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Using Long-Run Restrictions to Investigate the Sources of Exchange Rate Fluctuations*

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Abstract

This paper makes use of long-run restrictions to identify macroeconomic shocks and evaluate their relative importance for exchange rate fluctuations. Unlike previous studies that employ a similar approach, I consider a large eight variable vector autoregressive system that includes short term interest rates rather than money stocks in order to help identify monetary policy shocks. Results for the U.S. and the U.K. show that monetary policy shocks and other macroeconomic shocks behave according to theory. However, monetary shocks account for only a small fraction of the variance of the real exchange rate. Instead, "taste shocks" that can be associated with the degree of trade openness, terms of trade, and current account appear to be the key factor driving the U.S.-U.K. real exchange rate. Results for other countries under consideration (Canada, Germany, and Japan) are similar.

JEL Classification: F31

Key Words: *vector autoregression; taste shocks; monetary shocks; exchange rate movements; long-run identifying restrictions.*

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1. Introduction

The volatile nature of exchange rate movements since the collapse of the Bretton Woods system of fixed exchange rates has led economists to consider three important questions: What are the sources of exchange rate fluctuations? Are monetary policy shocks the main factor? Do other macroeconomic shocks matter more? In this paper, I present an eight-variable vector autoregression (VAR) model using long-run identification restrictions to address the long-standing issue of whether nominal/monetary shocks matter for exchange rates. In contrast to other empirical papers using a similar identification strategy, my larger VAR system allows for the inclusion of many relevant domestic and foreign macroeconomic variables. In addition, I use short-term interest rates, as advocated by Bernanke and Blinder (1992), rather than the usual money stocks to help identify monetary policy shocks. Results in this paper show that nominal shocks are not important for the bilateral exchange rates between the U.S. and four other G7 countries (U.K., Canada, Germany, and Japan). Other macroeconomic shocks identified in the VAR, such as supply or commodity price shocks, also have little influence on the real exchange rate. Instead, exchange rate fluctuations appear to arise from "taste" shocks that can be related to the international trade sector of the economies under investigation.

The remainder of the paper will proceed as follows: Section 2 gives a brief literature survey. Section 3 details the VAR model under consideration, the data, and the shock identification strategy. Section 4 shows the results of the VAR estimation. Section 5 investigates the properties and the sources of the taste shock. Finally, Section 6 offers concluding remarks.

2. Literature Review

The nature of the shocks that lead to exchange rate fluctuations has been a source of contention for economists for a long time. In an influential paper, Mussa (1986) argues that sluggish price adjustment must be the key factor in explaining the short-run movements in real and nominal exchange rates. This of course implies that the interaction of sticky prices and monetary shocks could have been the source of volatile exchange rates in the post-Bretton Woods era. On the other hand, Stockman (1987) disputes the idea that monetary shocks are to blame for the behavior of real exchange rates after the collapse of Bretton Woods. He argues that real shocks with large permanent components are the main culprits. With competing theories explaining exchange rate fluctuations, the debate must be brought to the data. Indeed, there exists a large body of empirical work in the area. However, even empirical studies on exchange rates have failed to reach consensus on whether monetary shocks matter for exchange rate variability. Some papers suggest little or no role for monetary shocks while other papers have found that monetary shocks are the most important in driving exchange rate movements. For example, Grilli and Roubini (1995) report the share of the dollar-pound exchange rate variance accounted for by monetary shocks to be as low as 2 percent while Rogers (1999) reports a share as high as 41 percent. This large discrepancy primarily reflects the major difficulty in empirical work on exchange rates: how to correctly identify monetary policy shocks and judge their relative importance.

The most common approach to identification of economic shocks involves the imposition of short-run restrictions within a VAR model. In particular, some of the contemporaneous effects of shocks on the variables in the VAR are restricted to zero. These restrictions can be either recursive or non-recursive, though under both categories the assumptions made can be rather implausible. For example, Eichenbaum and Evans (1995) assume that foreign interest rates do not respond to Federal Reserve policy shocks until a month after policy is changed, which is inconsistent with large movements in foreign rates immediately after the Federal Open Market Committee's (FOMC) policy announcements. In works that employ this kind of identification procedure, the estimated range of the share of monetary shocks in the total variance in real or nominal exchange rate is quite large, from around 2 percent (Grilli and Roubini 1995, U.S.-U.K. nominal exchange rate on impact) to 34 percent (Kim and Roubini 2000, U.S.-U.K. 7 variable model, six-month horizon for nominal exchange rate).

Identification of VAR models can also be achieved with long-run restrictions, as originally advocated by Blanchard and Quah (1989). With this method, some shocks (most likely nominal/monetary shocks) are assumed to have no long-run effect on real economic activity. These restrictions often make intuitive economic sense. They also allow for easy structural interpretation of all of the shocks in the VAR system. Most other popular techniques of identification, such as recursive and non-recursive short-run restrictions or sign/shape restrictions,[1] typically only allow for "partial identification" of the shock of interest without giving an interpretation to all the other shocks in the system. Clarida and Gali (1994) provide the seminal investigation of the effects of real and nominal shocks on real exchange rate by using long-run identification restrictions. I implement the same identification approach in this paper. However, unlike Clarida and Gali (1994) or other previous papers that employ a similar identification strategy, I estimate a much larger VAR system that includes many potentially relevant variables. (Clarida and Gali's model has only three variables, while I include eight variables.) In addition, I use short-term interest rates rather than the usual money stocks to identify monetary policy shocks.

Similar to studies using short-run identification restriction, results on the importance of monetary policy shocks for real exchange rates using long-run restrictions are often at odds with each other. Clarida and Gali (1994) suggest monetary shocks are unimportant (their highest estimated share of variance due to monetary shocks is 2.2 percent for U.S.-U.K. real exchange rate), while Rogers (1999) finds that the contribution of monetary shocks can be as high as 40.6 percent for U.S.-U.K. real exchange rate. My results correspond well with the findings in Clarida

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[1] Sign and shape restrictions are fairly recent developments in the area of VAR identification. In Canova and De Nicoló (2002), Faust (1998), and Uhlig (2005), the general idea is to systematically examine a variety of identification schemes, and then, through elimination by penalty functions or sign/shape restrictions on the impulse-response functions, find a unique solution. This approach is well suited to assessing the robustness of certain claims from identified VAR work. However, the formal restrictions imposed to arrive at the final choice, such as sign restrictions, are still subjective and some would argue even more restrictive than the short-run or long-run restriction approaches. Also, Faust (1998) and Uhlig (2005) only "partially" identify the structural model. Hence, besides the shock of interest, one cannot examine what other shock in the system may have important effect on the real exchange rate. Farrant and Peersman (2006) modified the Uhlig (2005) method to allow the full set of shocks to be identified by imposing a larger collection of sign restrictions. However, this is only feasible in relatively small VAR systems (the largest system in Farrant and Peersman 2006 has just four variables) as it becomes increasingly difficult to impose credible sign restrictions when the number of variables in the model gets larger.

and Gali (1994) in spite of the differences in the included variables in our models. In particular, I find that monetary policy shocks only account for about 2 percent of the total variance of the dollar-pound real exchange rate on impact. Contrast that to the "taste" shock, which accounts for close to 70 percent of the total variance of the dollar-pound real exchange rate on impact.

Unlike other studies in the area, this paper goes a step further to investigate the potential sources for the taste shock. Based on the VAR analysis, this taste shock does not appear to be associated with traditional demand-type disturbances that would have only a short-lived impact on output and interest rates. Instead, using regression analysis, I find that changes in relative trade openness, relative terms of trade, and relative current account between the country pairs could be important factors driving the taste shock. The result that real shocks originating from the demand side (in particular the international trade sector of the economy) are the main sources of exchange rate fluctuations distinguishes this paper from previous studies that have shown monetary disturbances to be unimportant. Most would argue that if real shocks matter for exchange rates, it should be coming in from the supply side through productivity type disturbances that impact exchange rates via the Balassa-Samuelson effect. So the results here provide empirical evidence for economists who have argued that demand side macroeconomic fundamentals can be an important determinant of the exchange rate.[2]

3. VAR Model Specification and Identification Scheme

3.1 Data

Four country pairs are considered in this paper: U.S.-U.K., U.S.-Canada, U.S.-Germany, and U.S.-Japan. The U.S.-U.K. VAR model will be the benchmark for easy comparison with

[2] There is some theoretical support for the linkage between exchange rates and the international trade sector. Choi (2005) develops a theoretical macroeconomic model that justifies a trade based representation of the real exchange rate (real exchange rate as a function of international trade flows, among other things). She shows that this trade based representation is highly correlated with actual real exchange rates for a wide range of countries, leading her to conclude that real exchange rates are closely connected to international trade flows and macroeconomic fundamentals.

other papers in the literature. The national currencies of these countries are among the most heavily traded in the world. This may in part reflect their status as major trading partners with the U.S.[3] These countries are also selected to facilitate comparisons with Clarida and Gali (1994) and Rogers (1999) and to represent distinctly different trading areas (Non-continental Europe, North America, Continental Europe, and East Asia). The sample period is 1970Q2 to 2006Q1 for all countries except Canada (1970Q2 to 2006Q2) and Germany (1970Q2 to 2005Q4).

Eight variables, including output, exchange rate, prices, and interest rates, are used in the estimation of each U.S.-foreign country VAR. All the variables are in natural logs (except for the interest rate variables) and demeaned. Please refer to Table 1 for details of the variables and their corresponding data sources. Estimation of the VAR model requires that each of the variables entering the VAR is stationary. Series that are non-stationary should be transformed appropriately prior to estimation, otherwise finite-sample inferences may suffer serious distortions. Because the standard Augmented Dickey Fuller (ADF) test and the KPSS test developed by Kwiatkowski, Phillips, Schmidt, and Shin (1992) show that in general all variables except for interest rates are integrated of order one in levels, they enter into the VAR in first differences.

An important assumption in my model is that there are permanent shocks to the real exchange rate. This is only possible if the real exchange rate series are non-stationary in levels. It is common knowledge that testing for the presence of unit roots in exchange rates is extremely difficult. Research in this area has presented evidence on both sides of the argument. Rogoff (1996) provides an excellent summary of the debate. The general consensus is that in the short-run, real exchange rate fluctuations are too persistent to be justifiably monetary in nature. Hence the fluctuations can be treated as effectively permanent, implying the presence of unit root in real exchange rates. But over the very long-run (over one hundred years of data) one can find more

[3] In terms of total trade, according to the most recent trade data (May 2009) from the Census Bureau, Canada is the U.S.'s number 1 trading partner, Japan is number 4, Germany is number 5, and the U.K. is number 6. These rankings have not changed much for the last twenty years or so. All countries considered here have been top ten trading partners of the U.S. over the past decades, though the rankings for Japan and the U.K. have dropped back somewhat in recent years due to the increase in U.S. trade with China and Mexico.

evidence that the exchange rate conforms to some version of the purchasing power parity (PPP) condition (see Edison 1987). Since the purpose of this paper is not to determine if PPP holds over very long-run periods, but to determine what factors are important over the relatively shorter horizons (i.e. 20 to 30 years), I will proceed as if real exchange rates are non-stationary.

It has long been argued that there are structural breaks in the mean of the level of U.S. inflation during the sample period considered in this paper. Ignoring such breaks could yield spuriously high degree of persistence in the inflation series that might affect the VAR estimates. A wide range of potential break dates have been suggested for the U.S. inflation rate. Levin and Piger (2003) found a break in mean in 1991Q1 or 1991Q2 using four different measures of inflation, while Rapach and Wohar (2005) located three break dates (1967Q3, 1973Q1, and 1982Q1). Instead of adopting break dates reported in earlier studies, I test for structural break dates using a procedure based on Bai and Perron (1998). The results are presented in Table 2 along with some details of the procedure.[4] I have uncovered only one break for the U.S. inflation rate over my sample period. The break date 1981Q2 is in agreement with findings in the literature that inflation persistence was exceptionally high during the period from 1965 to the early 1980s, though whether persistence continued to be high since then, or has declined, is more hotly contested. Since a break in U.S. inflation could induce a break in the relative inflation variables as well, the same break date found for U.S. inflation is allowed for each of the relative inflation variables. Then a search for additional breaks is carried out. All of the relative inflation measures appear to have multiple structural breaks. Rapach and Wohar (2005) have also found multiple breaks in thirteen industrialized countries' inflation rates.

