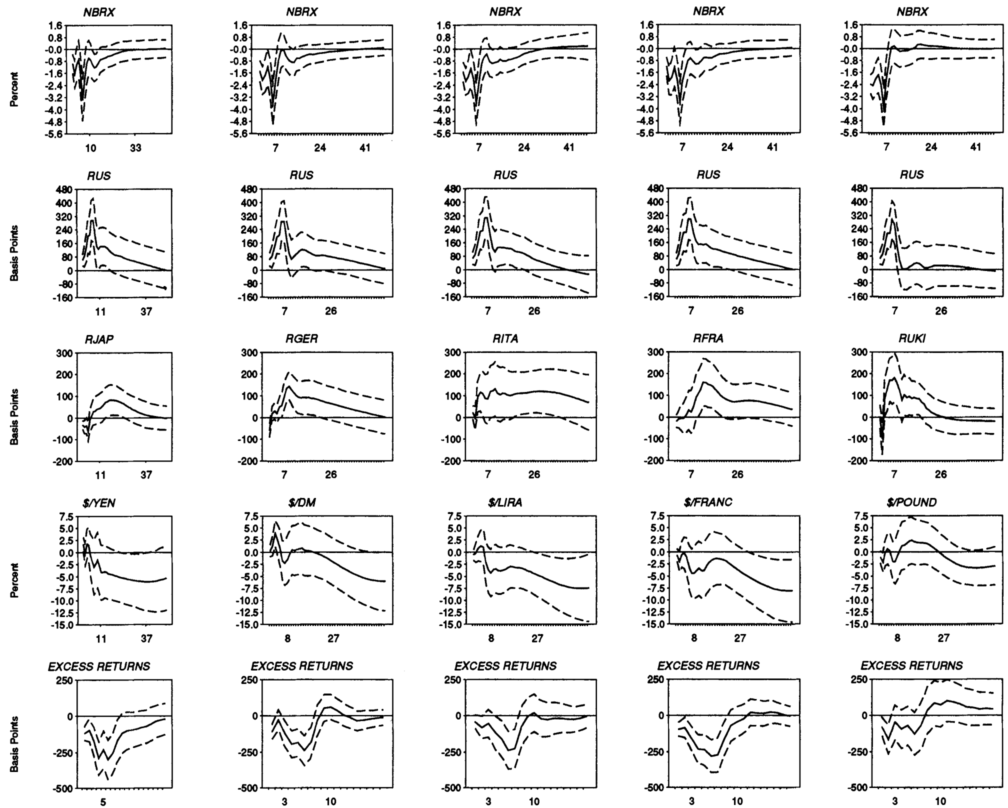

FIGURE IV

Dynamic Response Functions: Romer and Romer Episode, Eight-Variable System

Columns 1 displays the dynamic effect of a Romer and Romer episode on *NBRX*, the Federal Funds rate (RUS), the Japanese interest rate (RJAP), the nominal U.S.-Japan exchange rate, and an uncovered interest parity condition. Columns 2 through 5 do the same for are in logarithms except R^{For} and FF. In addition, the VAR includes the Romer and Romer index of monetary policy. Specifically, we consider a VAR for the vector of variables Z_t:

$$Z_t = A(L)Z_{t-1} + \beta(L)d_t + \epsilon_t. \tag{5}$$

Here $A(L)$ and $\beta(L)$ are one-sided polynomials in the lag operator L, and the vector Z_t equals $[Y_t, P_t, Y_t^{For}, R_t^{For}, NBRX_t, s_{Rt}^{For}, FF_t]'$. The variable d_t denotes the time t value of the Romer and Romer index. This variable equals one for the month at which a Romer and Romer episode begins, and zero otherwise. The response of Z_{t+k} to a time t Romer and Romer monetary contraction ($d_t = 1, d_{t+k} = 0$ for $k > 0$) is given by the coefficient on L^k in the polynomial $[I - A(L)]^{-1}\beta(L)$.[12]

