Exchange Rates and Uncovered Interest Differentials: The Role of Permanent Monetary Shocks*

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Abstract

We estimate an empirical model of exchange rates with transitory and permanent monetary shocks. Using monthly post-Bretton-Woods data from the United States, the United Kingdom, and Japan, we report four main findings: First, a transitory increase in the nominal interest rate appreciates the currency, whereas a permanent increase depreciates it. Second, transitory increases in the interest rate cause short-run deviations from uncovered interest-rate parity in favor of domestic assets, whereas permanent increases cause deviations against domestic assets. Third, permanent monetary shocks explain the majority of short-run movements in the level of the nominal exchange rate. Fourth, there is no exchange rate overshooting in response to either temporary or permanent monetary shocks.

JEL Classification: E52.
Keywords: Exchange rates, permanent monetary shocks, neo Fisher effect, uncovered interest rate parity, overshooting.

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1 Introduction

How do monetary disturbances affect exchange rates and cross-country return differentials? A defining characteristic of the vast existing literature devoted to elucidating this question is the assumption that monetary shocks come in only one type. The contribution of the present paper is to distinguish between temporary and permanent monetary shocks. Doing so reveals that the effects of these two types of shock on exchange rates and cross-country asset return differentials are diametrically opposed.

The empirical model estimated in the present study is an extension of the closed economy model with permanent and transitory monetary shocks developed by Uribe (2018). Permanent monetary disturbances are identified by assuming that the domestic nonstationary monetary shock is cointegrated with domestic inflation and the domestic interest rate. Domestic transitory monetary policy shocks are identified by assuming that transitory increases in the policy rate cause a nonpositive impact effect on domestic inflation and output. In addition to domestic transitory and permanent monetary shocks, the present formulation introduces a foreign permanent monetary shock with the property of being cointegrated with the foreign nominal interest rate. Finally, the empirical model assumes that the depreciation rate is cointegrated with the cross-country inflation differential, which implies that the real depreciation rate is stationary. The model is estimated on monthly data covering the period 1974:1 to 2018:3 using Bayesian methods. The domestic economy is taken to be the United States and the foreign economy either the United Kingdom or Japan.

The paper finds that transitory increases in the nominal interest rate cause a persistent appreciation of the domestic currency, which is in line with the results of earlier studies. However, permanent increases in the nominal interest rate are found to cause a depreciation of the exchange rate already in the short run. It follows that it is not the case that in general a monetary policy shock whereby the policy increases leads to an appreciation of the domestic currency. The paper further documents that for both types of monetary shocks the short-run effect of the monetary policy shock on the exchange rate is weaker than at any point in the future, implying the absence of an overshooting effect.

The distinction between transitory and permanent monetary shocks also has important consequences for the dynamics of interest rate differentials. As in the related empirical literature, a transitory increase in the nominal interest rate causes a short-run departure from uncovered interest rate parity in favor of domestic assets. By contrast, a permanent increase in the nominal interest rate causes a departure from uncovered interest rate parity against domestic assets. This result suggests that carry-trade speculation conditional on monetary shocks might involve more risk than is implied by estimates that do not distinguish permanent from transitory shocks.

The effects of permanent monetary shocks would be of little economic importance if such disturbances were found to account for a small fraction of exchange rate movements. But we find that this is not the case. The estimated empirical model predicts that domestic and foreign permanent monetary shocks jointly explain the majority of the forecast error variance of the nominal exchange rate at horizons 12 to 48 months. On the other hand, transitory monetary shocks are found to play an insignificant role. Finally, we find that real exchange rates are driven primarily by real permanent disturbances with a negligible contribution of monetary shocks.

A large number of empirical studies, starting with the seminal contribution by Eichenbaum and Evans (1995), have characterized the response of the nominal exchange rate and cross-country return differentials to identified monetary policy shocks. These studies use a variety of identification schemes of monetary policy shocks including recursive identification, structural vector autoregression models, sign restrictions, and high frequency identification. The main conclusions that have emerged from this body of work are: First, the domestic currency appreciates in response to a tightening in domestic monetary policy. Second, the empirical evidence is consistent with the Dornbusch exchange-rate overshooting result, namely, that the exchange rate appreciates more in the short run than in the long run. Some papers find that overshooting occurs on impact, as suggested by Dornbusch's model (Kim and Roubini, 2000; Faust and Rogers, 2003; Bjørnland, 2009; Kim, Moon, and Velasco, 2017), while others find that it takes place in a delayed fashion (Eichenbaum and Evans, 1995; Scholl and Uhlig, 2008). Third, much of the available econometric evidence suggests that a contractionary monetary shock causes a persistent deviation from uncovered interest rate parity in favor of domestic interest-bearing assets.[1]

There also exist papers that study the behavior of exchange rates in response to monetary policy shocks identified using the Romer and Romer narrative approach (see Eichenbaum and Evans (1995, especially figure IV) and Hettig, Müller, and Wolf (2019)). These studies also find that a monetary tightening leads to an appreciation of the exchange rate. However, these studies show that there is no exchange rate overshooting, that is, the exchange rate does not appreciate more on impact than in the long run. The results reported in the present paper for the response of the exchange rate to temporary tightenings is consistent with these findings.

The present paper departs from the literature surveyed in this and the previous paragraphs in that it distinguishes between temporary and permanent monetary disturbances. A related paper that also considers the effects of permanent monetary disturbances on the exchange rate is De Michelis and Iacoviello (2016). These authors find that in response to an increase in the U.S. inflation target, the U.S. real exchange temporarily depreciates. Unlike the present study, the focus of this paper is not on the effects of monetary policy shocks on nominal and exchange rates or on uncovered interest rate differentials. In addition, De Michelis and Iacoviello do not jointly estimate the effects of temporary and permanent monetary disturbances. To the best of our knowledge the present study represents the first attempt to implement this distinction in the context of an international empirical model and to document that it has significant consequences for the dynamics of exchange rates and excess returns.

[1]Exceptions in this regard are the findings of Kim and Roubini (2000) and Kim, Moon, and Velasco (2017). The latter authors argue that deviations from uncovered interest rate parity were observed during the Volcker period but not in the post-Volcker period.

More broadly, the present paper is also related to Engel (2016), who estimates a vector error correction system in the nominal exchange rate, the cross country price-level differential, and the nominal interest-rate differential and extracts the permanent component of the nominal exchange rate. An important difference between our analysis and that of Engel is that his analysis focuses on the unconditional correlation of foreign-exchange risk premia and real interest-rate differentials, whereas the present study is concerned with the effects of identified temporary and permanent monetary shocks on the level of exchange rates and uncovered interest differentials. In addition, the error-correction model of Engel's imposes stationarity in the level of the real exchange rate and the nominal interest rate differential, whereas the formulation in the present paper allows for the possibility that both of these variables have a permanent component.

The remainder of the paper is laid out as follows. Section 2 presents the empirical model and the identification scheme. Section 3 explains the estimation procedure. Section 4 describes the data used for estimation and their sources. Section 5 presents the main finding of the paper, namely, the responses of exchange rates and interest rate differentials to permanent and transitory monetary shocks. Section 6 documents the importance of permanent monetary shocks as drivers of nominal exchange rates. Section 7 estimates a variant of the model in which monetary policy in the United States and the United Kingdom are assumed to share a common permanent component. Section 8 shows that the main results of the paper are robust to restricting the sample to the Volcker era. Section 9 concludes.

2 The Empirical Model

The empirical model adapts the closed-economy model of temporary and permanent monetary shocks developed in Uribe (2018) to allow for a foreign block. The model is cast in five variables: the logarithm of real domestic output, denoted y_t, domestic inflation, denoted π_t, the domestic nominal interest rate, denoted i_t, the foreign interest rate, denoted i^*_t, and the depreciation rate of the domestic currency, denoted $\epsilon_t \equiv \ln(S_t/S_{t-1})$, where S_t denotes the spot nominal exchange rate, defined as the price of one unit of foreign currency in terms of units of domestic currency in period t. The variables π_t, i_t, i_t^*, and ϵ_t are expressed in percent per year. The domestic economy is meant to be the United States, and the foreign economy is meant to be either the United Kingdom or Japan.

All five variables are assumed to be nonstationary. We assume that output is cointegrated with an exogenous variable X_t, which can be interpreted as a stochastic output trend. In addition, we assume that the domestic and foreign nominal interest rates are cointegrated, respectively, with exogenous variables X_t^m and X_t^{m*}, which we interpret as permanent domestic and foreign monetary shocks.[2] Finally, we assume that the variable $\epsilon_t + \pi_t^* - \pi_t$, which captures the change in the bilateral real exchange rate, is stationary. We then define the following vector of stationary variables

$$\begin{bmatrix} \hat{y}_t \\ \hat{\pi}_t \\ \hat{i}_t \\ \hat{\epsilon}_t \\ \widehat{i_t^*} \end{bmatrix} \equiv \begin{bmatrix} y_t - X_t \\ \pi_t - X_t^m \\ i_t - X_t^m \\ \epsilon_t - X_t^m + X_t^{m^*} \\ i_t^* - X_t^{m^*} \end{bmatrix}$$

and assume that it evolves according to the following autoregressive process:[3]

$$\begin{bmatrix} \hat{y}_t \\ \hat{\pi}_t \\ \hat{i}_t \\ \hat{\epsilon}_t \\ \widehat{i_t^*} \end{bmatrix} = \sum_{i=1}^{L} B_i \begin{bmatrix} \hat{y}_{t-i} \\ \hat{\pi}_{t-i} \\ \hat{i}_{t-i} \\ \hat{\epsilon}_{t-i} \\ \widehat{i_{t-i}^*} \end{bmatrix} + C \begin{bmatrix} \Delta X_t^m \\ z_t^m \\ \Delta X_t \\ z_t \\ \Delta X_t^{m*} \end{bmatrix}, \tag{1}$$

[2]In section 7, we allow for the possibility that domestic monetary policy is cointegrated with foreign monetary policy. In that specification, the nominal devaluation rate is assumed to be stationary.

[3]The exposition of the model omits intercepts.

where the variable z_t^m is a stationary domestic monetary shock, and the variable z_t is interpreted as a combination of stationary shocks affecting the variables in the system, which we do not wish to identify individually. The objects B_i for $i = 1, \ldots, L$ and C are 5×5 matrices of coefficients, and L denotes the number of lags. We assume that the five driving forces follow univariate AR(1) processes of the form

$$\begin{bmatrix} \Delta X_{t+1}^m \\ z_{t+1}^m \\ \Delta X_{t+1} \\ z_{t+1} \\ \Delta X_{t+1}^{m*} \end{bmatrix} = \rho \begin{bmatrix} \Delta X_t^m \\ z_t^m \\ \Delta X_t \\ z_t \\ \Delta X_t^{m*} \end{bmatrix} + \psi \begin{bmatrix} \nu_{t+1}^1 \\ \nu_{t+1}^2 \\ \nu_{t+1}^3 \\ \nu_{t+1}^4 \\ \nu_{t+1}^5 \end{bmatrix}, \tag{2}$$

where $\nu_t^i \sim$ i.i.d. $\mathcal{N}(0,1)$ for $i = 1, \ldots, 5$ and ρ and ψ are 5×5 diagonal matrices of coefficients.