The structural break for U.S. inflation implies that the short-term interest rate variables in the VAR could also have structural breaks. Caporale and Grier (2000) and Bai and Perron (2003) have both found that multiple structural breaks exist in U.S. real interest rates. To locate potential structural breaks in the interest rate variables, I start by imposing the break date found for U.S.

[4] Since GDP deflator is used as the price data for Germany, I have to construct relative prices with U.S. GDP deflator as well. Due to this complication, the structural breaks for the Δp and $\Delta(p - p^*)$ variables in the U.S.-German case differs somewhat from the other country pairs.

inflation rate (Δp) on the 3-month treasury bill rate (i_{tbr}) and then implement the same procedure as used for inflation rates to search for additional breaks in the interest rate variables. As shown in Table 2, four break dates are found for i_{tbr}.[5] To be consistent, I impose these same break dates on the relative short-term rates ($i - i^{*}$) and search for further breaks. The relative interest rates tend to have more breaks than the U.S. domestic rates. This is possibly related to the findings in Rapach and Wohar (2005) of multiple structural breaks in real interest rates using international data.

3.2 Structural VAR Framework

For the benchmark model, the vector of variables of interest $\mathbf{x} \equiv [\Delta(y - y^*), \Delta y, \Delta q, \Delta p^c, \Delta(p - p^*), \Delta p, i - i^*, i_{tbr}]'$ is assumed to follow a multivariate covariance stationary process. The typical VAR representation assumes that the vector $\mathbf{x}$ depends on lags of itself and some vector of structural shocks $\boldsymbol{\varepsilon}$:

$$(1) \qquad A_0 \mathbf{x}_t = \sum_{j=1}^{k} A_j \mathbf{x}_{t-j} + \boldsymbol{\varepsilon}_t\,, \qquad \boldsymbol{\varepsilon}_t \sim N(0,D).$$

Note that the structural shocks are normally distributed with mean zero and that the variance covariance matrix D is diagonal (shocks are uncorrelated with each other). Provided that the coefficient matrix A_0 is invertible, equation (1) can be rewritten more compactly as

$$(2) \qquad A(L)\mathbf{x}_t = A_0^{-1}\boldsymbol{\varepsilon}_t\,,$$

where $A(L) = I - \sum_{j=1}^{k} A_0^{-1} A_j L^j$, and L is the lag operator. The Wold moving average (Wold MA) representation of equation (2) is then

[5] The same four break dates found for i_{tbr} are assumed for the federal funds rate i_{ffr} as well.

(3) $\mathbf{x}_t = C(L)\boldsymbol{\varepsilon}_t$,

where $C(L) = A(L)^{-1} A_0^{-1}$. Note here that *A(L)* would have to be invertible for equation (3) to make sense. The reduced-form Wold moving average representation of **x** is given by

(4) $\mathbf{x}_t = E(L)\mathbf{v}_t$, $\mathbf{v}_t \sim N(0,\Sigma)$.

Comparing equation (4) with equation (2) above, one can interpret *E(L)* to be equal to $A(L)^{-1}$, hence $E(0) = I$ and *E(L)* is invertible. As a consequence, $\mathbf{v}_t = A_0^{-1}\boldsymbol{\varepsilon}_t$ is the vector of innovations where each element in **v** is some linear combination of the structural shocks in **x**. The (reduced-form) autoregressive representation of the system in equation (4) can be given by

(5) $A(L)\mathbf{x}_t = \mathbf{v}_t$,

which is the same as equation (2) except that the right hand side of (2), $A_0^{-1}\boldsymbol{\varepsilon}$, is denoted by **v**. *A(L)* can be consistently estimated using standard ordinary least squares (OLS).[6] The residuals from the OLS regression can then be used to calculate *Σ*. The structural model, i.e., the coefficients of $C(L) = A(L)^{-1} A_0^{-1}$ will be identified to the extent that there are enough restrictions to determine the elements of *C(L)* uniquely. In the case of long-run restrictions, by making assumptions about the long-run behavior of the variables in the model that will render *C(1)* to be lower triangular, one can invoke the relationship that the spectral density for **x** at frequency zero is proportional to the long-run variance-covariance matrix denoted Λ:

(6) $\Lambda = E(1)\Sigma E(1)' = C(1)DC(1)'$,

[6] This is equivalent to estimating the model using conditional maximum likelihood under normality or using the SUR model with identical regressors in all equations.

such that Cholesky decomposing Λ provides a unique lower triangular matrix that is equivalent to $C(1)D^{1/2}$. Given $C(1)$ and $A(1)$, the impact matrix A_0^{-1} can be obtained, and the vector of structural shocks $\boldsymbol{\varepsilon}$ can then be recovered.

While long-run identification procedures are popular, there are some issues with their implementation. Faust and Leeper (1997) present two major criticisms. The first one is the problem of inference regarding the estimated $C(1)$ coefficients. As $C(1)$ estimates are inherently imprecise even in large samples, imposing long run restrictions transfers this uncertainty to all the structural parameters including coefficients of the impulse-response functions. To address this issue, I assume that the true model driving the data is a VAR with a known maximum lag order K, where K is determined by standard model selection procedures and is small relative to the sample size. In addition, I construct confidence intervals for the impulse-responses and variance decompositions with the more reliable bias corrected bootstrap method proposed by Kilian (1998). Kilian and Chang (2000) have shown that these confidence intervals (along with the Sims and Zha 1999 Bayesian Monte Carlo integration confidence intervals) have superior coverage accuracy when compared with the more common ways of constructing confidence intervals for impulse-responses, such as Runkle (1987) and Lütkepohl (1990).

Faust and Leeper (1997) were also concerned with the problem posed by multiple shocks. Since VAR is usually applied in low dimensional models, the identified shocks must be viewed as aggregates of a larger number of underlying shocks. So if one identified structural shock consists of two independent shocks, then the Blanchard and Quah long-run identification method is valid only if the underlying macroeconomic variables respond to the two shocks in the same way. My eight variable benchmark model, rather large for a VAR with long-run restrictions, should allow me to address this concern. Due to the size of the VAR, the identified shocks are disaggregated into a larger number of sensible categories: the supply and monetary shocks are decomposed into those that are common to both countries and those that are particular to only one of the countries; the monetary shocks are refined into money supply and money demand shocks. A commodity price shock is also introduced to allow for disturbances coming from the commodities market to be separated from productivity related supply shocks. Finally, a "taste" shock is allowed, which has permanent effects on the real exchange rate but not on output.

3.3 Identification of the VAR Model

The long-run restrictions imposed on the benchmark model can be expressed in the following Wold MA form:

$$(7) \qquad \mathbf{x} = C(1)\boldsymbol{\varepsilon}$$

$$\Downarrow$$

$$\begin{bmatrix} \Delta(y-y^*) \\ \Delta y \\ \Delta q \\ \Delta p^c \\ \Delta(p-p^*) \\ \Delta p \\ i-i^* \\ i_{tbr} \end{bmatrix} = \begin{bmatrix} C_{11}(1) & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ C_{21}(1) & C_{22}(1) & 0 & 0 & 0 & 0 & 0 & 0 \\ C_{31}(1) & C_{32}(1) & C_{33}(1) & 0 & 0 & 0 & 0 & 0 \\ C_{41}(1) & C_{42}(1) & C_{43}(1) & C_{44}(1) & 0 & 0 & 0 & 0 \\ C_{51}(1) & C_{52}(1) & C_{53}(1) & C_{54}(1) & C_{55}(1) & 0 & 0 & 0 \\ C_{61}(1) & C_{62}(1) & C_{63}(1) & C_{64}(1) & C_{65}(1) & C_{66}(1) & 0 & 0 \\ C_{71}(1) & C_{72}(1) & C_{73}(1) & C_{74}(1) & C_{75}(1) & C_{76}(1) & C_{77}(1) & 0 \\ C_{81}(1) & C_{82}(1) & C_{83}(1) & C_{84}(1) & C_{85}(1) & C_{86}(1) & C_{87}(1) & C_{88}(1) \end{bmatrix} \begin{bmatrix} \varepsilon^{s} \\ \varepsilon^{s-c} \\ \varepsilon^{d} \\ \varepsilon^{cp} \\ \varepsilon^{ms} \\ \varepsilon^{ms-c} \\ \varepsilon^{md} \\ \varepsilon^{md-c} \end{bmatrix}$$

The lower triangularity of $C(1)$ can be justified in a straightforward manner. Output is supply-driven in the long run,[7] hence shocks unrelated to the supply side of the economy should not have long run effects on output. For relative output, a supply shock that is common to both countries ($\varepsilon^{s\text{-}c}$) should not lead to long run differences between the two countries. Take a technological advancement as an example of a positive common supply shock. There may be short-run variations in the rate at which the countries incorporate this new technology into production of output, but over time there should be no major gaps in the outputs of the two countries that would lead to a shift in the relative output. Hence only a country specific supply shock (ε^{s}) would have long-run impact on relative output. For U.S. output, both common and

[7] Here I follow the arguments in Blanchard and Quah (1989). Demand factors may indeed have long-run impact on output, but the magnitude of the effect would be very small relative to that of supply disturbances. Hence I make the assumption that output is only influenced by supply shocks in the long-run.

relative supply shocks would have long run effects. This justifies the zeros in the first and second rows of $C(1)$.

For the real exchange rate, I only allow the supply shocks and the "taste" shock (ε^{d}) to have permanent effects (hence the zero restrictions on the third row of $C(1)$). Research on exchange rate determination shows that real factors from both the supply and demand sides of the economy may lead to long run changes in the real exchange rate, whereas nominal shocks such as monetary shocks only have temporary impact. The "taste" shock is meant to capture a variety of disturbances that would permanently impact exchange rates but not output. For example, it could represent a shift in preferences towards or away from traded goods, changes in trade policy that may alter the relative demand for traded goods, etc. The properties of the "taste" shock will be analyzed in much more detail later in the paper.

One would expect that commodity prices in the long-run are driven by changes in supply and demand of goods and by shocks directly to the commodities market (ε^{cp}), like an oil price shock. However, monetary shocks should have no reason to leave permanent effects on commodity prices. This underlies the zero restrictions on the fourth row of $C(1)$.

For the consumer price variables in the VAR, real shocks should play a role in their long-run value; however, not all nominal shocks would. Money supply shocks (ε^{ms} and $\varepsilon^{ms\text{-}c}$), defined as adjustments in the nominal interest rate in excess of the Federal Reserve's reaction to changing output and inflation, have a long run impact on prices. On the other hand, money demand shocks (ε^{md} and $\varepsilon^{md\text{-}c}$) may not have the same effect. For example, the monetary authority, in an effort to keep prices stable, may adjust monetary policy (i.e. interest rates) when a money demand shock hits, leaving prices unchanged. This implies that these money demand shocks will not have a long run impact on prices. Again, relative shocks would affect the relative and non-relative variables, but the common shocks would only affect the non-relative variables; hence the zero entries in the fifth and sixth rows of $C(1)$.

Finally, note that because interest rates enter the VAR in levels (i.e. it is stationary), none of the structural shocks should lead to permanent changes in them. However, as interest rates respond quickly to any changes in the economy, all structural shocks are allowed to have short-term impacts on the interest rate variables. The only exception is that the common money demand shock ($\varepsilon^{md\text{-}c}$) is assumed not to have a permanent impact on an accumulation of the relative interest rates. This justifies the remaining zero restriction.