Rows 1 and 2 of Figures IV provide corroborating evidence that the Romer and Romer [1989] dummy variables do indeed correspond to monetary policy contractions. In particular, a unit increase in the Romer and Romer index is associated with a sharp, persistent increase in the federal funds rate and a decrease in *NBRX*. Notice that the maximal increase in the federal funds rate and the maximal decrease in *NBRX* do not occur at the time of the change in the index. Instead both occur six months later. The initial change in the federal funds rate equals roughly 50 points. Six months later the federal funds rate is almost 300 basis points higher than it was initially. So Romer and Romer episodes correspond to large monetary contractions, at least relative to the types of shocks considered earlier. Recall that we obtain very similar results irrespective of whether we work with *NBRX* or federal funds rate-based measures of monetary shocks. In light of this, it is not surprising that the dynamic impulse responses functions of *NBRX* and the federal funds rate to a change in the Romer and Romer index appear to be mirror images of each other.

The fact that the peak effect of a change in the Romer and Romer on *NBRX* and the federal funds rate occurs with a six-month delay helps explain the dynamic response functions of s_{Rt}^{For} and s_t^{For}.[13] The initial response of real and nominal exchange rates is either very close to zero or slightly negative. But in all cases, after six months, real and nominal exchange rates undergo

12. The dates of the Romer and Romer [1989] episodes are 1974:4, 1978:8, and 1979:10. Since our sample ends after theirs, we included a dummy variable for the period 1988:8 suggested by Oliner and Rudebusch [1992].

13. Since these response functions are so similar, only the first is reported in Figure IV.

TABLE IVa
ROMER-BASED MEASURES OF POLICY SHOCKS, SEVEN-VARIABLE SYSTEM, REAL EXCHANGE RATES

	Dynamic response functions				
	Japan	Germany	Italy	France	United Kingdom
(1) 1–6 months	0.188	1.011	−0.095	−1.757	−1.964
Std. error	3.088	2.701	2.497	2.622	2.515
Significance	0.524	0.646	0.485	0.251	0.217
(2) 7–12 months	−3.433	−0.832	−3.675	−5.207	−2.647
Std. error	5.859	4.627	4.538	5.057	4.369
Significance	0.279	0.429	0.209	0.152	0.272
(3) 13–18 months	−4.843	−0.130	−3.390	−2.985	−0.592
Std. error	5.451	4.644	4.350	5.004	4.569
Significance	0.187	0.489	0.218	0.275	0.448
(4) 19–24 months	−5.602	−1.120	−3.830	−3.118	−0.928
Std. error	5.131	4.472	3.928	4.738	4.518
Significance	0.137	0.401	0.165	0.255	0.419
(5) 25–30 months	−5.739	−1.690	−4.239	−3.842	−1.603
Std. error	5.168	4.555	3.986	4.740	4.488
Significance	0.133	0.355	0.144	0.209	0.361
(6) 31–36 months	−5.706	−4.161	−5.704	−6.293	−4.202
Std. error	5.525	5.438	4.682	5.580	4.501
Significance	0.151	0.222	0.112	0.130	0.175
(7) Max impact	−8.932	−7.496	−8.513	−10.126	−7.735
Std. error	5.144	4.530	5.061	7.035	4.206
Significance	0.041	0.049	0.046	0.075	0.033
(8) Max month	24.126	31.888	30.322	29.224	23.092
Std. error	13.856	17.854	16.599	18.110	18.591
Significance	0.041	0.037	0.034	0.053	0.107

persistent appreciations. This is consistent with our earlier results. The same is true for excess returns, $E_t\Psi_t^{For}$. The large responses of FF_t, R_t^{For}, s_t^{For}, s_{Rt}^{For}, and Ψ_t^{For} reflect the magnitude of the Romer and Romer episodes. The main impact of working with the Romer and Romer index is that the dynamic response functions s_t^{For} and s_{Rt}^{For} are measured with much less precision than they were when we worked with the other policy shock measures. This is not surprising in light of the small number of Romer and Romer contractions. Tables IVa and IVb, which report the estimated values of $\mu_{for,R}(i,j)$ and $\mu_{For}(i,j)$, $\{(i,j) = (1,6), \ldots, (31,36)\}$, provide additional evi-