In sum, the vector of stationary endogenous variables $\hat{y}_t$, $\hat{\pi}_t$, $\hat{i}_t$, $\hat{\epsilon}_t$, and $\hat{i}_t^*$ follows an autoregressive process driven by innovations in the growth rate of the domestic output trend, ΔX_t, innovations in changes in the permanent components of domestic and foreign monetary policy, ΔX_t^m and ΔX_t^{m*}, stationary domestic monetary shocks, z_t^m, and other stationary shocks, z_t.

The system (1)-(2) is unobservable, because neither the detrended endogenous variables nor the driving forces are observed. Thus, we think of that system as describing the evolution of state variables in a state-space representation. To be able to estimate the model, we add equations linking the unobservable variables to variables with an empirical counterpart. Specifically, we include as observable variables the growth rate of real output, Δy_t, the domestic interest-rate inflation differential, $r_t \equiv i_t - \pi_t$, the changes in the domestic and foreign nominal interest rates, Δi_t and Δi_t^*, and the change in the nominal depreciation rate, $\Delta\epsilon_t$. These variables are linked to the unobservable states as follows:

$$\Delta y_t = \hat{y}_t - \hat{y}_{t-1} + \Delta X_t$$

$$r_t = \hat{i}_t - \hat{\pi}_t$$

$$\Delta i_t = \hat{i}_t - \hat{i}_{t-1} + \Delta X_t^m$$

$$\Delta\epsilon_t = \hat{\epsilon}_t - \hat{\epsilon}_{t-1} + \Delta X_t^m - \Delta X_t^{m*}$$

and

$$\Delta i_t^* = \hat{i}_t^* - \hat{i}_{t-1}^* + \Delta X_t^{m*}.$$

We assume that the variables on the left-hand sides of these expressions are observed with measurement error. Specifically, we assume that the econometrician observes the vector o_t defined as

$$o_t = \begin{bmatrix} \Delta y_t \\ r_t \\ \Delta i_t \\ \Delta\epsilon_t \\ \Delta i_t^* \end{bmatrix} + \mu_t, \tag{3}$$

where μ_t is a 5×1 vector of measurement errors distributed i.i.d. $\mathcal{N}(\emptyset, R)$, and R is a 5×5 diagonal variance-covariance matrix.

2.1 Identification

We have already introduced model restrictions that allow for the identification of the permanent sources of uncertainty. Specifically, the assumptions that y_t is cointegrated with X_t, that i_t, and π_t are cointegrated with X_t^m, that i_t^* is cointegrated with X_t^{m*}, and that ϵ_t is cointegrated with $X_t^m - X_t^{m*}$, allow for the identification of the three permanent shocks. In addition, we introduce restrictions that allow us to identify the transitory domestic monetary shock. In the spirit of the sign restriction approach pioneered by Uhlig (2005), we assume that increases in the domestic nominal interest rate caused by temporary innovations in domestic monetary policy have nonpositive impact effects on inflation or output. Formally, we impose the following sign restrictions on the coefficients of the matrix C in equation (1):

$$C_{12} \leq 0 \quad \text{and} \quad C_{22} \leq 0.$$

These assumptions are less restrictive than the ones often imposed in models in which the monetary shock is identified recursively. For example, Eichenbaum and Evans (1995) and Christiano, Eichenbaum, and Evans (2005) impose the restriction that output and inflation cannot respond contemporaneously to a monetary policy shock, which in the context of the present formulation is equivalent to imposing $C_{12} = C_{22} = 0$.

Without loss of generality, we normalize the impact effect of a unit innovation in the transitory monetary shock, z_t^m, on the domestic nominal interest rate to unity. Similarly, we normalize the impact effect of a unit innovation in the transitory shock, z_t, on domestic output to unity. Thus, we set

$$C_{32} = C_{14} = 1.$$

In line with the discussion in Faust and Rogers (2003), we leave unconstrained the contemporaneous response of the foreign interest rate to U.S. monetary policy, C_{51} and C_{52}, thus allowing the foreign monetary authority to respond within the month to U.S. monetary policy shocks. Similarly, we leave unconstrained the contemporaneous response of the U.S. interest rate to foreign monetary shocks, C_{35}.

A related issue has to do with the identifiability of the parameters of the model. We check for identifiability by applying the test proposed by Iskrev (2010). In essence the Iskrev test checks whether the derivatives of the predicted autocovariogram of the observables with respect to the vector of estimated parameters has rank equal to the length of the vector of estimated parameters. After estimating the model using the procedure and data described in sections 3 and 4, we find that, regardless of whether the foreign country is the United Kingdom or Japan, the derivative of the vectorized predicted autocovariogram of the vector of observables with respect to the parameters has full column rank when evaluated at the posterior mean of the Bayesian estimate. Full column rank obtains starting with the inclusion of covariances of order 0 to 8. According to this test, therefore, in both cases the parameter vector is identifiable in the neighborhood of our estimate. Specifically, the test result indicates that in the neighborhood of the estimate all values of the vector of parameters different from the estimated one give rise to autocovariograms that are different from the one associated with the posterior mean estimate.

3 Estimation

For the purpose of estimating the model, it is convenient to express it in a first-order state-space form. To this end let

$$
\begin{aligned}
\hat{Y}_t &\equiv \begin{bmatrix} \hat{y}_t & \hat{\pi}_t & \hat{i}_t & \hat{\epsilon}_t & \widehat{i_t^*} \end{bmatrix}', \\
u_t &\equiv \begin{bmatrix} \Delta X_t^m & z_t^m & \Delta X_t & z_t & \Delta X_t^{m*} \end{bmatrix}', \\
\nu_t &\equiv \begin{bmatrix} \nu_t^1 & \nu_t^2 & \nu_t^3 & \nu_t^4 & \nu_t^5 \end{bmatrix}',
\end{aligned}
$$

and

$$
\xi_t \equiv \begin{bmatrix} \hat{Y}_t' & \hat{Y}_{t-1}' & \ldots & \hat{Y}_{t-L+1}' & u_t' \end{bmatrix}'.
$$

Then we can write the system (1)-(3) as

$$
\xi_{t+1} = F\xi_t + P\nu_{t+1}, \tag{4}
$$

Table 1: Prior Distributions

Parameter	Distribution	Mean.	Std. Dev.
Main diagonal elements of B_1	Normal	0.95	0.5
All other elements of B_i, $i = 1, \ldots, L$	Normal	0	0.25
C_{21}, C_{31}, C_{55}	Normal	-1	1
$-C_{12}, -C_{22}$	Gamma	1	1
All other estimated elements of C	Normal	0	1
ψ_{ii}, $i = 1, \ldots, 5$	Gamma	1	1
ρ_{ii}, $i = 1, 2, 3, 5$	Beta	0.3	0.2
ρ_{44}	Beta	0.7	0.2
R_{ii}, $i = 1, \ldots, 5$	Uniform $\left[0, \frac{\mathrm{var}(o_t)}{10}\right]$	$\frac{\mathrm{var}(o_t)}{10\times 2}$	$\frac{\mathrm{var}(o_t)}{10\times\sqrt{12}}$
Elements of A	Normal	mean(o_t)	$\sqrt{\frac{\mathrm{var}(o_t)}{T}}$

Notes. T denotes the sample length, which equals 513 months. The vector A denotes the mean of the vector o_t, and is defined in appendix A.

and

$$o_t = H'\xi_t + \mu_t, \tag{5}$$

where the matrices F, P, and H are known functions of the matrices B_i for $i = 1, \ldots, L$, C, ρ, and ψ and are shown in appendix A. This representation is of particular use for calculating the likelihood of the data $\{o_1\, o_2\, \ldots o_T\}$, where T is the number of observations, via the Kalman filter.

We estimate the model at a monthly frequency using Bayesian techniques. The specification includes 6 lags in equation (1), $L = 6$. In assigning priors to the estimated parameters, we follow Uribe (2018). Table 1 describes the assumed prior distributions of the estimated parameters. We impose normal prior distributions to all elements of B_i, for $i = 1, \ldots, L$. We assume a prior parameterization in which at the mean of the prior parameter distribution the elements of $\hat{Y}_t$ follow univariate autoregressive processes. So when evaluated at their prior mean, only the main diagonal of B_1 takes nonzero values and all other elements of B_i for $i = 1, \ldots, L$ are nil. We impose an autoregressive coefficient of 0.95 in all equations, so that all elements along the main diagonal of B_1 take a prior mean of 0.95. We assign a prior standard deviation of 0.5 to the diagonal elements of B_1, which implies a coefficient of variation close to one half (0.5/0.95). We impose lower prior standard deviations on all other elements of the matrices B_i for $i = 1, \ldots, L$, and set them to 0.25.

All elements of the matrix C are assumed to have normal prior distributions with mean zero and unit standard deviation, with the following exceptions: First, as mentioned earlier, C_{32} and C_{14} are normalized to one, and are therefore not estimated. Second, elements C_{21} and C_{31}, which govern the responses of $\hat{\pi}_t \equiv \pi_t - X_t^m$ and $\hat{i}_t \equiv i_t - X_t^m$, respectively, to an innovation in ΔX_t^m, are assumed to have a prior mean of -1. This means that a shock that increases the U.S. nominal rate in the long run by 1 percentage point, under the prior, has a zero impact effect on inflation and the nominal rate. Third, we assume that C_{55} has a prior mean of -1. This implies that a shock that increases the foreign nominal interest rate in the long run by 1 percentage point has a prior mean impact effect on the foreign nominal interest rate of zero. Fourth, we assume that the prior distributions of $-C_{12}$ and $-C_{22}$ have a positive support. This guarantees, in accordance with our identification assumptions, that transitory increases in the nominal interest rate are limited to have nonpositive impact effects on inflation and output. We assume Gamma distributions with mean and standard deviation equal to 1 for both of these parameters. Thus, as in Uribe (2018), identification of the domestic transitory monetary shock is achieved via restrictions on prior distributions.