4. Estimation Results for Structural VAR Model

4.1 Benchmark U.S.-U.K. Case

The reduced-form U.S.-U.K. benchmark VAR is estimated using 4 lags for quarterly data.[8] Figure 1 shows the estimated dynamic response of the variable of interest to a one-standard deviation realization of a particular structural shock. The estimates have been suitably transformed to reflect the effect of shocks on the levels of the variables rather than their growth rates. I have omitted the results on relative variables in the VAR for brevity. The impulse-response functions shown in Figure 1 correspond to predictions of macroeconomic theory in general, although the point estimates (solid lines in Figure 1) are not always statistically significant. For example, consider the exchange rate and monetary policy shock. Looking at the fifth row in Figure 1, the relative money supply shock ε^{ms}, which can best be interpreted as a monetary policy shock, is associated with a drop in the 3-month treasury bill rate i_{tbr}. This indicates an expansionary monetary policy shock,[9] which should and does lead to an immediate depreciation of the real exchange rate, and an increase in output and the price level. Focusing on

[8] Standard lag selection criterions select fewer lags [The Akaike Information Criterion (AIC), the Bayesian Information Criterion (BIC), and the Hannan-Quinn Criterion (HQC) suggest 3, 1, and 1 lags respectively]. However, using just one lag suggested by the BIC and HQC leads to non-white-noise like residuals. It is possible that in large size VARs such as the one in this paper, the lag selection criteria penalize additional lags to a greater degree than for smaller dimension VARs. Because four lags were found to fully capture the serial correlation in the data, they are used for the benchmark case and for all other cases considered in this paper.

[9] It is possible that an expansionary monetary policy shock may not lead to a drop in interest rates if a liquidity effect does not dominate. Bernanke and Mihov (1998) have shown that there is no reason to reject the liquidity effect under their VAR framework, and as the sign of the monetary policy shock cannot be identified in any other fashion in this context, I will stick to the conventional assumption.

the impact of shocks on the real exchange rate, the most striking feature in Figure 1 is the impulse-response of real exchange rate q to the taste shock ε^d. The taste shock displays a large and statistically significant effect on the real exchange rate both on impact and beyond. This shock, however, does not appear to be picking up demand-side factors that are related to output; it has essentially zero influence on output both in the short and long-run. Hence I would argue that it is appropriate to label the shock as a "taste" shock rather than a "demand" shock.

A related way to examine the impact of individual shocks on the real exchange rate is to consider the variance decomposition presented in Table 3, which reports the share of the variance of the forecasting error made due to any one structural shock at any given time horizon. Looking at the first row of Table 3, one can easily see that the most important shock, explaining about 70 percent of the variance in the real exchange rate on impact, is the "taste" shock ε^d. No other shock even comes close. The relative monetary policy shock (ε^{ms}) accounts for a mere 2 percent. However, if we consider monetary shocks more generally as the combination of money supply and demand shocks, then their importance grows, accounting for about 28 percent of the variance in the real exchange rate on impact. As the forecast horizon expands, the taste shock becomes even more dominant while the monetary shocks' combined effect declines quickly. After four quarters, the total effect of the monetary shocks is less than half of that on impact. As for the other shocks in the system, neither the supply shocks nor the commodity price shocks have much influence over the real exchange rate at the short or long horizons (the supply shocks are allowed to have permanent effect on the real exchange rate, but empirically they do not appear to be important).

Despite the differences in modeling assumptions and data, the results here bear many similarities to those in Clarida and Gali (1994) and Rogers (1999), both of which used the long-run identification schemes to investigate the effect of monetary policy shocks on the U.S.-U.K. real exchange rate. Demand-type disturbances (including the "taste" shock because it is assumed to have no long-run impact on output) are always very important (accounting for over 95 percent of real exchange rate variance over any horizon in Clarida and Gali, and over 45 percent in

Rogers). Whereas supply disturbances do not play much of a role (proportions of variance that can be attributed to supply are lower than 10 percent for both Clarida and Gali and Rogers regardless of the time horizon). The main differences in our results arise from monetary shocks. Using a simple three variable VAR model, Clarida and Gali found that the maximum impact of monetary shock on the real exchange rate is only about 2 percent. In contrast, Rogers' five-variable VAR model showed that at a maximum, monetary policy shocks (shocks to the monetary base) account for around 15 percent of the variance in the real exchange rate, and this number almost triples if one also considers the other monetary shock that he identified (shocks to the money multiplier).

Overall though, it is a robust finding under long-run identification schemes that monetary policy and supply shocks are not major sources of exchange rate fluctuations. In addition, the effects of monetary shocks on real exchange rate that I reported in Figure 1 and Table 3 are well within the estimated range found in the literature using non-long-run identification schemes. What the results here suggest is that one should focus more on demand side shocks as the main source for fluctuations in the real exchange rate. This will be the topic of Section 5 of the paper.

4.2 Results for Other Countries

Three other countries from the G7 (Canada, Germany, and Japan) are considered in this paper in additional to the benchmark U.S.-U.K. case. Figure 2 displays impulse-responses of U.S. output, real exchange rate, U.S. price level, and U.S. interest rate (3-month treasury bill rate for Canada and the federal funds rate for Germany and Japan) to a one standard deviation monetary policy shock. These impulse-responses are produced from VAR models with the same specification as the benchmark case using four lags. From the analysis in the previous section, we know that the long-run restrictions imposed on the U.S.-U.K. VAR model appear to have identified an expansionary monetary policy shock that lowers the treasury bill rate i_{tbr} on impact, produces a depreciation of the real exchange rate, and leads to a rise in output and prices over time. For the other country pairs, the monetary policy shock does not appear to be as sharply identified. From Figure 2 one can observe some counter-intuitive responses of the price level and exchange rate to a one standard deviation monetary policy shock.

The presence of these "puzzles" is not uncommon and has been discussed at length in the literature. Putting the puzzles aside, we can still assess the relative strengths of each structural shock on the real exchange rate for these countries pairs using variance decomposition, and the results presented in Table 4 through 6 make it clear that the taste shock is still by far the most important shock that contribute to exchange rate variability, perhaps with the exception of Japan in the short-run. The variance decomposition results for Japan in Table 6 show that monetary policy shocks have the strongest effect on impact relative to all the other country pairs, accounting for almost 16 percent of the variance in the real exchange rate on impact, though the most important shock for this particular time horizon is the relative money demand shock, accounting for over 47 percent of the variance in the real exchange rate. Together all the monetary shocks explain the majority of exchange rate variability (about 75 percent) immediately after the shocks hit the economy. These monetary shocks are very persistent as well, even at the 40-quarter horizon, all the monetary shocks combined account for about 10 percent of total variance in the dollar-yen exchange rate.[10] Due to the overwhelming importance of the monetary shocks, the taste shock in the U.S.-Japan case is relatively small on impact, only accounting for about 16 percent of the variance in the exchange rate. But as the effect of monetary shocks die off, the taste shock gains in importance, though in the very long-run the taste shock still accounts for about thirty percentage points less than in the benchmark case.

These non-benchmark country pair results are roughly consistent with what was found in a number of other papers in the literature using a variety of different identification schemes, such as Clarida and Gali (1994), Grilli and Roubini (1995), Eichenbaum and Evans (1995), and Faust

[10] The strong monetary policy shock results for Japan and to some extent Germany are not due to the fact that the federal funds rate was used as the U.S. interest rate variable i in the VAR models (instead of the 3-month treasury bill rate). Robustness checks show that if I replace the federal funds rate with the 3-month treasury bill rate in the German and Japanese VARs, the monetary policy shock comes out slightly weaker and the taste shock slightly stronger. If I replace the 3-month treasury bill rate with the federal funds rate in the VARs for the U.K. and Canada instead, there is practically no difference in the results, in fact, the taste shock actually comes out slightly stronger than what is reported in the tables.

and Rogers (2003). It is rather curious that most of these papers find much stronger monetary policy shock effects on the dollar-yen and dollar-mark exchange rates compared to the dollar-pound or dollar-Canadian dollar exchange rates, but none elaborated on the possible reasons why. A potential explanation could be that the financial systems in Germany and Japan are much more bank-based than Canada or the U.K., which may exacerbate the impact of monetary policy shocks on real and financial macro variables.

5. A Further Investigation of the "Taste" Shock

From the results reported in the previous section, it is clear that the "taste" shock plays the main role in exchange rate fluctuations, both at short and long horizons. However, it is less clear what this taste shock represents. Because the taste shock is identified as a shock that has permanent effects on all the variables in the VAR except output, and there are no other non-monetary demand-type shocks identified in the system, it is very likely that the taste shock captures a variety of demand side disturbances unrelated to money.

In the exchange rate determination literature, besides the usual discussions of supply side factors (productivity and price differentials working through the Balassa-Samuelson effect to influence exchange rate), a variety of demand side factors have also been suggested. Froot and Rogoff (1991), Rogoff (1992), and DeGregorio and Wolf (1994), among many others, have emphasized the importance of government spending shocks in the absence of perfect capital mobility, which are shown to be empirically important in Froot and Rogoff (1991) and Rogers (1999). Also, since exchange rate is an essential element in international trade, factors related to trade may be crucial to exchange rate determination as well, such as terms of trade (Gregorio and Wolf 1994 and Stockman 1980), trade openness and changes in trade policy (see Li 2003 for an excellent summary of theoretical and empirical studies in the area), and the current account (Krugman 1990). Economists have also been interested in the impact of more abstract factors like risk and expectations on exchange rates. Dornbusch (1976) provides a classic model of expectation and exchange rate dynamics. Alvarez, Atkeson and Kehoe (2006) construct a general equilibrium monetary model with endogenous risk variations that is able to reproduce some key features of actual exchange rates. More recent empirical research on exchange rate determination has focused on the microstructures of the foreign exchange or financial market. These studies have suggested that shocks related to information dispersions or foreign exchange order flow (Evans and Lyons 2002) can have a significant impact on the exchange rate. However, as these studies often make use of very high frequency data, the findings that shocks to information dispersion and order flow affect the exchange rate may not be as relevant for longer horizon variations considered here. The taste shock identified in the structural VAR model could be a combination of some or all of the factors mentioned above, so I will now extend my analysis in the previous section to determine the main factor or factors behind the taste shock and hence find the driving force of real exchange rate variability.

Standard theory predicts that a positive demand shock to the U.S. economy leads to a short-run increase in output, a long-run increase in U.S. prices, and a short-run appreciation of the U.S. dollar in real terms. Figures 3 and 4 show the impulse-response functions of the taste shock on the relevant output, price and exchange rate variables for the four countries under investigation. Let us focus our attention on the benchmark U.S.-U.K. results first. Figure 3 shows that U.S. output y does not seem to be affected much by the taste shock, with the impulse-response function hovering around zero. Relative output exhibits more of a response, showing that the shock could be related to a relative demand shock favoring U.K. output and producing a real depreciation of the U.S. dollar as exhibited by the impulse-response of the real exchange rate q. Meanwhile, the price variables show little response to the taste shock. One could imagine that if the taste shock captures traditional demand-like shocks such as government spending shocks or shocks to income and consumption, the reaction of output and prices would be much larger than what is shown in Figure 3.

Looking at the results for the other country pairs in Figure 4, the story is similar to the benchmark case. There are very small responses of output and prices to the taste shock, with the 95 percent confidence interval always including zero. Japan is the only exception to the rule. It is rather peculiar that the relative price variable for the U.S.-Japan pair exhibits such strong and significant reaction to the taste shock despite little movement on the relative output front. This is further evidence that the taste shock is not a typical demand shock (price response with no output effect), and it appears to have differing effects on different country pairs as well.

As mentioned earlier, factors related to risk and expectations could be a driving force behind exchange rates, and shifts in these elements are likely to show up in interest rate variables, especially relative interest rates. Figure 5 illustrates the impulse-responses of the interest rate variables to a one standard deviation taste shock for the four country pairs. Wide confidence intervals covering the zero line are the dominating trait for all the country pairs. This indicates that reactions of the interest rate variables to the taste shock are statistically insignificant. Movements in relative short-term interest rates are slightly larger than for the U.S. short-term rate. The reactions are also slightly larger for the German and Japanese cases.[11]

The only other demand side factors that have not been considered yet are those related to international trade and those that have their roots in the micro structures of the foreign exchange market. Microeconomic factors such as "order flow" have been shown to be important in exchange rate determination (see, for example, Evans and Lyons 2002), hence could be a significant source for the taste shock. However, it is beyond the scope of this paper to investigate market microstructure that operates at high frequencies and requires high frequency data for analysis, therefore, I will concentrate on the potential effects of international trade factors on real exchange rate. Using regression analysis, I examine the relationship between the taste shocks and three variables that have been emphasized by previous theoretical and empirical studies: trade openness, terms of trade, and the current account.