TABLE IVb
ROMER-BASED MEASURES OF POLICY SHOCKS, SEVEN-VARIABLE SYSTEM, NOMINAL EXCHANGE RATES

	Dynamic response functions				
	Japan	Germany	Italy	France	United Kingdom
(1) 1–6 months	0.318	1.167	−0.001	−1.299	−0.817
Std. error	3.094	2.553	2.625	2.725	2.471
Significance	0.541	0.676	0.500	0.317	0.370
(2) 7–12 months	−3.016	−0.484	−3.815	−3.994	0.285
Std. error	5.589	4.856	4.804	5.426	4.699
Significance	0.295	0.460	0.214	0.231	0.524
(3) 13–18 months	−4.356	−0.415	−3.307	−1.927	2.067
Std. error	5.281	5.164	4.457	5.518	4.662
Significance	0.205	0.532	0.229	0.363	0.671
(4) 19–24 months	−5.083	−0.593	−3.603	−2.172	1.214
Std. error	5.127	4.870	4.204	4.993	4.262
Significance	0.161	0.452	0.196	0.332	0.612
(5) 25–30 months	−5.319	−1.191	−4.080	−3.027	0.316
Std. error	5.213	4.879	4.289	4.898	4.166
Significance	0.154	0.404	0.171	0.268	0.530
(6) 31–36 months	−6.000	−4.048	−6.209	−6.094	−2.761
Std. error	5.705	5.213	5.134	5.365	3.700
Significance	0.146	0.219	0.113	0.128	0.228
(7) Max impact	−8.908	−7.740	−9.346	−9.944	−5.903
Std. error	5.784	5.226	5.705	6.050	3.355
Significance	0.062	0.069	0.051	0.050	0.039
(8) Max month	24.716	32.944	32.480	32.210	24.910
Std. error	14.721	17.618	16.717	17.710	18.906
Significance	0.035	0.031	0.026	0.034	0.094

dence on this point. Notice that we cannot reject, at conventional significance levels, the hypothesis that $\mu_{For,R}(i,j)$ and $\mu_{For}(i,j)$ are equal to zero. Still, even with this method of measuring policy shocks, we can reject, at the 7 percent and 8 percent significance levels, the hypothesis that the maximal impact on s_t^{For} and s_{Rt}^{For} is equal to zero (see row (7) of Tables IVa and IVb, respectively). In addition, with the exception of the United Kingdom, there is strong evidence that the maximal effect of a policy shock on real and nominal exchange rates does not occur in the initial period of the shock (see row (8) of Tables IVa and IVb, respectively). Perhaps

TABLE V
DEVIATIONS FROM UNCOVERED INTEREST PARITY FOLLOWING U. S. MONETARY POLICY SHOCKS*

	Japan	Germany	Italy	France	United Kingdom
Panel A: Benchmark specification					
1–6 months	−22.276	−28.136	−27.235	−8.678	−39.906
Std. error	8.799	6.991	9.496	7.703	10.132
Significance	0.011	0.000	0.004	0.260	0.000
7–12 months	2.009	−3.688	4.731	11.917	−13.217
Std. error	12.844	8.164	12.223	8.909	12.902
Significance	0.876	0.651	0.699	0.181	0.306
Panel B: *NBRX* shocks, 7-variable system					
1–6 months	−24.279	−16.306	−18.525	−10.067	−29.077
Std. error	7.715	5.865	6.800	6.826	8.741
Significance	0.002	0.005	0.006	0.140	0.001
7–12 months	2.683	−2.126	5.298	12.139	−4.175
Std. error	12.643	7.698	11.478	10.331	13.311
Significance	0.832	0.782	0.644	0.240	0.754
Panel C: Federal funds rate shocks, 7-variable system					
1–6 months	−48.759	−37.817	−42.211	−41.300	−35.850
Std. error	7.977	6.784	7.140	7.096	9.112
Significance	0.000	0.000	0.000	0.000	0.000
7–12 months	−20.197	−8.101	1.284	5.363	−5.015
Std. error	9.916	7.805	10.385	9.409	13.238
Significance	0.042	0.299	0.902	0.569	0.705
Panel D: Romer and Romer shock, 8-variable system					
1–6 months	−203.410	−147.222	−114.344	−181.952	−96.945
Std. error	86.487	62.900	82.061	73.026	95.418
Significance	0.019	0.019	0.164	0.013	0.310
7–12 months	−137.337	−8.795	−55.551	−91.294	53.952
Std. error	122.721	80.948	117.199	96.603	139.979
Significance	0.263	0.913	0.636	0.345	0.700