The parameters ψ_{ii}, for $i = 1, \ldots, 5$, representing the standard deviations of the five exogenous innovations in the AR(1) process (2) are all assigned Gamma prior distributions with mean and standard deviation equal to one. We impose nonnegative serial correlations on the exogenous shocks ($\rho_{ii} \in (0, 1)$ for $i = 1, \ldots, 5$), and adopt Beta prior distributions for these parameters. We assume relatively small means of 0.3 for the prior serial correlations of the three monetary shocks (z_t^m, ΔX_t^m, and ΔX_t^{m*}) and the nonmonetary nonstationary shock (ΔX_t) and a relatively high mean of 0.7 for the stationary nonmonetary shock (z_t). The prior distributions of all serial correlations are assumed to have a standard deviation of 0.2. The variances of all measurement errors, R_{ii}, are assumed to have a uniform prior distribution with lower bound 0 and upper bound of 10 percent of the sample variance of the corresponding observable indicator. Although not explicitly discussed thus far, the estimated model includes constants. These constants appear in the observation equation (5), for details see Appendix A. The unconditional means of the five observables are assumed to have normal prior distributions with means equal to their sample means and standard deviations equal to their sample standard deviations divided by the square root of the length of the sample period.

To draw from the posterior distribution of the estimated parameters, we apply the Metropolis-Hastings sampler. We construct a Monte-Carlo Markov chain (MCMC) of two million draws and burn the initial one million draws. Posterior means and error bands around the impulse responses shown in later sections are constructed from a random subsample of the MCMC chain of length 100 thousand with replacement.

4 The Data

The estimation uses monthly data from 1974:1 to 2018:3. The domestic country is meant to be the United States and the foreign country the United Kingdom or Japan. We proxy real domestic output, y_t, by the log of an index of industrial production for the United States. The source is OECD MEI database Production of Total Industry, seasonally adjusted, index 2010=100, downloaded from stats.oecd.org on June 1, 2018. Domestic inflation, π_t, is proxied by the growth rate of the U.S. consumer price index. The source is OECD MEI dataset Consumer Price Index, all items, index 2010=100, downloaded from stats.oecd.org on June 1, 2018. The domestic nominal interest rate, i_t, is proxied by the U.S. federal funds rate (monthly averages of daily rates). The source is the file frb_H15.xls available at www.federalreserve.gov, downloaded on June 2, 2018.

When the foreign economy is taken to be the United Kingdom the data used for estimation are as follows: The devaluation rate, ϵ_t, is measured by the growth rate of the dollar-pound nominal exchange rate (dollar price of one British pound). The source is fred.stlouisfed.org series EXUSUK, downloaded on June 1, 2018. The foreign nominal interest rate, i_t^*, is measured as the Official Bank Rate of the Bank of England obtained as follows. From January 1975 to March 2018, the source is www.bankofengland.co.uk, and the series name is IUMABEDR, downloaded on June 1, 2018. For the period January 1974 to December 1974, the source is fred.stlouisfed.org, and the series name is BOERUKM, retrieved on June 1, 2018. The estimation of the model that replaces the nominal exchange rate with the real exchange rate uses the bilateral dollar-pound real exchange rate, defined as $e_t \equiv P_t^* S_t / P_t$, where P_t^* denotes the UK consumer price index in period t, P_t denotes the U.S. consumer price index in period t, and S_t denotes the dollar-pound nominal exchange rate. The measure for P_t^* is Consumer Prices, All Items, index 2010=100, from the OECD Consumer Prices (MEI) dataset, downloaded from stats.oecd.org on June 1, 2018.

When the foreign country is assumed to be Japan, the foreign nominal interest rate, i_t^*, is proxied by the call rate, series IR01MADR1M. The source is www.stat-search.boj.or.jp and the data were downloaded on June 4, 2018. The dollar-yen nominal exchange has the same source and download date, and is given by the series FM08FXERM07. The foreign price level, P_t^* is proxied by the series Consumer Prices, All Items, index 2010=100, from the OECD Consumer Prices (MEI) dataset, downloaded from stats.oecd.org on June 1, 2018.

5 Results

Figure 1 displays the impulse responses to temporary and permanent U.S. monetary shocks when the foreign economy is taken to be the United Kingdom. The variables included in the figure are the nominal U.S. interest rate, i_t, the nominal exchange rate (dollar price of one British pound), $\ln S_t$, the real dollar-pound exchange rate, $\ln e_t$, and the uncovered interest rate differential, $i_t - i_t^* - \epsilon_{t+1}$. The left column of the figure displays impulse responses to a permanent U.S. monetary shock (X_t^m) that increases the federal funds rate by 1 percentage point in the long run. The right column displays impulse responses to a transitory U.S. monetary shock (z_t^m) that increases the federal funds rate by 1 percentage point on impact. The figure includes 95-percent asymetric error bands computed using the Sims-Zha (1999) method.[4]

The main message conveyed by figure 1 is that the responses of the exchange rate and the uncovered interest-rate differential are quite different depending on whether the monetary shock is permanent or transitory.

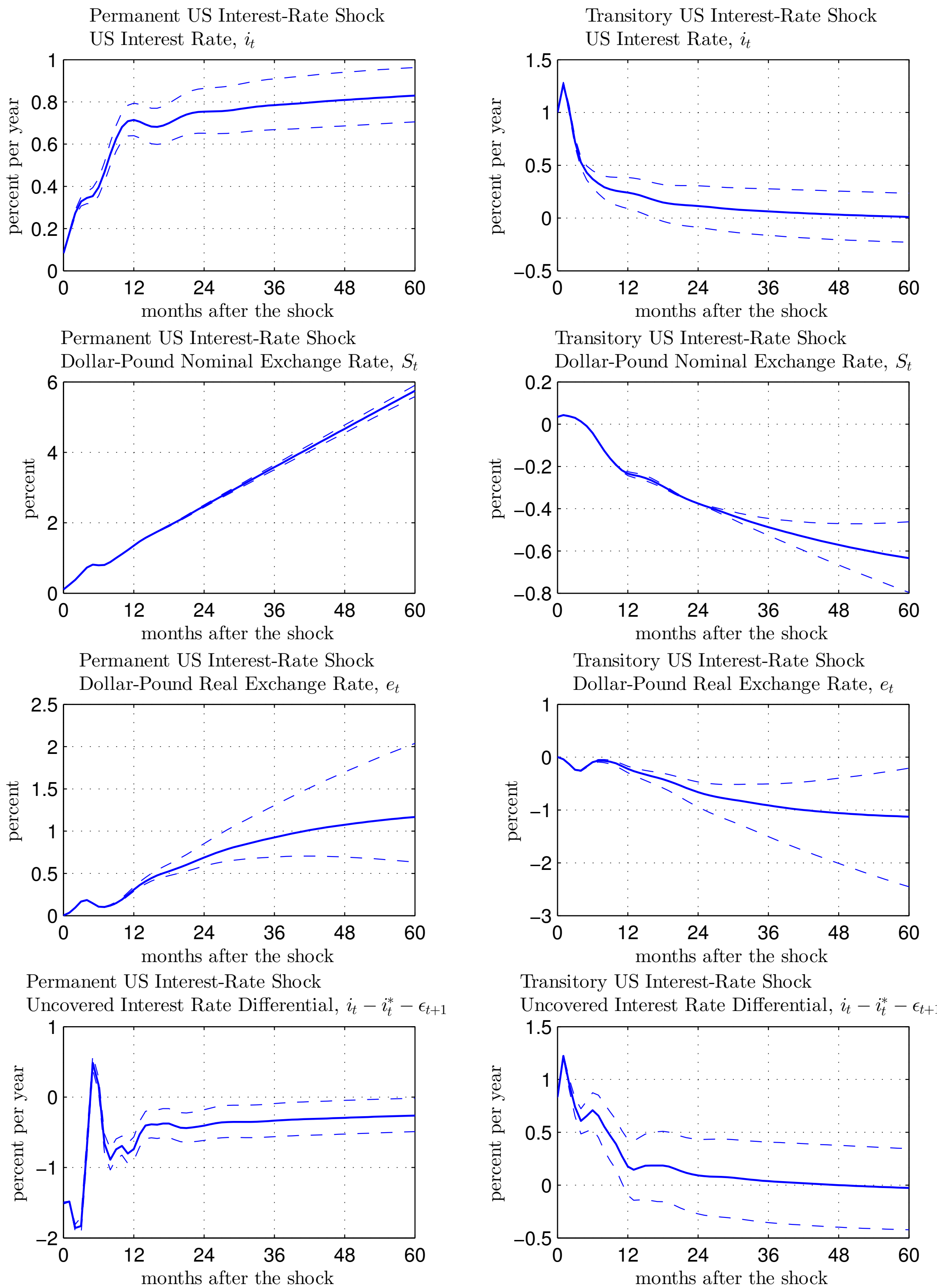
A permanent increase in the nominal interest rate produces an immediate depreciation of the nominal exchange rate, whereas a transitory increase in the nominal interest rate causes a persistent appreciation.[5] Importantly, the temporary monetary shock does not produce overshooting of the exchange rate, as its response is weaker in the short run than at any point in the future. This finding is in contrast with the existing body of empirical studies that have shown the existence of overshooting either immediately (Kim and Roubini, 2000; Faust and Rogers, 2003; Kim, Moon, and Velasco, 2017), or in a delayed fashion (Eichenbaum and Evans, 1995; Scholl and Uhlig, 2008). One possible interpretation of the difference between the findings presented here and those just cited is that the overshooting effect found in models that include a single monetary shock are the result of confounding impulses stemming from transitory and permanent sources.

The results obtained for the nominal exchange rate extend to the real exchange rate, as shown in the third row of figure 1. To compute these impulse responses we replace in equation (1) the stationary version of the nominal devaluation rate, $\widehat{\epsilon}_t$, by the real devaluation rate, $\epsilon_t^r \equiv \ln(e_t/e_{t-1})$, thus imposing, in accordance with our maintained assumption, that the real devaluation rate is stationary. The estimation is performed replacing $\Delta\epsilon_t$ by ϵ_t^r in the vector of observables, which is justified by the assumption that the latter is stationary in

[4]The error bands were computed based on the variance covariance matrix of the vector of impulse responses to a given disturbance, either X_t^m or z_t^m. The error bands are little changed if one were to base their construction on the variance covariance matrix of the impulse response of a given variable to a given shock.

[5]By construction, the nominal exchange rate depreciates without bounds in the long run in response to the permanent monetary shock. Thus, the noteworthy feature of the impulse response of $\ln S_t$ to a positive innovation in X_t^m is the predicted depreciation in the short run.