Following related literature in the area of international trade, I consider the ratio of exports plus imports to the gross domestic product, sometimes referred to as "trade intensity," as

[11] As in the monetary policy shock case, robustness checks show that the larger impact of the taste shock on the interest rate variables for Germany and Japan is not due to the fact that I use the federal funds rate instead of the three-month treasury bill rate in the VAR model estimations for these two countries. Indeed, the reactions of the interest rate variables are even larger if I replace the federal funds rate with the treasury bill rate here.

the proxy for trade openness. Trade intensity is one of the most commonly used measures of openness in the literature, where the data needed for the construction of the variable are reliable and readily available for the countries and sample period being considered. Theoretical models point to a real depreciation of the domestic currency after an increase in trade openness. As a country liberalizes its trade, demand for imports increases and demand for non-tradables decreases in response to relative price change. Then, if the Marshall-Lerner condition holds, a real depreciation would be necessary to maintain internal and external balance. However, Calvo and Drazen (1998) have argued that if the trade liberalization is non-credible (or of uncertain duration) then it is potentially possible to see real appreciation instead. There is little empirical evidence on this subject, though Li (2003) shows that using event studies, *ceteris paribus*, the real exchange rate depreciates after a country's most recent episode of trade liberalization.

The "terms of trade" variable is measured as the ratio of export prices to import prices. Intuitively, this is a ratio that quantifies a country's welfare. An increase (or improvement) in the terms of trade implies the home country gets more units of imported good for each unit of good it exports. This variable could affect the real exchange rate through both income and substitution effects. An increase in the terms of trade, for example, would mean a boost to real income and, therefore, a rise in demand and hence the relative price for non-tradables. The general price level would increase as a result, so this income effect eventually leads to appreciation of the real exchange rate. The substitution effect is less straightforward. Assuming that non-tradables and tradables are substitutes, an improvement of the terms of trade would cause the non-tradable prices to increase relative to imports, but decrease relative to exports, leaving ambiguous the change in the relative price of non-tradables to tradables as a whole. Hence, if income effect dominates, an improvement in, say, the U.S. terms of trade relative to a foreign country would mean a real appreciation of the U.S. dollar relative to the foreign country's currency. De Gregorio and Wolf (1994) find that for their sample of OECD countries, an improvement in the terms of trade does lead to a real appreciation.

The final trade variable under consideration here is the current account, which enters the analysis as a ratio (current account to GDP). Theoretically speaking, it is rather natural to see a link between real exchange rates and current account, as both exports and imports are affected when the real exchange rate changes. However, the direction of causation may go the other way also. It is well documented that sustained current account deficits are associated with long-run real exchange rate depreciation. Wright and Gagnon (2006) presented results that show the current account to GDP ratio has a modest but statistically significant effect on the estimated probability of a large depreciation; and Krugman (1990) argues that current account changes lead to transfers of wealth across countries, and as the spending pattern differs across home and foreign residents, it is likely to induce significant real exchange rate changes.

Table 7 presents the regression results. The counterfactual real exchange rate with only the taste shock on is the dependent variable.[12] The explanatory variables include the three trade measures mentioned previously, each entering the regression as a country specific trade variable. Since the dependent variable is non-stationary by construction, and most of the trade variables are also non-stationary according to standard unit root tests,[13] a percentage change specification is necessary to avoid the spurious regression problem.[14]

In general, results presented in Table 7 show that for all country pairs, at least one of the trade measures show up as a statistically significant explanatory variable for the hypothetical real exchange rate with only the taste shock on, hence could be potential sources of the taste shock. Specifically, for the benchmark U.S.-U.K. case, U.S. and U.K. terms of trade as well as U.K. trade openness appear to be significant. The coefficients on these variables have the expected

[12] The hypothetical series reflects the effect of the accumulation of taste shocks on the real exchange rate.

[13] Unit root test results available upon request.

[14] Another solution to the spurious regression problem is to check for potential cointegrating relationships between the dependent and explanatory variables, and if there is cointegration, one can apply the dynamic ordinary least squares (DOLS) specification with Newey-West standard error correction. However, it was not possible for me to reject the null of no cointegration at the 5 percent level using the Z_t-test developed by Phillips (1987) for any of the country pairs under investigation here. The critical values used for the Z_t-test are produced from the FORTRAN programs provided by MacKinnon (1996).

signs except for U.S. terms of trade. A positive coefficient for the U.S. terms of trade indicates that as the U.S. terms of trade improves, we observe a depreciation of the U.S. real exchange rate. This is somewhat counter intuitive, though if the substitution effect dominates in the U.S. economy, we could potentially see a real depreciation despite the improvement in terms of trade. For Canada, the Canadian trade openness and terms of trade measures both show up as significant explanatory variables for the hypothetical exchange rate. Both of these variables have the expected signs. For Germany, only one trade variable comes up significant, the German terms of trade. The coefficient is highly significant though and has the expected sign. Finally, for Japan, U.S. trade openness appears to be highly significant while the Japanese terms of trade is significant at the 10% level. Again, both of these variables have the expected signs.

On the whole, the regression analysis appears to have pinned down a potential source of the taste shock, namely the international trade sector of the economy.[15] Foreign terms of trade measures seem particularly important, showing up as a significant explanatory variable in all country pairs. Current account, on the other hand, does not seem to matter. One can interpret the results here as being fairly consistent with the earlier findings as shocks in trading terms and conditions are frequent, but may not have immediate or prominent effects on output or interest rate differentials.

6. Conclusion

The results in this paper provide evidence against the prevalent idea that nominal factors such as monetary policy shocks have been the dominant source of real exchange rate fluctuations during the post Bretton Woods era. This is not to say that nominal shocks are unimportant. In

[15] As a robustness check, I included the percentage change in government expenditure (relative to GDP) as an explanatory variable in some specifications of the regression analysis. If the taste shock is capturing fiscal type shocks to the economy, this variable should be statistically significant in the regressions. For all country pairs, neither the U.S. nor the foreign government expenditure measures show up as statistically significant even at the 10% level. The only exception to the rule is for the U.S.-German case, where the U.S. government expenditure variable is a significant explanatory variable for the hypothetical real exchange rate at the 5% level. Because of the general result that government expenditure does not affect the taste shock, Table 7 reports the specification with only the trade variables.

many cases they are, especially for the dollar-yen exchange rate. However, the dominant factor dictating exchange rate movements in my results appear to be non-monetary in nature. For the benchmark U.S.-U.K. case, at a maximum, total monetary shocks (money supply plus money demand shocks) account for less than thirty percent of the forecasted error variance in the dollar-pound real exchange rate. The number is even smaller for the U.S.-Canada and U.S.-Germany cases. Japan is the only exception where total monetary shocks account for over seventy-five percent of forecasted error variance in the real exchange rate at the peak horizon (on impact). In contrast, a real factor, which I termed a "taste" shock, appears to be the main source of real exchange rate variability over the short, medium, and long horizons. This taste shock accounts for over fifty percent of real exchange rate forecast error variance over all horizons for every country in the sample except for Japan. Even in the Japanese case, nominal shocks are the most important only in the first four quarters, after which the taste shock takes over as the predominant factor.

My results complement those in Clarida and Gali (1994), who also found real shocks to be important for exchange rate fluctuations. However, the model in this paper offers a refinement over theirs in the sense that it contains a richer information set and hence can distinguish between a larger number of possible shocks that may drive the exchange rate. The results presented are also along the lines of empirical papers that make use of other identification strategies, such as Eichenbaum and Evans (1995), who have shown that monetary policy shocks are not the main driving force behind exchange rate fluctuations.

These findings imply that policy makers and economists alike will be misguided if all they worry about is the effects of monetary shocks on exchange rates. Instead, more emphasis should be placed on real shocks such as shocks to the trading sector of the economy. I have shown that the taste shock, which plays the essential role in dictating the movements in the exchange rate, appears to be associated with important trade measures such as trade openness, terms of trade, and the current account. Perhaps surprisingly, the taste shock does not seem to be related to macroeconomic shocks such as shocks to government spending.

The model here could be further extended in subsequent research to consider a wider range of short-run and long-run restrictions (including combinations of both) that would allow other potentially important shocks to be identified individually (for example government spending shocks). Also, it would be interesting to see if the results are robust to using the sign/shape restriction method for identification, which typically has shown monetary shocks to be quite important for exchange rates. Finally, a larger set of real exchange rates could be considered to see if the trade factors found important here for the bilateral exchange rates between the U.S. and the four industrialized countries will hold up for bilateral exchange rates between the U.S. and developing or less developed countries.

References

Alvarez, F., Atkeson, A., and Kehoe, P., 2006. Time-Varying Risk, Interest Rates, and Exchange Rates in General Equilibrium. Federal Reserve Bank of Minneapolis, Staff Report: 371.

Bai, J., and Perron P., 1998. Estimating and Testing Linear Models with Multiple Structural Changes. Econometrica 66-1, 47-78.

Bai, J., and Perron, P., 2003. Computation and Analysis of Multiple Structural Change Models. Journal of Applied Econometrics 18-1, 1-22.

Bernanke, B., and Blinder, A., 1992. The Federal Funds Rate and the Channels of Monetary Transmission. American Economic Review 82-4, 901-921.

Bernanke, B., and Mihov, I., 1998. Measuring Monetary Policy. Quarterly Journal of Economics, 113-3, 869-902.

Blanchard, O., and Quah, D., 1989. The Dynamic Effects of Aggregate Demand and Supply Disturbances. American Economic Review 79-4, 655-673.

Calvo, G., and Drazen A., 1998. Uncertain Duration of Reform: Dynamic Implications. Macroeconomic Dynamics 2-4, 443-455.

Canova, F., and De Nicoló, G., 2002. Monetary Disturbances Matter for Business Fluctuations in the G-7. Journal of Monetary Economics 49-6, 1131-1159.

Caporale, T., and Grier K., 2000. Political Regime Change and the Real Interest Rate. Journal of Money, Credit, and Banking 32-3, 320-334.

Choi, W.Y., 2005. Real Exchange Rates, International Trade, and Macroeconomic Fundamentals. Ph. D. Dissertation, Department of Economics, Stanford University.

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

De Gregorio, J., and Wolf, H., 1994. Terms of Trade, Productivity and the Real Exchange Rate. NBER Working Paper: 4807.

Dornbusch, R., 1976. Expectations and Exchange Rate Dynamics. Journal of Political Economy 84-6, 1161-1176.

Edison, H., 1987. Purchasing Power Parity in the Long Run: A Test of the Dollar/Pound Exchange Rate (1890-1978). Journal of Money, Credit, and Banking, 19-3, 376-387.

Eichenbaum, M., and Evans, C., 1995. Some Empirical Evidence on the Effects of Shocks to Monetary Policy on Exchange Rates. Quarterly Journal of Economics 110, 975-1009.

Evans, M. D. D., and Lyons R., 2002. Order Flow and Exchange Rate Dynamics. Journal of Political Economy 110-1, 170-180.

Farrant, K., and Peersman, G., 2006. Is the Exchange Rate a Shock Absorber or a Source of Shocks? New Empirical Evidence. Journal of Money, Credit, and Banking, 38-4, 939-961.

Faust, J, 1998. The Robustness of Identified VAR Conclusions About Money. Carnegie-Rochester Conference Series on Public Policy 49, 207-244.

Faust, J, and Leeper, E., 1997. When Do Long Run Identifying Restrictions Give Reliable Results? Journal of Business and Economic Statistics, 15-3, 345-353.

Faust, J., and Rogers, J. H., 2003. Monetary Policy's Role in Exchange Rate Behavior. Journal of Monetary Economics 50, 1403-1424.

Froot, K. A., and Rogoff, K., 1991. The EMS, the EMU, and the transition to a common currency. NBER Macroeconomics Annual, Cambridge and London: MIT Press, pp. 269-317.

Grilli, V., and Roubini, N., 1995. Liquidity and Exchange Rates: Puzzling Evidence from the G-7 Countries. New York University, Leonard N. Stern School of Business, Department of Economics, Working Paper.

Kilian L., 1998. Confidence Interval for Impulse Responses under Departures from Normality. Econometric Reviews 17-1, 1-29.

Kilian L., and Chang, P.L., 2000. How Accurate are Confidence Intervals for Impulse Responses in Large VAR Models? Economics Letters 69-3, 299-307.