*This table reports the expected excess returns (in annualized basis points) that can be earned from investing in foreign one-month bills relative to U. S. one-month bills in the periods following a U. S. monetary policy shock. The point estimates refer to the average response over six-month horizons.

most surprisingly, Table V indicates that at least for Japan, Germany, and France, we can reject, at the 2 percent significance level, the hypothesis that the average value of $E_t\Psi_t^{For}$ in the first half years after the onset of a Romer and Romer episode equals zero. So once again, there is sharp evidence against uncovered interest rate parity and in favor of the view that a contractionary shock to U. S. monetary policy generates negative excess returns associated with holding foreign short-term interest-bearing assets.

IV. Relating the Unconditional and Conditional Forward Premium Biases

We have found that contractionary shocks to U. S. monetary policy are followed by sharp, persistent decreases in the spread between various foreign and U. S. interest rates, and sustained, persistent appreciations in the U. S. exchange rate. These findings are related to a classic result in the exchange rate literature: in regressions of the form,

$$\Delta s_{t+1}^{For} = \alpha + \beta(f_{t+1} - s_t^{For}) + \epsilon_{t+1}, \tag{6}$$

the coefficient β is typically estimated to be negative, rather than unity as would be the case under risk neutrality and rational expectations. Here, f_{t+1} is the logarithm of the time t dollar price of a unit of foreign currency to be delivered at time $t + 1$. The finding that the change in the future exchange rate is negatively related to the forward premium is often referred to as the "forward premium bias."[14] Under covered interest rate parity, (6) is equivalent to the regression,

$$\Delta s_{t+1}^{For} = \alpha + \beta(R_t^{US} - R_t^{For}) + \epsilon_{t+1}.$$

With $\beta < 0$, the more the U. S. interest rate exceeds the foreign interest rate, the more the dollar tends to appreciate over the holding period. So rather than offset the differential gains associated with investing in the United States, the expected appreciation of the U. S. exchange rate magnifies those returns.

The estimated impulse response functions of time t excess returns $E_t\Psi_t^{For}$ discussed in the previous section can be viewed as reflecting a "conditional forward premium bias." In particular, we found that a very specific shock to the system—a contractionary shock to U. S. monetary policy—leads to a fall in $R_t^{For} - R_t^{US}$ and a persistent appreciation in the dollar that magnifies, rather than dampens the expected returns associated with investing in the United States. So our results are complementary to those in the

14. See Hodrick [1987], Engel [1995], and Frankel and Rose [1994] for detailed reviews of the empirical evidence on the forward premium bias.

literature and shed light on a specific shock to agents' environments that helps generate the "unconditional forward premium bias."

The literature contains a variety of competing explanations for the unconditional forward premium bias. These may be useful in thinking about the delayed response of exchange rates to monetary policy shocks. Engel [1995] provides a critical review of attempts to account for these puzzles by modeling risk aversion on the part of market participants. Included in this work are tests of the CAPM, tests of latent variable models, portfolio-balance models of risk premiums and general equilibrium models of risk premiums. Frankel and Rose [1994] survey recent work on exchange rates that departs from the assumption of rational expectations. Included here is work that allows for groups of agents whose irrational expectations lead to speculative bubbles via bandwagon effects. A closely related literature uses survey data on exchange rate expectations to shed light on the hypothesis of rational expectations. See Froot and Frankel [1989], Takagi [1991], and Frankel and Rose [1994]. Finally, various authors have pursued the possibility that the puzzles discussed above represent small sample phenomena. These might arise because of peso problems or learning about regime shifts. Lewis [1994] provides a survey of work in this area.