Figure 1: Impulse Responses to Permanent and Transitory U.S. Monetary Shocks: United Kingdom



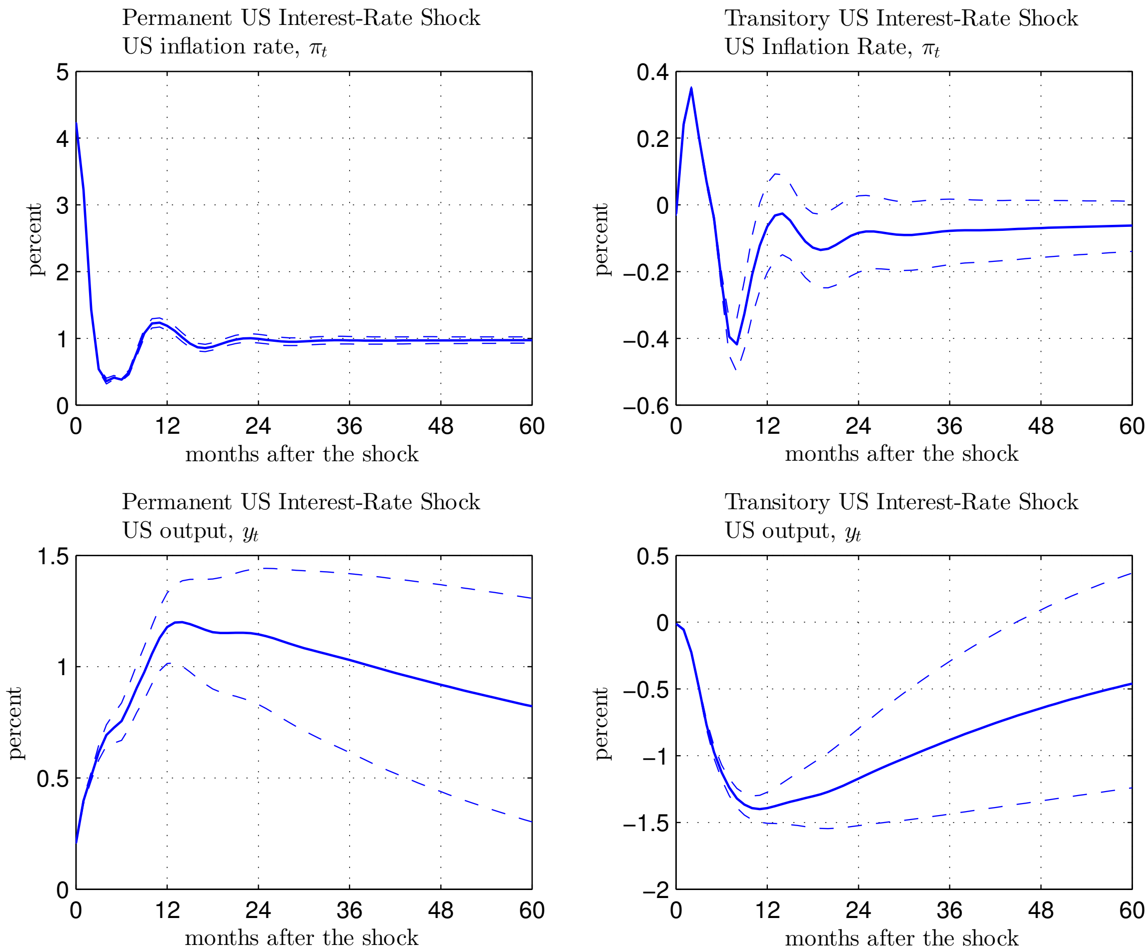
Notes. The left column displays the response to a permanent monetary shock that increases the U.S. nominal interest rate by 1 annual percentage point in the long run. The right column displays the response to a transitory monetary shock that increases the U.S. nominal interest rate by 1 annual percentage point on impact. Solid lines are posterior mean estimates. Broken lines are asymmetric 95-percent confidence bands computed using the Sims-Zha (1999) method.

levels. As in the case of nominal exchange rates, the dollar-pound real exchange rate depreciates in response to a permanent increase in the U.S. nominal interest rate, but appreciates in response to a transitory monetary tightening. Neither shock causes overshooting in the sense that the initial response of the real exchange rate is larger than its long-run response. Quantitatively the responses of the real and nominal exchange rates are similar when the monetary shock is transitory but different when the shock is permanent. Specifically, when the shock is permanent, the depreciation of the real exchange rate is much smaller than that of the nominal exchange rate even in the short run (which is the relevant horizon for this comparison). This suggests that, conditional on monetary shocks, the Mussa (1986) Puzzle—the argument that nominal and real exchange rates move in tandem in the short run—applies to transitory monetary shocks but not significantly to permanent monetary shocks.

Consider now the response of the uncovered interest-rate differential, shown in the last row of figure 1. In line with results documented extensively in the related empirical literature (see, for instance, the papers cited above), the bottom right panel of the figure shows that a temporary increase in the nominal interest rate causes a deviation from uncovered interest-rate parity (UIP) in favor of domestic assets. The novel result is that, contrary to what happens under a temporary shock, a permanent increase in the nominal interest rate causes a deviation from UIP against domestic assets. This finding suggests that deviations from uncovered interest rate parity caused by monetary shocks might carry more risk than previously thought, unless investors have the ability to tell apart permanent from transitory monetary shocks as they take place.

Figure 2 displays the responses of domestic (that is, U.S.) inflation and output to the U.S. monetary shocks, X_t^m and z_t^m. Consistent with the vast existing literature on the effects of transitory monetary policy shocks, see, for example, Christiano, Eichenbaum, and Evans (2005), we find that a temporary tightening causes a contraction in aggregate activity and a fall in inflation. In addition, in line with Uribe (2018), the left panels of figure 2 show that a

Figure 2: Impulse Responses of U.S. Inflation and Output to Permanent and Transitory U.S. Monetary Shocks: United Kingdom



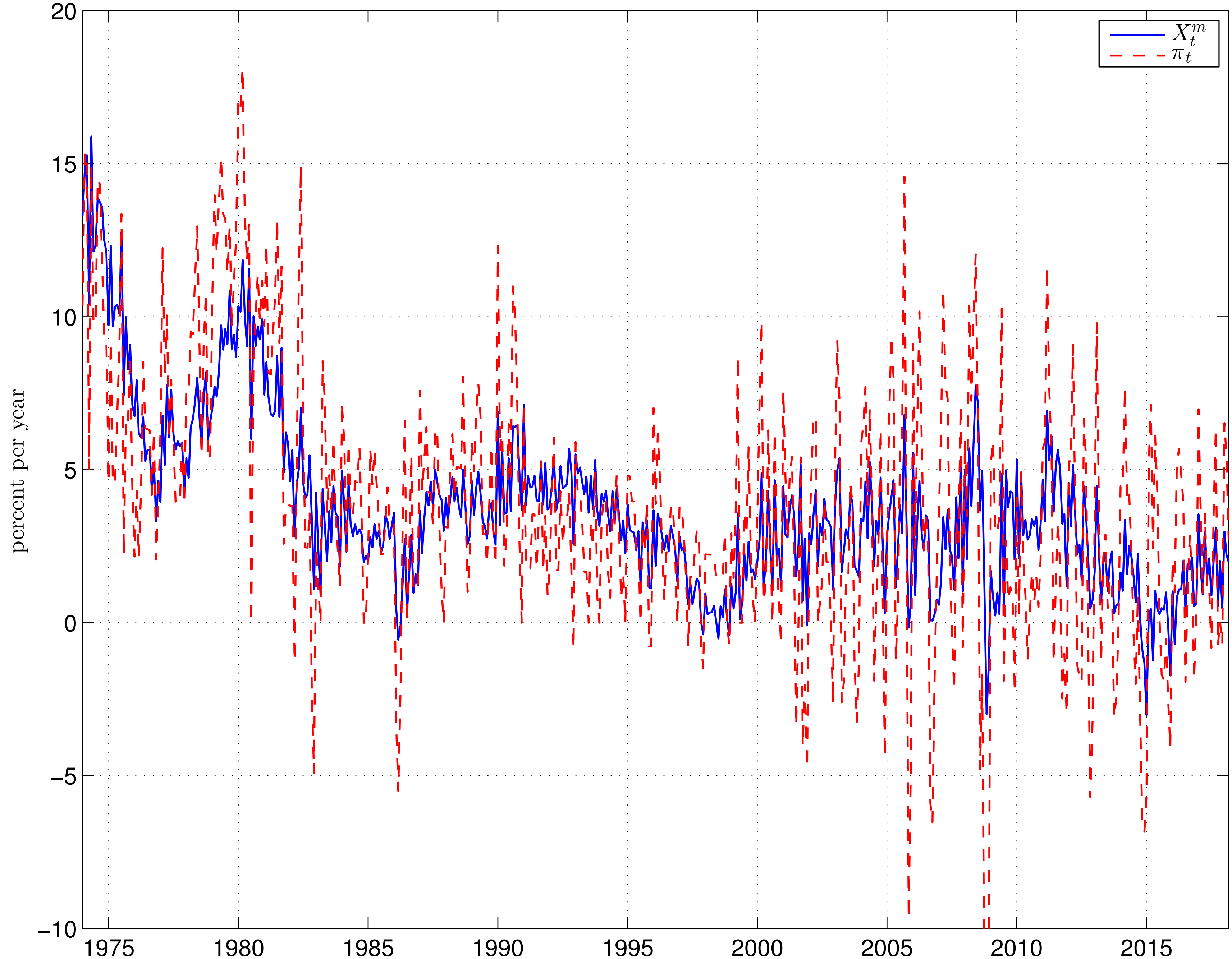
Note. See notes to figure 1.

permanent increase in the nominal interest rate causes an immediate increase in inflation and an expansion in aggregate activity. Thus, the neo-Fisher effect, whereby a monetary policy shock that raises the nominal interest rate in the long run causes an increase in inflation already in the short run, appears to be present not only at a quarterly frequency as in the work of Uribe (2018) but also at a monthly frequency.

It is of interest to ascertain the behavior of the permanent component of U.S. monetary policy as viewed through the lens of the empirical model. Figure 3 displays with a solid line the estimate of the permanent U.S. monetary shock, X_t^m, and with a broken line the actual U.S. inflation rate, π_t. A constant was added to the former to ensure that its sample mean is equal to that of π_t. Because the data is monthly the path of π_t is quite jaggy. Nonetheless it is discernible from the figure that the permanent monetary component, X_t^m, tracks well low frequency movements in inflation. In particular, X_t^m increases during the high inflation years of the late 1970s and falls gradually during the Volcker disinflation of the early 1980s. Also as reported in Uribe (2018) the estimation suggests the presence of a significant permanent component in both the low inflation following the great contraction of 2008 and the increase in inflation that started when the Fed decided to embark on a gradual normalization of interest rates in late 2015.