Kim, S., and Roubini, N., 2000. Exchange Rate Anomalies in the Industrial Countries: A Solution with a Structural VAR Approach. Journal of Monetary Economics 45, 561-586.

Krugman, P. R., 1990. Equilibrium Exchange Rates. International Policy Coordination and Exchange Rate Fluctuations, University of Chicago Press, Chicago, pp. 159-187.

Kwiatkowski, D., Phillips, P. C. B., Schmidt, P., and Shin, Y., 1992. Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root: How Sure are We that Economic Time Series Have a Unit Root? Journal of Econometrics 54, 159-178.

Levin, A., and Piger, J. M., 2003. Is Inflation Persistence Intrinsic in Industrial Economies? The Federal Reserve Bank of St. Louis Working Paper Series, working paper 2002-023E.

Li, X., 2003. Trade Liberalization and Real Exchange Rate Movement. IMF Working Paper.

Lütkepohl, H., 1990. Asymptotic Distributions of Impulse Response Functions and Forecast Error Variance Decompositions of Vector Autoregressive Models. Review of Economics and Statistics 72-1, 116-125.

MacKinnon, J. G., 1996. Numerical Distribution Functions for Unit Root and Cointegration Tests. Journal of Applied Econometrics 11, 601-618.

Mussa, M., 1986. Nominal Exchange Rate Regimes and the Behavior of Real Exchange Rates: Evidence and Implications. Carnegie-Rochester Conference Series on Public Policy 25, 117-213.

Phillips, P. C. B., 1987. Time Series Regression with a Unit Root. Econometrica 55-2, 277-302.

Rapach, D. E., and Wohar, M. E., 2005. Regime Changes in International Real Interest Rates: Are They a Monetary Phenomenon? Journal of Money, Credit, and Banking 37-5, 887-906.

Rogers, J. H., 1999. Monetary Shocks and Real Exchange Rates. Journal of International Economics 49-2, 269-288.

Rogoff, K., 1992. Traded Goods Consumption Smoothing and the Random Walk Behavior of the Real Exchange Rate. Bank of Japan Monetary and Economic Studies 10-2, 1-29.

Rogoff, K., 1996. The Purchasing Power Parity Puzzle. Journal of Economic Literature 34-2, 647-668.

Runkle, D. E., 1987. Vector Autoregressions and Reality. Journal of Business and Economic Statistics 5-4, 437-442.

Sims, C., and Zha, T., 1999. Error Bands for Impulse Responses. Econometrica 67-5, 1113-1155.

Stockman, A. C., 1980. A Theory of Exchange Rate Determination. Journal of Political Economy 88-4, 673-697.

Stockman, A. C., 1987. The Equilibrium Approach to Exchange Rates. Economic Review Federal Reserve Bank of Richmond 73-2, 12-30.

Uhlig, H., 2005. What Are the Effects of Monetary Policy on Output? Results From An Agnostic Identification Procedure. Journal of Monetary Economics 52-2, 381-419.

Wright, J. H., and Gagnon, J. E., 2006. Predicting Sharp Depreciations in Industrial Country Exchange Rates, Board of Governors of the Federal Reserve System, International Finance Discussion Papers, No. 881.

TABLE 1

DATA DETAILS AND SOURCES

Variable	Detail	Source
$y - y^*$	Relative output = ln(U.S. real GDP)[1] – ln(foreign real GDP)[2]	U.S. data from FRED[3] Foreign data from IFS[4]
y	U.S. output = U.S. real GDP	FRED
q	Real exchange rate = ln(nominal exchange rate)[5] + ln(foreign CPI)[6] – ln(U.S. CPI)[7]	U.S. data from FRED Exchange rate and foreign data from IFS
p^c	Real Intermediate materials price = ln(producer price index: intermediate materials)[8] – ln(U.S. CPI)	FRED
$p - p^*$	Relative prices = ln(U.S. CPI) – ln(foreign CPI)	U.S. data from FRED Foreign data from IFS
p	U.S. CPI = U.S. consumer price index	FRED
$i - i^*$	Relative short-term interest rate = U.S. 3-month treasury bill rate(or federal funds rate) – foreign short-term treasury bill rate (or money market rate)[9]	U.S. data from FRED Foreign data from IFS
i_{tbr}	U.S. 3-month treasury bill rate[10]	FRED
i_{ffr}	U.S. Federal Funds Rate[11]	FRED

1. Real gross domestic product (GDPC96), seasonally adjusted annual rate in billions of chained 2000 dollars.
2. Canada: GDP volume 1997 ref. chained (15699B.RXF...), Germany: GDP volume 2000 = 100 (13499BVRZF...), Japan: GDP volume 2000 = 100 (15899BVRZF...), U.K.: GDP volume 2003 ref. chained (11299B.RXF...). Chained volume is chosen whenever it is available for the sample period under consideration.
3. St. Louis Fed FRED database.
4. International Monetary Fund's International Financial Statistics database.
5. Nominal exchange rate defined as amount of U.S. dollars needed to purchase one unit of foreign currency. For nominal exchange rate between U.S. and Germany, it is the amount of U.S. dollars needed to purchase one German Mark. Since the DM is defunct after Dec. 1998, I patched on the Euro movement to the data after that date.
6. Canada: CPI all cities pop over 30,000 (15664...ZF...), Germany: GDP deflator 2000 = 100 (13499BIRZF...), Japan: CPI all Japan 485 items (15864...ZF...), U.K.: CPI all items (11264...ZF...). Had to use GDP deflator for Germany as CPI data is incomplete over sample period.
7. CPI For All Urban Consumers: All Items (SA) 1982-84 = 100 (CPIAUCSL).
8. Producer Price Index: Intermediate Materials: Supplies & Components (PPIITM), seasonally adjusted 1982 = 100.
9. Treasury bill rates are used for Canada and the U.K.; Germany and Japan use the money market rates. Decisions on which rate to use are based on availability of data during sample period.
10. 3-month treasury bill rate (TB3MS).
11. Effective Federal Funds Rate (FEDFUNDS).

Note: Data that are originally in monthly form (everything except for GDP) are converted to quarterly by taking the quarter's last monthly observation.

TABLE 2

LIST OF STRUCTURAL BREAK DATES

Country	Variable	Structural breaks
U.S.	Δp	1981Q2
	$\Delta ydef$ (ΔGDP deflator)	1980Q3
	i_{tbr}	1979Q2, 1981Q2, 1982Q2, 1985Q2
	i_{ffr}	1979Q2, 1981Q2, 1982Q2, 1985Q2
Canada	$\Delta(p - p^*)$	1981Q2, 1990Q4
	$i - i^*$	1979Q2, 1981Q2, 1982Q2, 1983Q4, 1985Q2, 1992Q2
Germany	$\Delta(p - p^*)$	1972Q4, 1980Q3, 1990Q4
	$i - i^*$	1979Q2, 1981Q2, 1982Q2, 1984Q2, 1985Q2, 1992Q3
Japan	$\Delta(p - p^*)$	1977Q2, 1981Q2, 1997Q1
	$i - i^*$	1979Q2, 1981Q2, 1982Q2, 1985Q2
U.K.	$\Delta(p - p^*)$	1974Q1, 1975Q1, 1981Q2, 1990Q1
	$i - i^*$	1979Q2, 1981Q2, 1982Q2, 1983Q4, 1985Q2, 1992Q3

Note: All calculations performed using Gauss. The Gauss program assumes that the series being tested has a simple AR(1) structure and has one break in the mean. It then calculates the F-statistic for all potential break dates over the sample period with trimming factor = 0.15. The date with the largest F-stat is then tested for significance using the critical values produced by Andrews (1993, 2003). If it is a valid break date, then the original sample is split into two on the break date, and the same procedure is repeated for the first sub-sample and second sub-sample to search for more break dates. This process continues until the sub-samples are too small to be used or when no more valid break dates are found. I start the structural break search with U.S. inflation (Δp), then apply the same break date found in Δp on relative inflation $\Delta(p - p^*)$ and the U.S. 3-month treasury bill rate (i_{tbr}). The search procedure is repeated to find additional breaks in $\Delta(p - p^*)$ and i_{tbr}. Structural breaks in i_{ffr} are simply assumed to be the same as the structural breaks found in i_{tbr}. The reason for this assumption is that the breaks found for i_{ffr} through the search procedure did not make the modified i_{ffr} series (with structural breaks imposed) stationary, whereas when we imposed the same breaks found in i_{tbr} on i_{ffr}, the modified series becomes stationary. This may seem a bit ad-hoc, but intuitively one would think that structural breaks on two closely related interest rates could and should be identical. Finally, I impose the break dates found in i_{tbr} on the relative short-term interest rates ($i - i^*$) then search for more breaks.

TABLE 3

VARIANCE DECOMPOSITION RESULTS FOR U.S.-U.K. BILATERAL REAL EXCHANGE RATE IN LEVELS

Time horizon with 95% confidence bands	ε^{s} (%)	$\varepsilon^{s\text{-}c}$ (%)	ε^{d} (%)	ε^{cp} (%)	ε^{ms} (%)	$\varepsilon^{ms\text{-}c}$ (%)	ε^{md} (%)	$\varepsilon^{md\text{-}c}$ (%)	Total monetary shocks
(1) 0 quarter	0.84	0.68	69.57	0.34	2.18	0.02	15.06	10.8	28.06
Lower band	(0.01)	(0.01)	(15.38)	(0.00)	(0.01)	(0.00)	(0.03)	(0.05)	
Upper band	(28.13)	(32.12)	(80.97)	(29.71)	(25.56)	(16.65)	(34.80)	(34.74)	
(2) 4 quarters	0.26	1.00	82.98	2.11	2.75	1.07	3.17	6.65	13.64
Lower band	(0.31)	(0.38)	(27.27)	(0.20)	(0.22)	(0.16)	(0.36)	(0.25)	
Upper band	(41.65)	(26.39)	(87.21)	(27.06)	(13.93)	(13.16)	(9.94)	(17.67)	
(3) 8 quarters	0.25	0.74	88.00	1.47	3.11	0.67	1.78	3.97	9.53
Lower band	(0.38)	(0.41)	(31.32)	(0.25)	(0.22)	(0.16)	(0.29)	(0.23)	
Upper band	(44.14)	(30.44)	(90.02)	(22.01)	(11.58)	(8.90)	(5.94)	(11.02)	
(4) 12 quarters	0.23	1.45	90.17	1.00	2.51	2.51	0.46	1.40	6.88
Lower band	(0.38)	(0.42)	(32.31)	(0.19)	(0.19)	(0.14)	(0.22)	(0.18)	
Upper band	(47.08)	(34.06)	(91.57)	(17.66)	(8.86)	(5.80)	(4.46)	(7.20)	
(5) 16 quarters	0.20	1.97	91.69	0.75	1.88	0.35	1.11	2.07	5.41
Lower band	(0.32)	(0.37)	(32.18)	(0.15)	(0.15)	(0.11)	(0.17)	(0.16)	
Upper band	(49.44)	(37.54)	(92.42)	(14.61)	(6.45)	(4.30)	(3.27)	(4.93)	
(6) 20 quarters	0.18	2.25	92.74	0.59	1.47	0.27	0.88	1.62	4.24
Lower band	(0.34)	(0.33)	(32.13)	(0.12)	(0.12)	(0.08)	(0.13)	(0.13)	
Upper band	(50.21)	(39.67)	(93.42)	(12.03)	(5.01)	(3.32)	(2.52)	(3.65)	
(7) 40 quarters	0.09	3.12	94.46	0.28	0.71	0.13	0.42	0.78	2.04
Lower band	(0.29)	(0.23)	(29.59)	(0.06)	(0.05)	(0.04)	(0.06)	(0.06)	
Upper band	(54.32)	(44.70)	(95.71)	(5.83)	(2.23)	(1.52)	(1.15)	(1.50)	

Note: Benchmark specification period is from 1970Q2 to 2006Q1 with 4 lags. Confidence intervals are obtained using the Kilian (1998) bias-corrected bootstrap methods. The lower band indicates the 2.5th percentile of the bootstrapped distribution. Similarly, the upper band indicates the 97.5th percentile. The estimated variance decomposition numbers in this table sometimes fall outside the confidence interval bands. This can be explained using a simple example. Suppose the true proportion of the forecast error variance of the real exchange rate explained by a particular structural shock is zero. Then confidence bands generated from random drawings will most likely not include zero (since it is bounded below by zero). Hence if the estimated proportion is really close to 0% or 100%, then it is likely to fall outside the 95% confidence interval constructed. The last column (total monetary shocks) is constructed by simply summing together the proportion of variances due to ε^{ms} , $\varepsilon^{ms\text{-}c}$, ε^{md} , and $\varepsilon^{md\text{-}c}$.