Olivier J. Blanchard has pointed out to us that a particular type of small sample problem might be able to rationalize the delayed response of the exchange rate that we documented (see also Gourinchas and Tornell [1995] for closely related work). Suppose that there are two types of shocks to U. S. monetary policy. These induce persistent and transitory shocks, R_t^P and R_t^T, respectively, to the difference between foreign and U. S. interest rates. A decrease in R_t^P or R_t^T corresponds to a contractionary U. S. monetary policy shock. Agents see only current and lagged realizations of $R_t^{For} - R^{US}$, not the separate realizations of R_t^P and R_t^T. In this environment our identification scheme is misspecified and will isolate some combination of R_t^P and R_t^T.

Uncovered interest parity (relationship (4) for $j = 0$) implies that the time t exchange rate depends on current and all expected future values of $R_t^{For} - R_t^{US}$. But the expected value of the future interest rate spread depends on agents' view of current and past realizations of R_t^P and R_t^T. Now consider the response of the exchange rate to a negative realization to R_t^P. In the impact period of the shock agents do not know whether the shock to $R_t^{For} - R_t^{US}$

reflects a realization of R_t^p or R_t^T. Over time, they will place increasing weight on the possibility that the time t shock was to R_t^p. The dollar continues to appreciate as more weight is placed on this possibility. Since the shock to R_t^p is persistent but not permanent, the exchange rate will eventually return to its pre-shock level. So as time evolves, the response of the exchange rate will be hump shaped. Could this account for the shape of our estimated impulse response functions? Not in and of itself. This is because here there are two types of interest rate shocks. As time evolves following a shock to R_t^T, we would observe simple Dornbusch type overshooting. In this example, where both policy shocks are operative, our policy reaction function is misspecified, and our estimated impulse response function represents some combination of the separate response to R_t^T and R_t^P. We have produced examples in which this specification error leads to hump-shaped impulse response functions. However, these examples rely critically on the assumption that the sample over which the VAR is estimated is marked by an *unusually* large proportion of shocks to R_t^P, relative to the population moments. So this explanation relies on small sample arguments and specification error. Formally pursuing this conjecture empirically is an interesting avenue of research.

V. CONCLUSION

This paper investigated the effects of shocks to monetary policy on nominal and real U. S. exchange rates. We did so using alternative measures of shocks to U. S. monetary policy. We found strong evidence that contractionary policy shocks lead to (i) significant, persistent appreciations in exchange rates, both nominal and real, and (ii) significant, persistent departures from uncovered interest rate parity. The negative interest rate differentials between foreign and U. S. assets are associated with appreciations of the U. S. dollar, rather than the depreciations implied by uncovered interest rate parity. This finding is consistent with the well-documented puzzle that future changes in exchange rates are negatively related to the forward premium.

We conclude by noting that according to our results, shocks to U. S. monetary policy contributed significantly to the overall variability of U. S. exchange rates in the post-Bretton Woods era. In conjunction with our other findings, this highlights important shortcomings of monetized international Real Business Cycle models. To be fair though, monetary shocks do not explain the majority of movements in U. S. exchange rates. So monetary policy was important, but it was by no means the sole determinant of changes in real exchange rates. Our results are entirely consistent with the notion that real changes which affect the relative prices of the different goods produced by different countries could have been at least as important as monetary policy in the process of exchange determination. Providing direct evidence on this possibility is an important task that we leave for future research.

APPENDIX

This appendix describes the data used in this study.

Nominal exchange rates:

The data are bilateral monthly average exchange rates between the U. S. dollar and Japanese Yen, German Deutschemark, French Franc, Italian Lira, and U. K. Pound. For the flexible exchange rate period, the data source is the Federal Reserve Board database.

U. S. data:

The source for the following data is the Federal Reserve database: Industrial Production Index, Consumer Price Index–Urban, Federal Funds rate, monthly average of daily rates, three-month Treasury bill rates, monthly average of daily rates, Total Reserves, Nonborrowed Reserves with Extended Credit, and Special Borrowings.

Foreign data:

For each country (Japan, Germany, Italy, France, and the United Kingdom), the data source is the *International Financial Statistics* database. Industrial Production (line 66) and Consumer Price Indices (line 64) are used to measure foreign output and foreign price levels. The choice of foreign interest rate depended upon availability over the sample period.

Japan:

Short-term money market rate.

Germany:

Short-term money market rate.

France:

Short-term money market rate.

Italy:

Short-term money market rate.

United Kingdom:

Short-term Treasury bill rate.

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