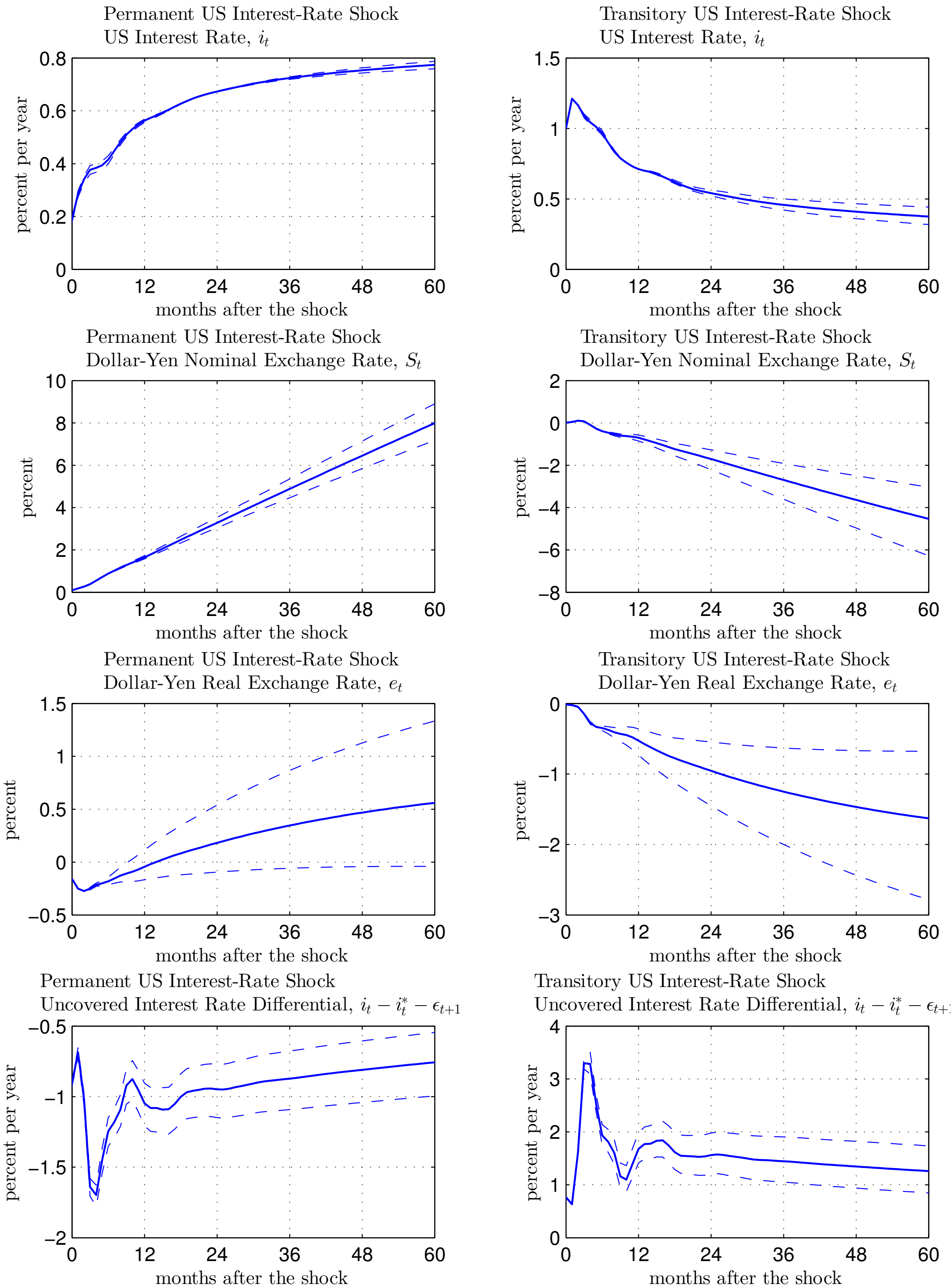
The results obtained when the foreign country is taken to be the United Kingdom continue to hold when the foreign country is assumed to be Japan. This is shown in Figure 4. In particular, a permanent increase in the U.S. policy rate causes a depreciation of the dollar vis-á-vis the Japanese yen already in the short run and a deviation from UIP in favor of Japanese assets, whereas the opposite results obtain when the monetary tightening in the United States is temporary. Importantly, neither the nominal nor the real exchange rate display overshooting. Finally, as shown in figure 5, the responses of U.S. inflation and real output are also broadly in line with the results reported when the model is estimated on U.S. and U.K. data, that is, a temporary monetary tightening has the conventional contractionary effect on real activity and prices, whereas a permanent tightening is associated with neo

Figure 3: U.S. Inflation and Its Permanent Component: Empirical model with United Kingdom



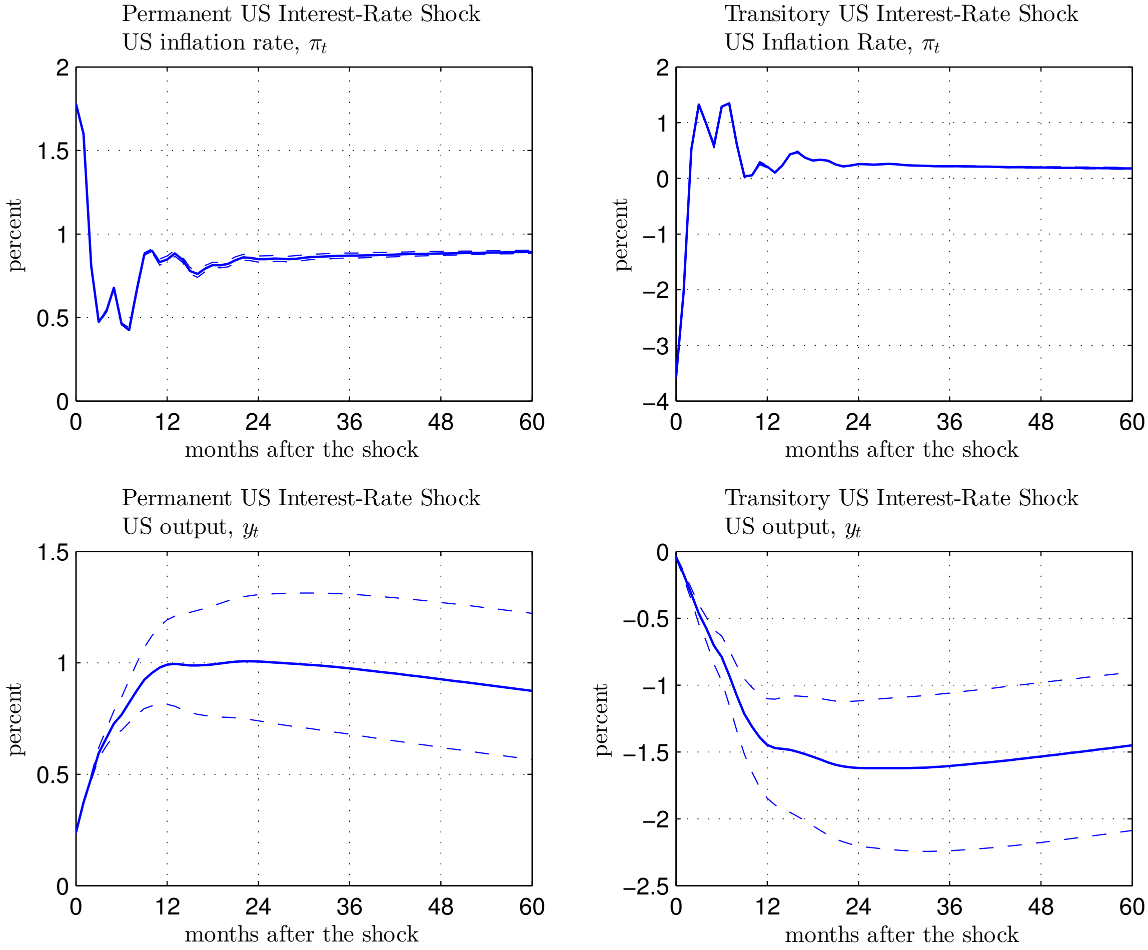
Note. The solid line shows the level of the permanent U.S. monetary shock X_t^m adjusted by a constant so that in sample it has the same mean as U.S. inflation. The dashed line shows month-over-month U.S. CPI inflation expressed in percent per year. The vertical axis is truncated below at -10 percent.

Figure 4: Impulse Responses to Permanent and Transitory U.S. Monetary Shocks: Japan



Note. See notes to figure 1.

Figure 5: Impulse Responses of U.S. Inflation and Output to Permanent and Transitory U.S. Monetary Shocks: Japan



Note. See notes to figure 1.

Table 2: Forecast Error Variance Decomposition at Horizon 36 months

	Δy_t	π_t	i_t	$\ln S_t$	$\ln e_t$	i_t^*	$i_t - i_t^* - \epsilon_{t+1}$
	A. United Kingdom						
Permanent Monetary Shock, X_t^m	0.10	0.84	0.35	0.37	0.07	0.01	0.05
Transitory Monetary Shock, z_t^m	0.08	0.01	0.06	0.00	0.01	0.01	0.01
Permanent Nonmonetary Shock, X_t	0.27	0.06	0.51	0.05	0.71	0.05	0.90
Transitory Nonmonetary Shock, z_t	0.50	0.03	0.00	0.00	0.20	0.00	0.00
Permanent Foreign Monetary Shock, X_t^{m*}	0.05	0.05	0.08	0.58	0.00	0.94	0.04
	B. Japan						
	Δy_t	π_t	i_t	$\ln S_t$	$\ln e_t$	i_t^*	$i_t - i_t^* - \epsilon_{t+1}$
Permanent Monetary Shock, X_t^m	0.26	0.82	0.57	0.50	0.01	0.00	0.12
Transitory Monetary Shock, z_t^m	0.04	0.07	0.08	0.01	0.01	0.00	0.03
Permanent Nonmonetary Shock, X_t	0.18	0.07	0.33	0.35	0.98	0.06	0.81
Transitory Nonmonetary Shock, z_t	0.44	0.04	0.01	0.02	0.00	0.00	0.01
Permanent Foreign Monetary Shock, X_t^{m*}	0.08	0.01	0.02	0.12	0.00	0.93	0.03

Notes. Notation: Δy_t, U.S. output growth; π_t, U.S. inflation; i_t, the federal funds rate; $\ln S_t$, dollar-pound or dollar-yen nominal exchange rate; $\ln e_t$, the dollar-pound or dollar-yen real exchange rate; i_t^*, U.K. or Japanese nominal interest rate; ϵ_t, devaluation rate.

Fisherian dynamics.

6 The Importance of Permanent Monetary Shocks for Exchange Rates: Variance Decompositions

Thus far, we have argued that permanent monetary shocks are important for explaining movements in the exchange rate on the basis that this variable responds quite differently depending on whether the monetary disturbance is permanent or transitory. This argument would be of limited relevance if permanent shocks had a minor role in driving exchange rates at business-cycle frequencies. To shed light on this issue, table 2 reports forecast-error variance decompositions of the nominal and real dollar-pound and dollar-yen exchange rates, the uncovered interest-rate differential, U.S. output growth, U.S. inflation, and the U.S., U.K., and Japanese policy rates at horizon 36 months. The choice of horizon follows Eichenbaum and Evans (1995).