TABLE 4

VARIANCE DECOMPOSITION RESULTS FOR U.S.-CANADA BILATERAL REAL EXCHANGE RATE IN LEVELS

Time horizon with 95% confidence bands	ε^{s} (%)	$\varepsilon^{s\text{-}c}$ (%)	ε^{d} (%)	ε^{cp} (%)	ε^{ms} (%)	$\varepsilon^{ms\text{-}c}$ (%)	ε^{md} (%)	$\varepsilon^{md\text{-}c}$ (%)	Total monetary shocks
(1) 0 quarter	1.73	0.83	87.48	0.09	3.73	2.65	0.46	3.03	9.87
Lower band	(0.01)	(0.01)	(17.81)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	
Upper band	(29.27)	(31.16)	(87.75)	(21.68)	(34.01)	(31.28)	(26.42)	(26.43)	
(2) 4 quarters	1.94	3.21	87.19	0.15	1.80	0.49	0.57	4.64	7.50
Lower band	(0.19)	(0.22)	(24.42)	(0.15)	(0.19)	(0.19)	(0.23)	(0.22)	
Upper band	(29.79)	(37.42)	(88.39)	(17.78)	(21.72)	(16.35)	(20.19)	(20.01)	
(3) 8 quarters	3.89	4.60	87.42	0.11	0.94	0.28	0.39	2.37	3.98
Lower band	(0.22)	(0.26)	(28.30)	(0.14)	(0.14)	(0.16)	(0.18)	(0.17)	
Upper band	(36.06)	(42.98)	(90.66)	(14.83)	(14.87)	(12.47)	(14.62)	(11.33)	
(4) 12 quarters	4.91	5.47	86.99	0.08	0.62	0.19	0.25	1.49	2.55
Lower band	(0.25)	(0.28)	(29.08)	(0.10)	(0.11)	(0.14)	(0.14)	(0.12)	
Upper band	(41.00)	(47.25)	(92.21)	(11.22)	(10.82)	(9.39)	(10.40)	(7.25)	
(5) 16 quarters	5.39	5.80	86.88	0.06	0.45	0.14	0.18	1.08	1.85
Lower band	(0.25)	(0.30)	(29.72)	(0.08)	(0.09)	(0.12)	(0.12)	(0.09)	
Upper band	(43.92)	(50.24)	(93.25)	(9.21)	(8.27)	(7.19)	(7.50)	(5.12)	
(6) 20 quarters	5.71	5.91	86.87	0.05	0.36	0.12	0.14	0.85	1.47
Lower band	(0.23)	(0.33)	(30.21)	(0.07)	(0.07)	(0.10)	(0.10)	(0.07)	
Upper band	(45.60)	(52.77)	(93.76)	(7.33)	(6.82)	(5.75)	(5.78)	(3.95)	
(7) 40 quarters	6.51	5.95	86.82	0.02	0.17	0.06	0.07	0.41	0.71
Lower band	(0.19)	(0.34)	(29.36)	(0.03)	(0.03)	(0.05)	(0.05)	(0.03)	
Upper band	(50.82)	(55.63)	(95.68)	(3.37)	(3.17)	(2.58)	(2.54)	(1.75)	

Note: Sample period for Canada is from 1970Q2 to 2006Q2 with 4 lags. Confidence intervals are obtained using the Kilian (1998) bias-corrected bootstrap methods. The lower band indicates the 2.5th percentile of the bootstrapped distribution. Similarly, the upper band indicates the 97.5th percentile. The estimated variance decomposition numbers in this table sometimes fall outside the confidence interval bands. This can be explained using a simple example. Suppose the true proportion of the forecast error variance of the real exchange rate explained by a particular structural shock is zero. Then confidence bands generated from random drawings will most likely not include zero (since it is bounded below by zero). Hence if the estimated proportion is really close to 0% or 100%, then it is likely to fall outside the 95% confidence interval constructed. The last column (total monetary shocks) is constructed by simply summing together the proportion of variances due to ε^{ms}, $\varepsilon^{ms\text{-}c}$, ε^{md}, and $\varepsilon^{md\text{-}c}$.

TABLE 5

VARIANCE DECOMPOSITION RESULTS FOR U.S.-GERMANY BILATERAL REAL EXCHANGE RATE IN LEVELS

Time horizon with 95% confidence bands	ε^{s} (%)	$\varepsilon^{s\text{-}c}$ (%)	ε^{d} (%)	ε^{cp} (%)	ε^{ms} (%)	$\varepsilon^{ms\text{-}c}$ (%)	ε^{md} (%)	$\varepsilon^{md\text{-}c}$ (%)	Total monetary shocks
(1) 0 quarter	0.40	0.81	76.34	8.75	10.43	2.37	0.18	0.72	13.70
Lower band	(0.01)	(0.19)	(1.03)	(0.01)	(0.13)	(0.01)	(0.00)	(0.00)	
Upper band	(36.93)	(59.97)	(69.75)	(25.72)	(58.57)	(37.09)	(18.92)	(11.67)	
(2) 4 quarters	0.90	0.70	85.43	8.62	3.05	0.53	0.30	0.47	4.35
Lower band	(0.54)	(0.97)	(6.93)	(0.36)	(1.06)	(0.22)	(0.17)	(0.15)	
Upper band	(40.21)	(58.64)	(77.18)	(25.38)	(32.29)	(24.08)	(10.33)	(6.00)	
(3) 8 quarters	1.26	2.64	87.26	5.76	1.85	0.24	0.42	0.56	3.07
Lower band	(0.63)	(1.17)	(17.40)	(0.31)	(0.67)	(0.21)	(0.15)	(0.13)	
Upper band	(40.30)	(51.35)	(83.53)	(20.29)	(24.30)	(17.60)	(7.39)	(4.37)	
(4) 12 quarters	1.94	4.54	87.06	3.85	1.36	0.20	0.49	0.56	2.61
Lower band	(0.56)	(1.05)	(23.82)	(0.24)	(0.46)	(0.19)	(0.10)	(0.09)	
Upper band	(41.52)	(46.19)	(86.69)	(15.93)	(18.39)	(14.54)	(5.61)	(3.09)	
(5) 16 quarters	2.98	5.35	86.74	2.85	1.02	0.19	0.43	0.44	2.08
Lower band	(0.57)	(0.92)	(27.10)	(0.20)	(0.35)	(0.17)	(0.08)	(0.07)	
Upper band	(42.94)	(45.11)	(89.01)	(12.63)	(13.76)	(11.98)	(4.27)	(2.23)	
(6) 20 quarters	4.11	5.79	86.14	2.26	0.82	0.16	0.36	0.35	1.69
Lower band	(0.59)	(0.88)	(30.37)	(0.18)	(0.29)	(0.15)	(0.06)	(0.96)	
Upper band	(43.89)	(44.51)	(89.99)	(10.39)	(10.93)	(9.84)	(3.44)	(1.81)	
(7) 40 quarters	6.86	7.04	84.14	1.11	0.41	0.08	0.18	0.17	0.84
Lower band	(0.63)	(0.71)	(32.46)	(0.09)	(0.14)	(0.08)	(0.03)	(0.03)	
Upper band	(48.67)	(47.11)	(93.57)	(5.43)	(5.42)	(5.25)	(1.68)	(0.85)	

Note: Sample period for Germany is from 1970Q2 to 2005Q4 with 4 lags. Confidence intervals are obtained using the Kilian (1998) bias-corrected bootstrap methods. The lower band indicates the 2.5th percentile of the bootstrapped distribution. Similarly, the upper band indicates the 97.5th percentile. The estimated variance decomposition numbers in this table sometimes fall outside the confidence interval bands. This can be explained using a simple example. Suppose the true proportion of the forecast error variance of the real exchange rate explained by a particular structural shock is zero. Then confidence bands generated from random drawings will most likely not include zero (since it is bounded below by zero). Hence if the estimated proportion is really close to 0% or 100%, then it is likely to fall outside the 95% confidence interval constructed. The last column (total monetary shocks) is constructed by simply summing together the proportion of variances due to ε^{ms}, $\varepsilon^{ms\text{-}c}$, ε^{md}, and $\varepsilon^{md\text{-}c}$.

TABLE 6

VARIANCE DECOMPOSITION RESULTS FOR U.S.-JAPAN BILATERAL REAL EXCHANGE RATE IN LEVELS

Time horizon with 95% confidence bands	ε^{s} (%)	$\varepsilon^{s\text{-}c}$ (%)	ε^{d} (%)	ε^{cp} (%)	ε^{ms} (%)	$\varepsilon^{ms\text{-}c}$ (%)	ε^{md} (%)	$\varepsilon^{md\text{-}c}$ (%)	Total monetary shocks
(1) 0 quarter	1.71	0.47	15.55	6.90	15.63	12.33	47.40	0.01	75.37
Lower band	(0.19)	(0.02)	(0.01)	(0.04)	(0.00)	(0.19)	(4.33)	(0.00)	
Upper band	(63.96)	(43.28)	(29.80)	(40.99)	(19.31)	(38.36)	(52.98)	(9.03)	
(2) 4 quarters	2.13	0.30	45.11	4.60	17.41	6.27	22.07	2.12	47.87
Lower band	(1.16)	(0.56)	(8.28)	(0.34)	(0.24)	(0.36)	(1.17)	(0.12)	
Upper band	(56.63)	(46.55)	(71.68)	(31.14)	(22.46)	(17.05)	(20.13)	(5.26)	
(3) 8 quarters	1.24	2.41	51.13	2.76	16.71	3.89	14.91	6.94	42.45
Lower band	(1.01)	(0.56)	(11.50)	(0.34)	(0.22)	(0.36)	(0.64)	(0.13)	
Upper band	(58.75)	(51.82)	(79.06)	(24.34)	(19.66)	(8.83)	(12.25)	(4.56)	
(4) 12 quarters	1.32	4.18	55.12	2.83	15.46	2.88	10.79	7.42	36.55
Lower band	(0.91)	(0.52)	(13.85)	(0.38)	(0.19)	(0.31)	(0.40)	(0.10)	
Upper band	(61.22)	(55.13)	(82.37)	(19.17)	(17.79)	(6.10)	(7.79)	(3.43)	
(5) 16 quarters	2.50	5.16	59.83	2.97	12.99	2.39	8.09	6.07	29.54
Lower band	(0.81)	(0.46)	(14.85)	(0.38)	(0.14)	(0.25)	(0.26)	(0.08)	
Upper band	(63.88)	(56.84)	(84.84)	(17.03)	(15.37)	(5.10)	(5.13)	(2.42)	
(6) 20 quarters	4.48	6.55	62.94	2.53	10.46	1.93	6.29	4.83	23.51
Lower band	(1.04)	(0.40)	(14.50)	(0.30)	(0.12)	(0.21)	(0.19)	(0.06)	
Upper band	(68.40)	(58.93)	(86.22)	(14.56)	(12.03)	(4.23)	(3.80)	(1.82)	
(7) 40 quarters	12.13	12.54	64.14	1.10	4.52	0.82	2.65	2.09	10.08
Lower band	(1.16)	(0.41)	(9.80)	(0.12)	(0.05)	(0.08)	(0.07)	(0.03)	
Upper band	(80.77)	(61.98)	(88.11)	(7.03)	(4.96)	(1.66)	(1.44)	(0.63)	

Note: Sample period for Japan is from 1970Q2 to 2006Q1 with 4 lags. Confidence intervals are obtained using the Kilian (1998) bias-corrected bootstrap methods. The lower band indicates the 2.5th percentile of the bootstrapped distribution. Similarly, the upper band indicates the 97.5th percentile. The estimated variance decomposition numbers in this table sometimes fall outside the confidence interval bands. This can be explained using a simple example. Suppose the true proportion of the forecast error variance of the real exchange rate explained by a particular structural shock is zero. Then confidence bands generated from random drawings will most likely not include zero (since it is bounded below by zero). Hence if the estimated proportion is really close to 0% or 100%, then it is likely to fall outside the 95% confidence interval constructed. The last column (total monetary shocks) is constructed by simply summing together the proportion of variances due to ε^{ms}, $\varepsilon^{ms\text{-}c}$, ε^{md}, and $\varepsilon^{md\text{-}c}$.