The picture that emerges from the table is that permanent monetary shocks are the main drivers of the nominal exchange rate. Jointly, the domestic and foreign permanent monetary shocks, X_t^m and X_t^{m*}, explain 95 and 62 percent of the forecast-error variance of the dollar-pound and dollar-yen nominal exchange rates, respectively. The domestic (U.S.) permanent monetary shock plays a significant role, explaining 37 and 50 percent of this variance for the dollar-pound and dollar-yen nominal exchange rates, respectively.

Further, independently of whether the foreign block is taken to be the United Kingdom or Japan, the domestic permanent monetary shock is estimated to be an important driver of domestic nominal variables. It accounts for the majority of the forecast-error variance of U.S. inflation, 82 to 84 percent, and for a significant fraction of the variance of the federal funds rate, 35 to 57 percent. The domestic permanent monetary shock also explains a nonnegligible fraction of the forecast-error variance of domestic real output growth, 10 to 26 percent. These last three findings are similar to those obtained by Uribe (2018) in the context of closed-economy empirical and optimizing models estimated on U.S. and Japanese quarterly data.

By contrast, the domestic monetary shock, z_t^m, plays a negligible role in accounting for short-run movements in the nominal exchange rate. Table 2 shows that z_t^m explains no more than one percent of the forecast-error variance of the dollar-pound or dollar-yen nominal exchange rate. This share is much smaller than the one reported in other studies. For example, Eichenbaum and Evans (1995) estimate that their identified monetary shock explains, respectively, 19 and 22 percent of the forecast-error variance of the dollar-pound and dollar-yen nominal exchange rates at horizons 31 to 36 months.[6] One possible source of discrepancy could be that their estimation does not distinguish between temporary and permanent monetary disturbances.

Unlike the nominal exchange rate, the real exchange rate is estimated to be driven primarily by real shocks. Table 2 shows that taken together, the three monetary shocks explain

[6]These point estimates, however, are reported to be statistically insignificant at a 10-percent confidence level (see their table Ib).

less than 10 percent of movements in $\ln e_t$. Of the two real shocks, the permanent one, X_t, accounts for the lion share of movements in the real exchange rate, with a share of 71 percent in the United Kingdom and 98 percent in Japan. The modest role of monetary shocks in accounting for movements in the real exchange rate is consistent with that reported in Clarida and Galí (1994). For the United Kingdom, these authors find that monetary shocks explain 0.4 percent of the forecast-error variance of the level of the real exchange rate at a horizon of 12 quarters. For Japan at the same forecasting horizon, they estimate a larger contribution of 15 percent, but with a standard error of 13 percent, reflecting significant sampling uncertainty.

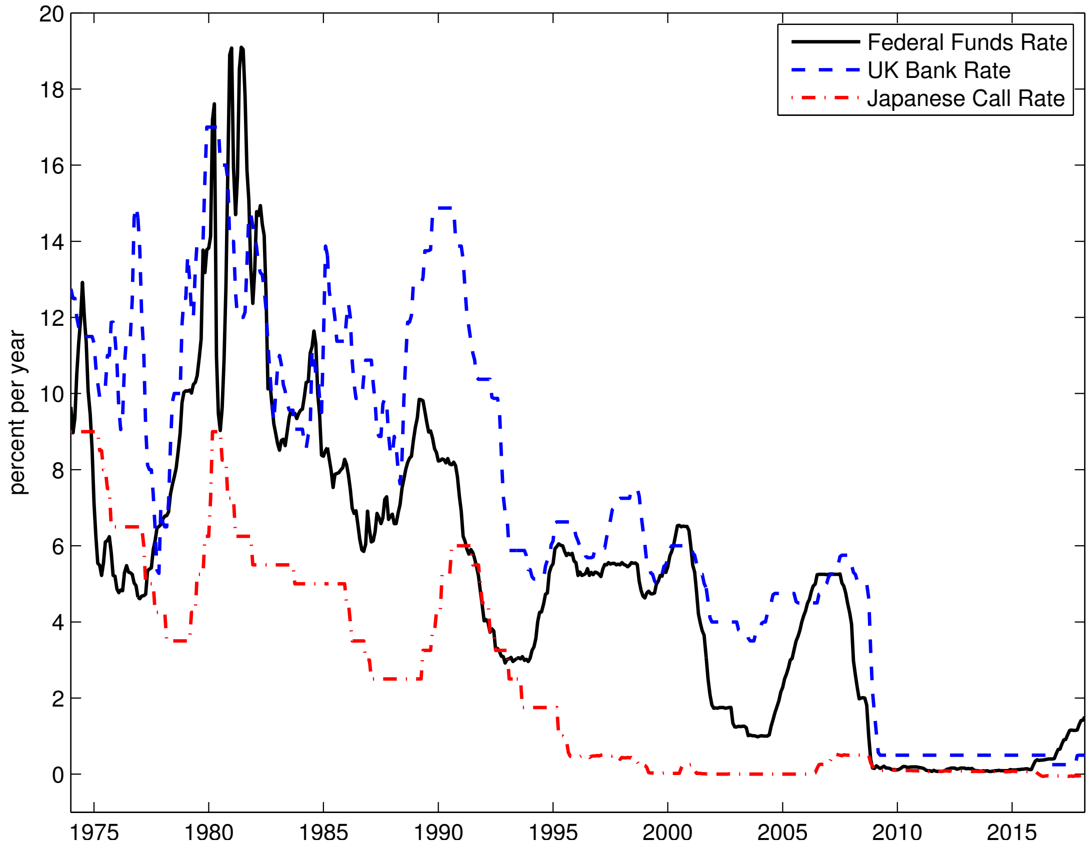
The main result of this section, namely, that permanent monetary shocks play an important role in explaining the variance of forecast errors of the nominal exchange rate, continues to hold at other horizons relevant for business-cycle analysis. Tables B.1 and B.2 of appendix B show results for horizons 12 to 48 months for the United Kingdom and Japan, respectively. As expected, permanent monetary shocks continue to be important at horizons longer than 36 months. The noteworthy result is that even at a horizon as short as 12 months, permanent monetary shocks explain 49 and 46 percent of the forecast-error variance of the nominal exchange rate in the United Kingdom and Japan, respectively.

In sum, this section documents that permanent monetary shocks represent an important source of short-run variation in nominal exchange rates. At the same time, permanent real shocks emerge as the main driver of movements in the real exchange rate.

7 Cointegrated Monetary Policies

Arguably, monetary policy could be coordinated across countries. For example, it is conceivable that an economy smaller than but similar in degree of development and institutional structure and related through trade in goods and financial assets to the United States might follow an interest rate policy with a permanent component that keeps pace with that of its

Figure 6: The Policy Interest Rate: United States, United Kingdom, and Japan, 1974:1 to 2018:3



Notes. Interest rates are expressed in percent per year and observed at a monthly frequency.

larger partner. Figure 6 displays the nominal interest rate in the United States, the United Kingdom, and Japan over the sample period 1974:1 to 2018:3. It is apparent from the figure that the changes in the deviations of the U.K. bank rate from the federal funds rate are quite transitory, suggesting that the two policy rates might share a common permanent component. Such relationship is less apparent in the case of Japan, especially since 1995 when the Bank of Japan embarked in a zero-rate policy from which it has not yet emerged. The conjecture that the U.S. and U.K. policy rates share a common permanent component, but the U.S. and Japanese pair does not, is supported by standard unit-root tests performed on the interest-rate differential $i_t - i_t^*$.[7]

Accordingly, we now formulate a variant of the empirical model that assumes that i_t and i_t^* are cointegrated but can exhibit short-run deviations from each other, and then estimate the new model on the dollar-pound dataset. In the new model, equation (1) is unchanged. In equation (2), the fifth row now features the variable $X_t^m - X_t^{m*}$ in lieu of ΔX_t^{m*}, to yield

$$\begin{bmatrix} \Delta X_{t+1}^m \\ z_{t+1}^m \\ \Delta X_{t+1} \\ z_{t+1} \\ X_{t+1}^m - X_{t+1}^{m*} \end{bmatrix} = \rho \begin{bmatrix} \Delta X_t^m \\ z_t^m \\ \Delta X_t \\ z_t \\ X_t^m - X_t^{m*} \end{bmatrix} + \psi \begin{bmatrix} \nu_{t+1}^1 \\ \nu_{t+1}^2 \\ \nu_{t+1}^3 \\ \nu_{t+1}^4 \\ \nu_{t+1}^5 \end{bmatrix}.$$

We continue to assume that the matrices ρ and ψ are diagonal.

The assumption that $X_t^m - X_t^{m*}$ is stationary induces stationarity in both the interest rate differential, $i_t - i_t^*$, and the devaluation rate, ϵ_t. Accordingly, we use these two variables

[7] An ADF test with 12 lags rejects the null hypothesis of the presence of a unit root with a p-value of 0.049 for the U.S.-U.K. interest-rate differential, but fails to rejected it with a p-value of 0.123 for the U.S.-Japan interest-rate differential.

as observables instead of Δi_t^* and $\Delta\epsilon_t$.[8] Thus, the vector of observables becomes

$$o_t = \begin{bmatrix} \Delta y_t \\ r_t \\ \Delta i_t \\ \epsilon_t \\ i_t - i_t^* \end{bmatrix} + \mu_t,$$