TABLE 7

REGRESSION RESULTS FOR HYPOTHETICAL REAL EXCHANGE RATE WITH ONLY TASTE SHOCK ON

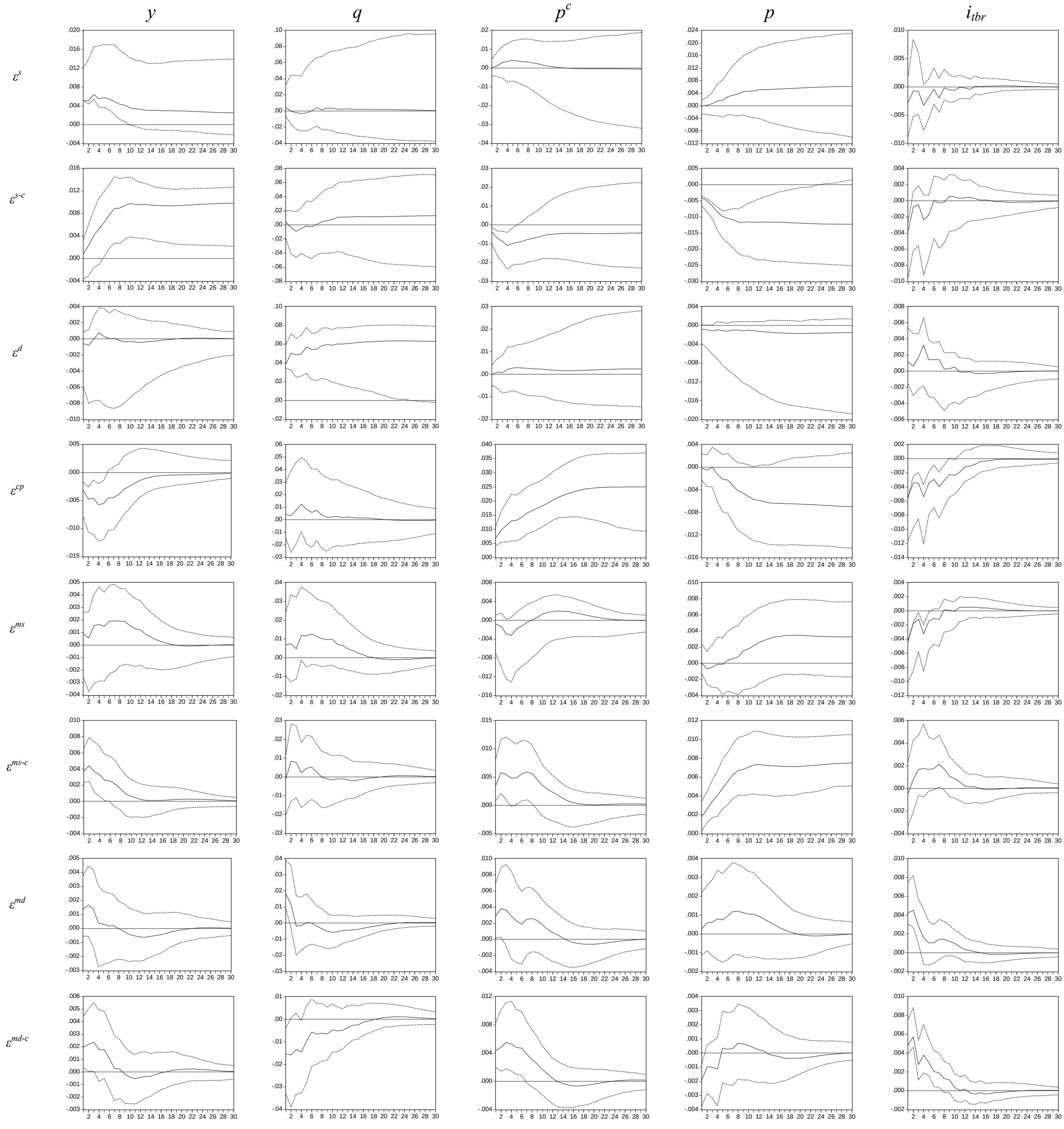
Country	C	%Δ of trade openness (U.S.)	%Δ of terms of trade (U.S.)	%Δ of current account (U.S.)	%Δ of trade openness (foreign)	%Δ of terms of trade (foreign)	%Δ of current account (foreign)
U.K.	0.0322	0.1747	0.1778*	-0.0830	-0.1897*	0.2582**	-0.1248
	(0.0885)	(0.1150)	(0.0845)	(0.0990)	(0.1037)	(0.0994)	(0.0882)
Canada	0.0124	-0.0258	0.0696	-0.0389	-0.2291**	0.1991**	0.0454
	(0.0893)	(0.1135)	(0.0948)	(0.0921)	(0.0910)	(0.0871)	(0.0839)
Germany	0.0573	0.1076	-0.0186	-0.1001	-0.0750	0.3639***	0.0174
	(0.0824)	(0.1045)	(0.0904)	(0.0845)	(0.0856)	(0.0878)	(0.0895)
Japan	0.1251	0.4202***	0.0283	-0.1960	-0.1984	0.2260*	0.1100
	(0.0859)	(0.1255)	(0.1203)	(0.3007)	(0.1289)	(0.1272)	(0.0863)

* significant at the 10% level using 2-tailed t-test
** significant at the 5% level using 2-tailed t-test
*** significant at the 1% level using 2-tailed t-test

Note: Table above presents a percentage change specification. This specification is implemented instead of natural log difference to accommodate the fact that current account data is mostly negative. Estimated coefficients are reported (with standard errors in brackets). All variables are standardized by their own standard deviation prior to estimation. This facilitates the interpretation of the coefficients as each coefficient would represent the effect of a one standard deviation change of the trade variable on the standardized real exchange rate.

FIGURE 1

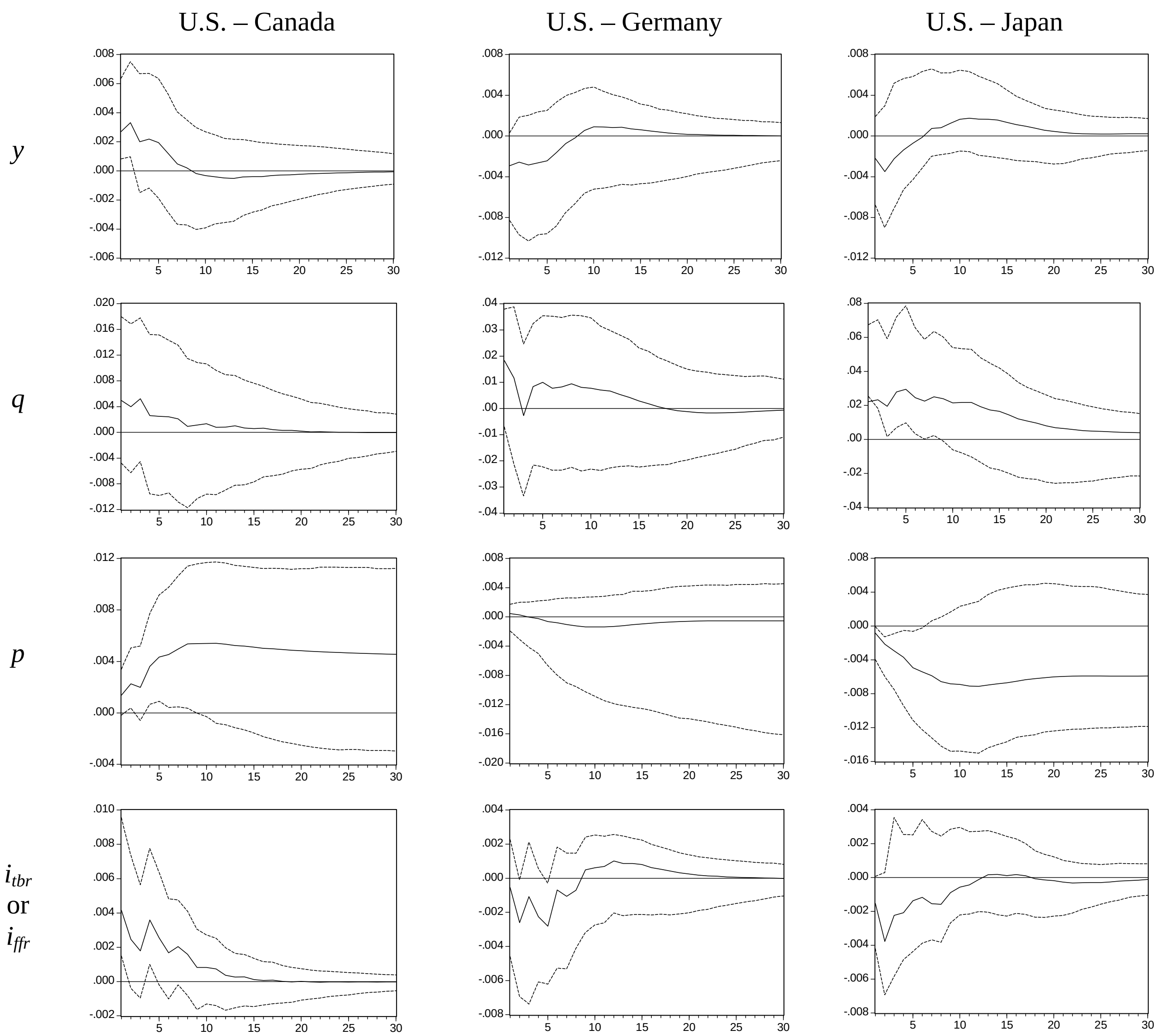
IMPULSE RESPONSES TO A ONE STANDARD DEVIATION
STRUCTURAL SHOCK FOR BENCHMARK U.S.-U.K. MODEL



Note: Benchmark specification period is from 1970Q2 to 2006Q1 with 4 lags. Impulse responses for up to 30 quarters are displayed. Confidence intervals (dashed lines) are obtained using the Kilian (1998) bias-corrected bootstrap methods. The lower dashed line indicates the 2.5th percentile of the bootstrapped distribution and the upper dashed line indicates the 97.5th percentile.