and, correspondingly, the observation equations become

$$\Delta y_t = \hat{y}_t - \hat{y}_{t-1} + \Delta X_t$$

$$r_t = \hat{i}_t - \hat{\pi}_t$$

$$\Delta i_t = \hat{i}_t - \hat{i}_{t-1} + \Delta X_t^m$$

$$\epsilon_t = \hat{\epsilon}_t + X_t^m - X_t^{m*}$$

and

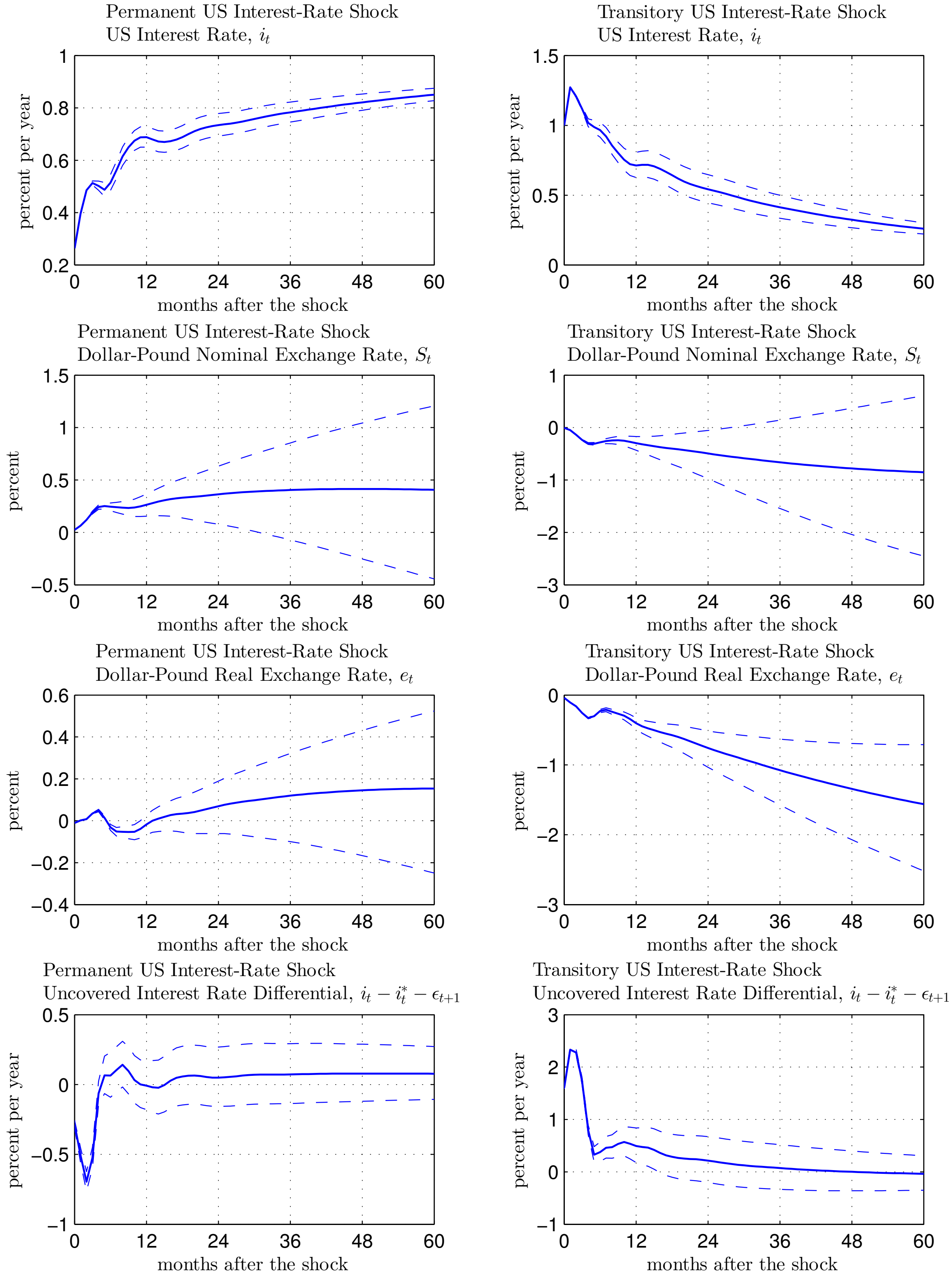
$$i_t - i_t^* = \hat{i}_t - \hat{i}_t^* + X_t^m - X_t^{m*}.$$

In this formulation, only the last two observation equations differ from their baseline counterparts. Using the identifiability test of Iskrev (2010), we find that the parameters of the model are identifiable in the neighborhood of the posterior mean.

Figure 7 displays impulse responses to permanent and transitory monetary shocks in the model with cointegrated policies. By construction, the response of the level of the nominal exchange rate to a permanent U.S. tightening is now bounded, since the devaluation rate is stationary. By and large, the results obtained under the baseline model carry over to the new specification. In particular, (a) the nominal exchange rate depreciates in response

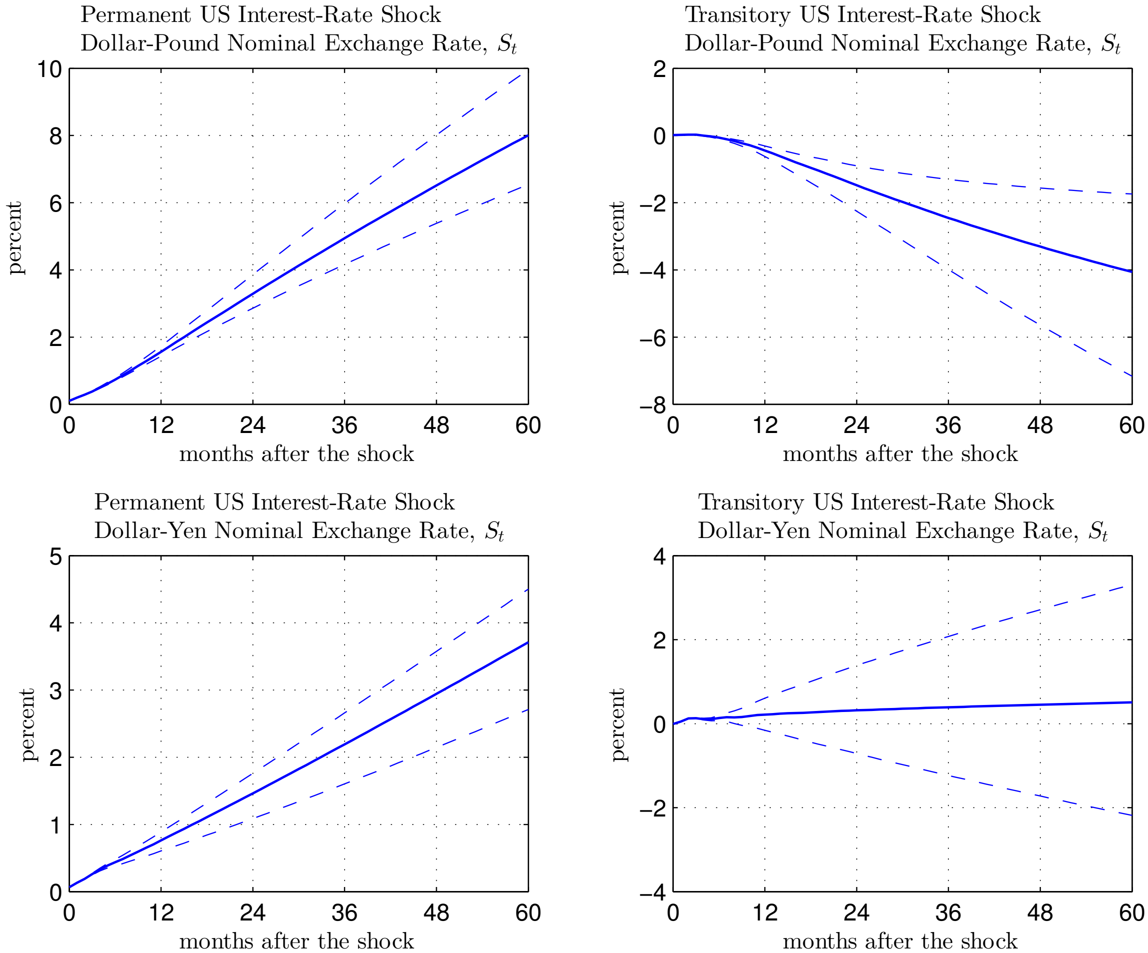
[8]One could have kept as observables Δi_t^* and $\Delta\epsilon_t$, as both are stationary and can be linked to the latent state variables. However, using ϵ_t and $i_t - i_t^*$ as observables has the advantage of requiring less time differencing.

Figure 7: Impulse Responses to Permanent and Transitory U.S. Monetary Shocks Under Cointegrated U.S. and U.K. Monetary Policies



Note. See notes to figure 1.

Figure 8: Impulse Responses of the Nominal Exchange Rate to Permanent and Transitory U.S. Monetary Shocks: 1974:1 to 1987:12



Notes. The model is estimated on the subsample period 1974:1 to 1987:12. See notes for figure 1.

to a permanent tightening and appreciates in response to transitory tightening. The same effects hold for real exchange rates; (b) there is no exchange-rate overshooting; and (c) the uncovered interest-rate differential moves in favor of U.S. assets in response to a temporary tightening but against U.S. assets in response to a permanent tightening.

8 The Volcker Era and Exchange Rate Overshooting

An open question in the existing empirical literature on the effect of monetary policy on exchange rates is whether in response to an interest-rate tightening the nominal exchange rate exhibits a delayed or an immediate overshooting effect. Kim, Moon, and Velasco (2017) argue that the delayed overshooting effect is a feature of the Volcker era. In this paper, we argue that once one allows for both transitory and permanent monetary shocks, the overshooting effect, either immediate or delayed, disappears altogether. In this section, we show that this result is robust to truncating the sample in December of 1987, the last year of Volcker's chairmanship.

Figure 8 displays the responses of the dollar-pound and dollar-yen nominal exchange rates to permanent and transitory increases in the federal funds rate when the model is estimated on the shorter sample. Naturally, due to the reduced estimation period, the impulse responses are estimated with significant sampling uncertainty, which is reflected in wide error bands. Nonetheless, the main result conveyed by the figure is the absence of the overshooting effect whether of a delayed or instantaneous type. A potential explanation of this finding is that overshooting could be the consequence of not explicitly distinguishing between temporary and permanent monetary disturbances.

9 Conclusion

Existing empirical studies have documented that a monetary shock that increases the domestic interest rate causes an appreciation of the domestic currency with an overshooting effect, that is, an appreciation of the domestic currency that is larger in the short run than in the long run and a persistent deviation from uncovered interest parity in favor of domestic assets.

In this paper, we estimate an empirical model of exchange rates that allows for permanent and transitory monetary shocks. Using monthly data from the United States, the United Kingdom, and Japan for the post Bretton-Woods period, we obtain the following four insights. First, in the short run permanent monetary policy shocks depreciate the domestic currency whereas temporary ones appreciate it. Second, once permanent and transitory monetary shocks are modeled separately, the overshooting effect disappears: Transitory increases in the nominal interest rate cause persistent appreciations of the domestic currency, whereas permanent increases cause persistent depreciations. This finding suggests that the existing overshooting results may be the consequence of confounding impulses. Third, both transitory and permanent increases in the domestic nominal interest rate cause persistent deviations from uncovered interest-rate parity, but of opposite signs, the former in favor of domestic assets and the latter against. Fourth, permanent monetary shocks explain the majority of short-run movements in the nominal exchange rate, while transitory monetary shocks play a minor role. Real exchange rates are driven primarily by permanent real shocks, with a negligible role for monetary disturbances.

A pending issue in this line of investigation is to capture the effects of monetary policy shocks on exchange rates documented in this paper in the context of an equilibrium optimizing model of nominal and real exchange rate determination. Certainly, a challenge that such a theoretical endeavor will have to confront is to generate short-run deviations from uncovered interest rate parity that are of opposite sign depending on whether the monetary shock is transitory or permanent. This aspect of the propagation mechanism will most likely prove crucial for explaining observed exchange rate dynamics. We leave these tasks for future research.

Appendix A

In this appendix, we present the empirical model in more detail showing explicitly its associated intercepts that we had omitted earlier to simplify the exposition. Let

$$\widehat{Y}_t \equiv \begin{bmatrix} y_t - X_t - E(y_t - X_t) \\ \pi_t - X_t^m - E(\pi_t - X_t^m) \\ i_t - X_t^m - E(i_t - X_t^m) \\ \epsilon_t - X_t^m + X_t^{m*} - E(\epsilon_t - X_t^m + X_t^{m*}) \\ i_t^* - X_t^{m*} - E(i_t^* - X_t^{m*}) \end{bmatrix}$$

and

$$u_t \equiv \begin{bmatrix} \Delta X_t^m - E(\Delta X_t^m) \\ z_t^m \\ \Delta X_t - E(\Delta X_t) \\ z_t \\ \Delta X_t^{m*} - E(\Delta X_t^{m*}) \end{bmatrix}.$$

The vector $\widehat{Y}_t$ evolves over time according to

$$\widehat{Y}_t = \sum_{i=1}^{L} B_i \widehat{Y}_{t-i} + C u_t$$

and the vector u_t according to

$$u_t = \rho u_{t-1} + \psi \nu_t.$$

Then to obtain a first-order state space representation let the vector ξ_t be given by

$$\xi_t \equiv \begin{bmatrix} \widehat{Y}_t' & \widehat{Y}_{t-1}' & \ldots & \widehat{Y}_{t-L+1}' & u_t' \end{bmatrix}'$$

With these definitions and notation in hand the empirical model becomes

$$\xi_{t+1} = F\xi_t + P\nu_{t+1}$$

and the observation equations can be written as

$$o_t = A' + H'\xi_t + \mu_t.$$

The relationship between the matrices B_i for $i = 1, \ldots, L$, C, ρ, and ψ and the matrices A, F, P, and H is as follows. Let V denote the number of variables included in the vector $\widehat{Y}_t$ and S the number of shocks in the vector ν_t. In the empirical implementation of the model $V = 5$ and $S = 5$. Further, let

$$B \equiv [B_1 \cdots B_L],$$

and let I_j denote an identity matrix of order j, $\emptyset_j$ denote a square matrix of order j with all entries equal to zero, and $\emptyset_{i,j}$ denote a matrix of order i by j with all entries equal to zero. Then, for $L \geq 2$ we have

$$F = \begin{bmatrix} B & C\rho \\ \left[I_{V(L-1)}\, \emptyset_{V(L-1),V}\right] & \emptyset_{V(L-1),S} \\ \emptyset_{S,VL} & \rho \end{bmatrix},$$

$$P = \begin{bmatrix} C\psi \\ \emptyset_{V(L-1),S} \\ \psi \end{bmatrix};$$

$$A' = \begin{bmatrix} E(\Delta X_t) \\ E(i_t - \pi_t) \\ E(\Delta X_t^m) \\ E(\Delta X_t^m - \Delta X_t^{m*}) \\ E(\Delta X_t^{m*}) \end{bmatrix}$$

and

$$H' = \begin{bmatrix} M_\xi & \emptyset_{V,V(L-2)} & M_u \end{bmatrix},$$

where the matrices M_ξ and M_u take the form

$$M_\xi = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & -1 & 0 & 0 & 0 & 0 \\ 0 & -1 & 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & -1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & -1 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & -1 \end{bmatrix}$$

and

$$M_u = \begin{bmatrix} 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & -1 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix}.$$

Appendix B: Variance Decomposition at Horizons between 12 and 48 months

Table B.1: Forecast Error Variance Decomposition at Horizons between 12 and 48 months: United Kingdom

	Δy_t	π_t	i_t	$\ln S_t$	$\ln e_t$	i_t^*	$i_t - i_t^* - \epsilon_{t+1}$
Permanent Monetary Shock, X_t^m							
Horizon 12 months	0.10	0.83	0.23	0.23	0.00	0.01	0.05
Horizon 24 months	0.10	0.83	0.31	0.36	0.03	0.01	0.05
Horizon 36 months	0.10	0.84	0.35	0.37	0.07	0.01	0.05
Horizon 48 months	0.10	0.85	0.37	0.36	0.10	0.00	0.05
Transitory Monetary Shock, z_t^m							
Horizon 12 months	0.08	0.01	0.23	0.00	0.00	0.01	0.01
Horizon 24 months	0.08	0.01	0.10	0.00	0.00	0.01	0.01
Horizon 36 months	0.08	0.01	0.06	0.00	0.01	0.01	0.01
Horizon 48 months	0.08	0.01	0.05	0.00	0.02	0.01	0.01
Permanent Nonmonetary Shock, X_t							
Horizon 12 months	0.26	0.07	0.47	0.50	0.98	0.06	0.93
Horizon 24 months	0.27	0.07	0.51	0.14	0.88	0.05	0.91
Horizon 36 months	0.27	0.06	0.51	0.05	0.71	0.05	0.90
Horizon 48 months	0.27	0.06	0.49	0.03	0.59	0.04	0.89
Transitory Nonmonetary Shock, z_t							
Horizon 12 months	0.52	0.05	0.00	0.00	0.01	0.00	0.00
Horizon 24 months	0.51	0.04	0.00	0.00	0.09	0.00	0.00
Horizon 36 months	0.50	0.03	0.00	0.00	0.20	0.00	0.00
Horizon 48 months	0.50	0.03	0.00	0.00	0.28	0.00	0.00
Permanent Foreign Monetary Shock, X_t^{m*}							
Horizon 12 months	0.03	0.04	0.06	0.26	0.00	0.92	0.01
Horizon 24 months	0.04	0.05	0.07	0.50	0.00	0.93	0.03
Horizon 36 months	0.05	0.05	0.08	0.58	0.00	0.94	0.04
Horizon 48 months	0.05	0.05	0.09	0.61	0.00	0.95	0.05

Notes. See notes to table 2.

Table B.2: Forecast Error Variance Decomposition at Horizons between 12 and 48 months: Japan

	Δy_t	π_t	i_t	$\ln S_t$	$\ln e_t$	i_t^*	$i_t - i_t^* - \epsilon_{t+1}$
Permanent Monetary Shock, X_t^m							
Horizon 12 months	0.28	0.73	0.55	0.39	0.01	0.01	0.11
Horizon 24 months	0.26	0.80	0.58	0.54	0.01	0.00	0.12
Horizon 36 months	0.26	0.82	0.57	0.50	0.01	0.00	0.12
Horizon 48 months	0.26	0.83	0.56	0.48	0.02	0.00	0.12
Transitory Monetary Shock, z_t^m							
Horizon 12 months	0.04	0.15	0.26	0.01	0.00	0.01	0.03
Horizon 24 months	0.04	0.10	0.13	0.01	0.00	0.00	0.03
Horizon 36 months	0.04	0.07	0.08	0.01	0.01	0.00	0.03
Horizon 48 months	0.04	0.05	0.06	0.01	0.01	0.00	0.03
Permanent Nonmonetary Shock, X_t							
Horizon 12 months	0.13	0.03	0.12	0.51	0.99	0.12	0.84
Horizon 24 months	0.17	0.05	0.26	0.32	0.99	0.08	0.82
Horizon 36 months	0.18	0.07	0.33	0.35	0.98	0.06	0.81
Horizon 48 months	0.19	0.08	0.35	0.38	0.96	0.05	0.81
Transitory Nonmonetary Shock, z_t							
Horizon 12 months	0.48	0.08	0.00	0.01	0.00	0.01	0.01
Horizon 24 months	0.45	0.05	0.00	0.02	0.00	0.00	0.01
Horizon 36 months	0.44	0.04	0.01	0.02	0.00	0.00	0.01
Horizon 48 months	0.44	0.03	0.01	0.01	0.00	0.00	0.01
Permanent Foreign Monetary Shock, X_t^{m*}							
Horizon 12 months	0.08	0.01	0.06	0.07	0.00	0.85	0.01
Horizon 24 months	0.08	0.01	0.03	0.11	0.00	0.90	0.02
Horizon 36 months	0.08	0.01	0.02	0.12	0.00	0.93	0.03
Horizon 48 months	0.08	0.01	0.02	0.12	0.00	0.94	0.03

Notes. See notes to table 2.

References

Bjørnland, Hilde C., "Monetary policy and exchange rate overshooting: Dornbusch was right after all," *Journal of International Economics 79,* 2009, 64-77.

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans, "Nominal Rigidities and the Dynamic Effects of a Shock to Monetary Policy," *Journal of Political Economy 113,* 2005, 1-45.

Clarida, Richard, and Jordi Galí, "Sources of real exchange-rate fluctuations: How important are nominal shocks?," *Carnegie-Rochester Conference Series on Public Policy 41,* 1994, 1-56.

De Michelis, Andrea, and Matteo Iacoviello, "Raising an inflation target: The Japanese experience with Abenomics," *European Economic Review 88,* 2016, 67-87.

Dornbusch, Rüdiger, "Expectations and Exchange Rate Dynamics," *Journal of Political Economy 84,* December 1976, 1161-1171.

Eichenbaum, Martin, and Charles L. Evans, "Some Empirical Evidence on the Effects of Shocks to Monetary Policy on Exchange Rates," *Quarterly Journal of Economics 110,* November 1995, 975-1009.

Engel, Charles, "Exchange Rates, Interest Rates, and the Risk Premium," *American Economic Review 106,* February 2016, 436-474.

Faust, Jon, and John H. Rogers, "Monetary policy's role in exchange rate behavior," *Journal of Monetary Economics 50,* 2003, 1403-1424.

Hamilton, James D., *Time Series Analysis,* Princeton University Press: Princeton, NJ, 1994.

Hettig, Thomas, Gernot Müller, and Martin Wolf, "Exchange Rate Undershooting: Evidence and Theory," mimeo, University of Tübingen, 2019.

Iskrev, Nicolay, "Local Identification in DSGE Models," *Journal of Monetary Economics 57,* 2010, 189-210.

Kim, Seong-Hoon, Seongman Moon, and Carlos Velasco, "Delayed Overshooting: Is It an

80s Puzzle?," *Journal of Political Economy 125,* 2017, 1570-1598.

Kim, Soyoung, and Nouriel Roubini, "Exchange rate anomalies in the industrial countries: A solution with a structural VAR approach," *Journal of Monetary Economics 45,* 2000, 561-586.

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

Scholl, Almuth, and Harald Uhlig, "New evidence on the puzzles: Results from agnostic identification on monetary policy and exchange rates?," *Journal of International Economics 76,* 2008, 1-13.

Sims, Christopher, and Tao Zha, "Error Bands for Impulse Responses," *Econometrica 67,* 1999, 1113-1156.

Uhlig, Harald, "What are the effects of monetary policy on output? Results from an agnostic identification procedure," *Journal of Monetary Economics 52,* 2005, 381-419.

Uribe, Martín, "The Neo-Fisher Effect: Econometric Evidence from Empirical and Optimizing Models," NBER working paper 25089, September 2018.