FIGURE 2

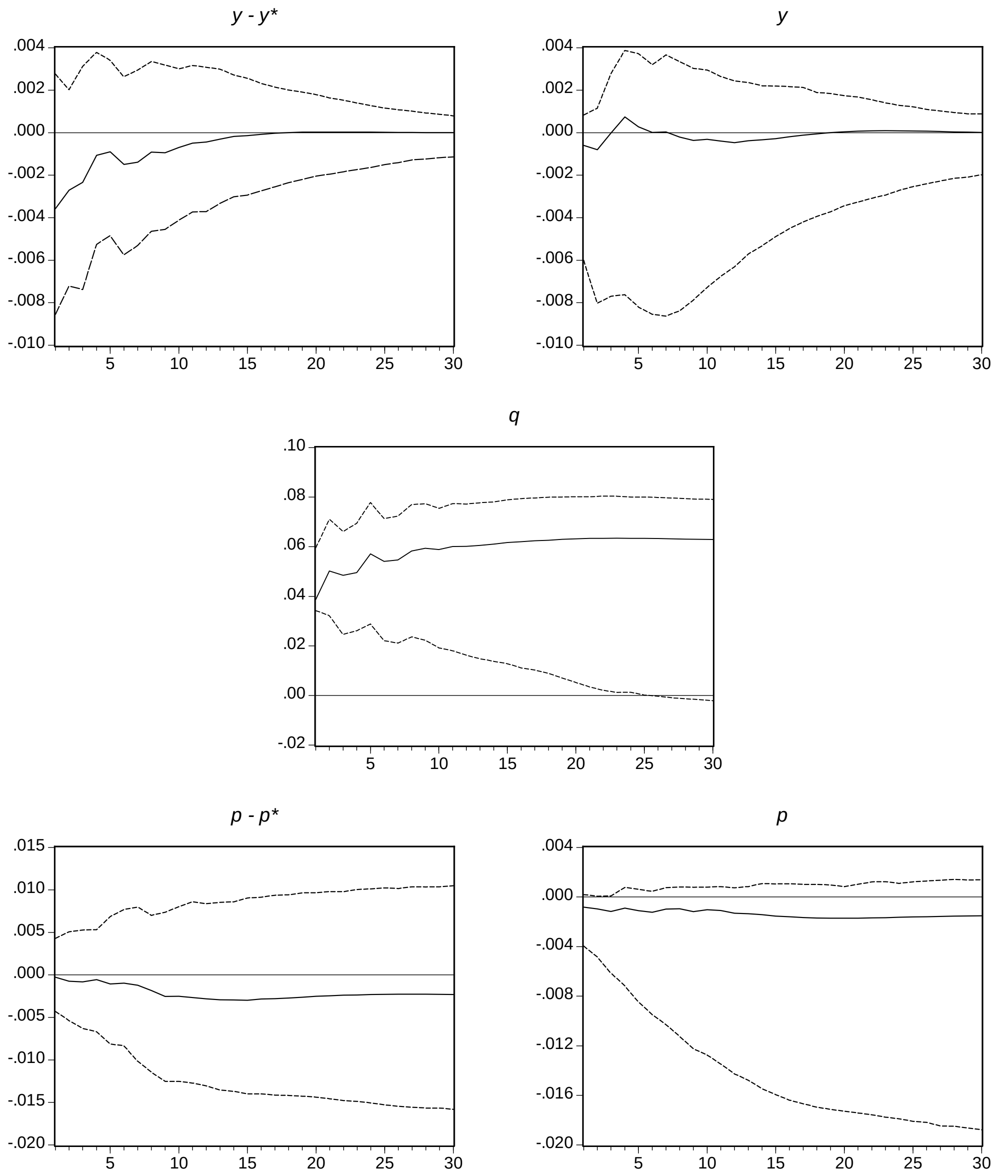
IMPULSE RESPONSES TO A ONE STANDARD DEVIATION U.S. MONETARY POLICY SHOCK FOR U.S.-CANADA, U.S.-GERMANY, AND U.S.-JAPAN MODELS



Note: Sample periods are 1970Q2 to 2006Q2 (Canada), 1970Q2 to 2006Q1 (Japan), and 1970Q2 to 2005Q4 (Germany) with 4 lags. Impulse responses for up to 30 quarters are displayed. Confidence intervals (dashed lines) are obtained using the Kilian (1998) bias-corrected bootstrap methods. The lower dashed line indicates the 2.5th percentile of the bootstrapped distribution and the upper dashed line indicates the 97.5th percentile.

FIGURE 3

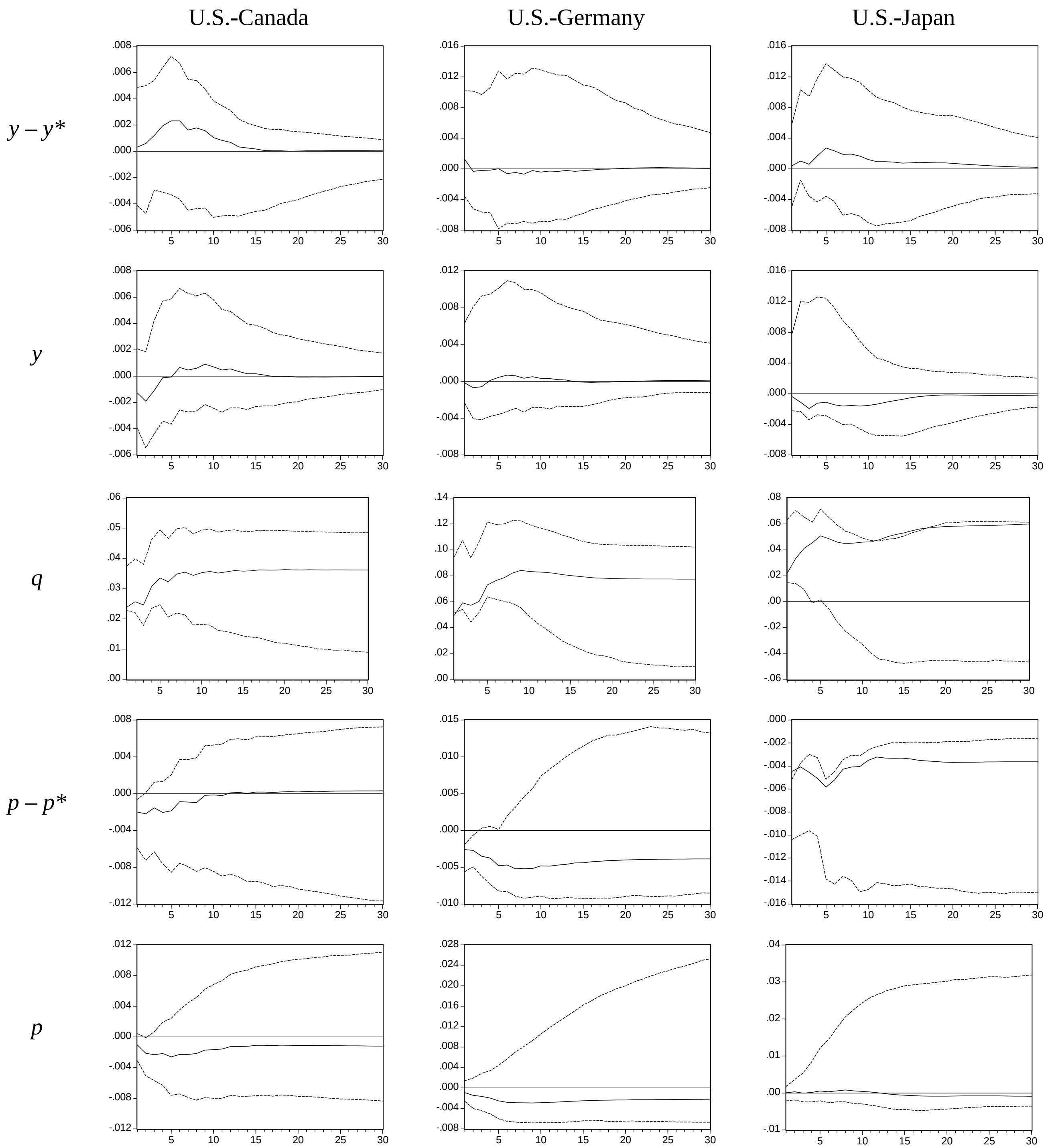
IMPULSE RESPONSES TO A ONE STANDARD DEVIATION TASTE SHOCK FOR U.S.-U.K. BENCHMARK MODEL



Note: Sample period 1970Q2 to 2006Q1 with 4 lags. Impulse responses for up to 30 quarters are displayed. Confidence intervals (dashed lines) are obtained using the Kilian (1998) bias-corrected bootstrap methods. The lower dashed line indicates the 2.5th percentile of the bootstrapped distribution and the upper dashed line indicates the 97.5th percentile.

FIGURE 4

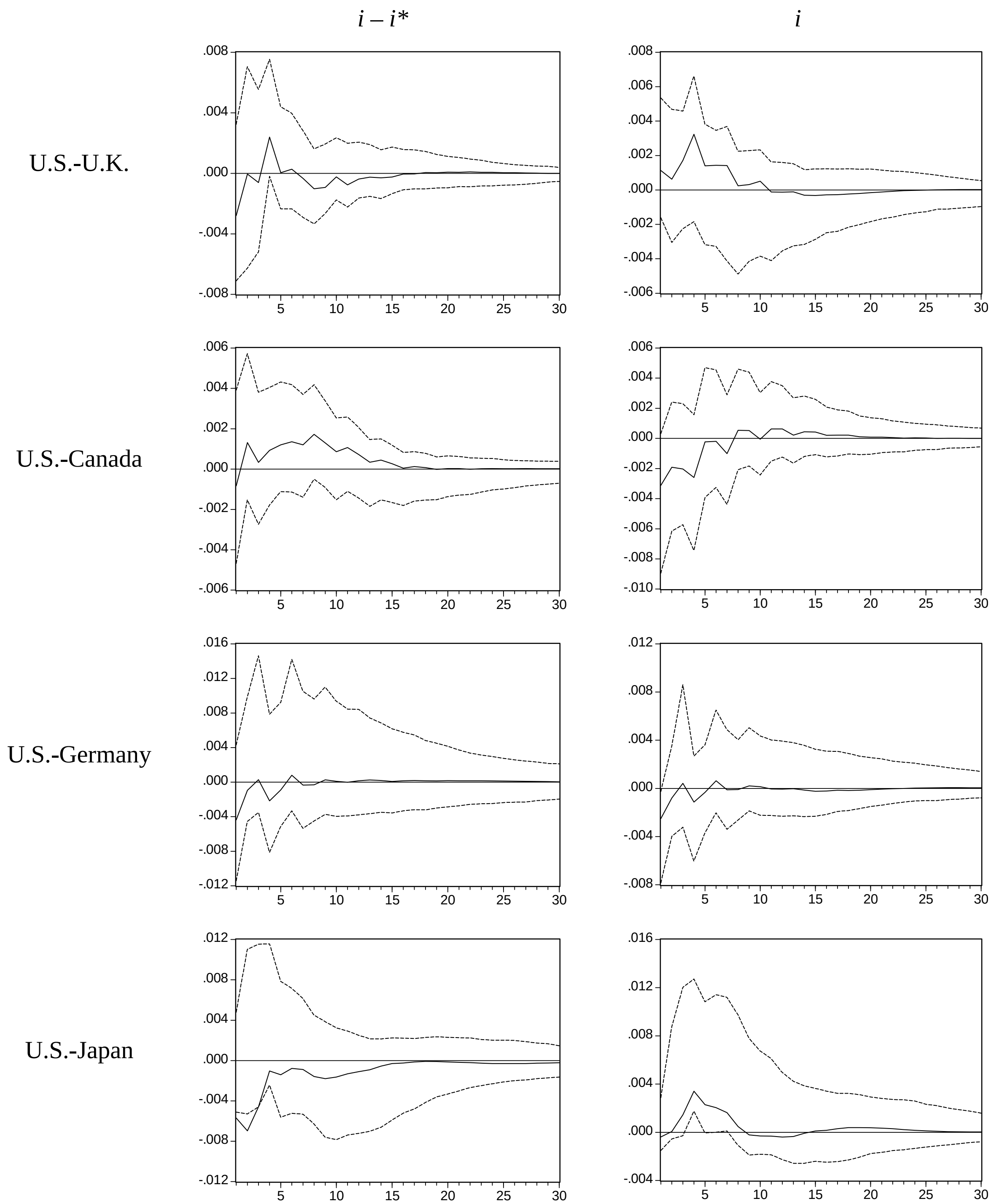
IMPULSE RESPONSES TO A ONE STANDARD DEVIATION TASTE SHOCK FOR U.S.-CANADA, U.S.-GERMANY, AND U.S.-JAPAN MODELS



Note: Sample periods are 1970Q2 to 2006Q2 (Canada), 1970Q2 to 2006Q1 (Japan), and 1970Q2 to 2005Q4 (Germany) with 4 lags. Impulse responses for up to 30 quarters are displayed. Confidence intervals (dashed lines) are obtained using the Kilian (1998) bias-corrected bootstrap methods. The lower dashed line indicates the 2.5th percentile of the bootstrapped distribution and the upper dashed line indicates the 97.5th percentile.

FIGURE 5

IMPULSE RESPONSES OF INTEREST RATE VARIABLES TO A ONE STANDARD DEVIATION TASTE SHOCK



Note: Sample periods are 1970Q2 to 2006Q2 (Canada), 1970Q2 to 2006Q1 (U.K. and Japan), and 1970Q2 to 2005Q4 (Germany) with 4 lags. Impulse responses for up to 30 quarters are displayed. Confidence intervals (dashed lines) are obtained using the Kilian (1998) bias-corrected bootstrap methods. The lower dashed line indicates the 2.5th percentile of the bootstrapped distribution and the upper dashed line indicates the 97.5th percentile.