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# Fluctuations in the Foreign Exchange Market: How Important are Monetary Policy Shocks?

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**Abstract:**
We study the effects of U.S. monetary policy shocks on the bilateral exchange rate between the U.S. and each of the G7 countries. We also estimate deviations from uncovered interest rate parity and exchange rate pass-through conditional on these shocks. The analysis is based on a structural vector autoregression in which monetary policy shocks are identified through the conditional heteroscedasticity of the structural disturbances. Unlike earlier work in this area, our empirical methodology avoids making arbitrary assumptions about the relevant policy indicator or transmission mechanism in order to achieve identification. At the same time, it allows us to assess the implications of imposing invalid identifying restrictions. Our results indicate that the nominal exchange rate exhibits delayed overshooting in response to a monetary expansion, depreciating for roughly ten months before starting to appreciate. The shock also leads to large and persistent departures from uncovered interest rate parity, and to a prolonged period of incomplete pass-through. Variance-decomposition results indicate that monetary policy shocks account for a non-trivial proportion of exchange rate fluctuations.

**Keywords:** Conditional heteroscedasticity, Delayed overshooting, Exchange rate pass-through, Identification, Structural vector autoregression, Uncovered interest rate parity

**JEL Classification:** C32, E52, F31, F41

## 1. Introduction

The dominant view in the literature on exchange rate determination is based on the premises that prices are sticky in the short run and that uncovered interest rate parity (UIP) holds continuously. As first established by Dornbusch (1976), these assumptions together imply that the nominal exchange rate must immediately overshoot its long-run level in response to a monetary policy shock. Price stickiness also implies that, conditional on a monetary policy shock, exchange rate pass-through to domestic prices is incomplete in the short run, which translates into violations of the law of one price and, therefore, deviations from purchasing power parity. Following the seminal work of Obstfeld and Rogoff (1995), recent theoretical studies on exchange rate determination have sought to incorporate these features into fully optimizing, rational-expectation models. Despite being more sophisticated, however, these models preserve the essence of the Dornbusch model, continuing to emphasize the interaction of nominal rigidities and monetary policy shocks as the main mechanism driving exchange-rate fluctuations.

The purpose of this paper is to evaluate the empirical relevance of this view. More specifically, we estimate the effects of U.S. monetary policy shocks on the bilateral exchange rate between the U.S. and each of the remaining G7 countries. We also estimate deviations from UIP and exchange rate pass-through to U.S. domestic prices conditional on these shocks. Finally, we measure the importance of monetary policy shocks in accounting for the variability of each of these variables.

Surprisingly, little empirical work has been done to measure the effects of monetary policy shocks on exchange rates. Furthermore, the handful of empirical studies that have attempted to examine this issue using vector autoregressions (VAR) did not reach a consensus regarding the direction in which and the extent to which monetary policy shocks affect the exchange rate. Indeed, some studies find that the nominal exchange rate does not immediately overshoot its long-run level in response to a monetary policy shock. Instead, it exhibits a hump-shaped profile, reaching its maximal response several months after the shock; a pattern often referred to as delayed overshooting (e.g., Eichenbaum and Evans 1995, Grilli and Roubini 1995, 1996). Others, in contrast, find that the exchange rate overshooting is nearly immediate (e.g., Kim and Roubini 2000 and Kalyvitisa and Michaelides 2001). Similarly, there is little agreement on the importance of monetary policy shocks in accounting for exchange-rate movements: estimates of the fraction of exchange rate variability that is attributed to monetary policy shocks range from essentially zero (e.g., Scholl and Uhlig 2005) to over 50 percent (e.g., Kim and Roubini 2000).

In the same vein, although it is now well established that there are significant departures from UIP, which imply the existence of predictable excess returns on the foreign-exchange market, there is little and mixed evidence on the extent to which these departures are due to unexpected changes in monetary policy. Deviations from UIP conditional on monetary policy shocks and the importance of these shocks in accounting for the variability of excess returns are found to be large in some studies (e.g., Eichenbaum and Evans 1995 and Faust and Rogers 2003) but fairly small in others (e.g., Kim and Roubini 2000).

To some extent, the discrepancy in results across these earlier studies is attributed to the method used to identify monetary policy shocks within structural vector autoregressions (SVAR). Although most of existing studies measure monetary policy shocks with innovations to the short term interest rate, they differ in the restrictions imposed on the interactions between the variables included in the SVAR, which in turn determine the mechanism through which shocks propagate. Four types of restrictions can be found in the literature: recursive zero restrictions (e.g., Eichenbaum and Evans 1995, Grilli and Roubini 1995, 1996), non-recursive zero restrictions (e.g., Kim and Roubini 2000), sign and shape restrictions (e.g., Faust and Rogers 2003 and Scholl and Uhlig 2005), and long-run restrictions (e.g., Clarida and Gali 1994 and Rogers 1999).[1]

These different types of restrictions have in common that they are arbitrary in nature. Faust and Rogers (2003) further argue that some of the commonly used zero restrictions are highly stylized and, therefore, unlikely to provide a plausible description of the transmission channels of monetary policy shocks. Based on an empirical exercise in which they eliminate all dubious identifying assumptions, they conclude that the peak response of the nominal exchange rate to a monetary policy shock may be delayed or nearly immediate, and that monetary policy may or may not be important in accounting for exchange-rate fluctuations. While this study provides useful insights into the consequences of imposing dubious identifying assumptions, it does not resolve the uncertainty surrounding the effects of monetary policy shocks on foreign-exchange variables. Scholl and Uhlig (2005) attribute this inconclusiveness to the fact that the restrictions imposed by Faust and Rogers (2003)

[1] Studies based on long-run restrictions have focused on the real rather than the nominal exchange rate.

are too weak to narrow down the range of plausible monetary policy shocks. Put differently, these restrictions lead to under-identified SVARs, so that monetary policy shocks are not uniquely determined. This in turn implies that the underlying restrictions are not testable.

This paper is in the spirit of the work of Faust and Rogers, but differs from it in several important respects. It estimates a flexible SVAR where monetary policy shocks are identified by exploiting the conditional heteroscedasticity of the structural innovations, a procedure that has recently been proposed by Normandin and Phaneuf (2004). Unlike the identification procedures used in earlier studies, which impose conditional homoscedasticity of the innovations, this data-based approach does not rely on any arbitrary assumption about the relevant indicator or transmission mechanism of monetary policy. It is therefore a judgement-free approach which, in addition, allows one to formally test the commonly used restrictions, including recursive zero restrictions, which are predominant in the literature.

Our results indicate that following an unanticipated monetary expansion, the nominal exchange rate exhibits delayed but rapid overshooting, reaching its maximal depreciation between 8 and 11 months after the shock. The monetary policy shock also triggers significant and persistent departures from UIP as well as a prolonged period of incomplete exchange rate pass-through, which exhibits a non-monotonic pattern. Interestingly, our approach generates empirically plausible results for all the variables included in the SVAR without having to impose arbitrary restrictions on their dynamic responses. Variance-decomposition results reveal that monetary policy shocks are relatively important in explaining the variability of the nominal exchange rate, with a contribution that exceeds 30 percent at the 36-month horizon in some cases. In contrast, there is no clear evidence that the empirical failure of UIP is mainly driven by monetary disturbances, at least at short horizons. Compared with the results reported by Faust and Rogers, our findings provide more conclusive evidence on the effects of monetary policy shocks. This is mainly because our identification procedure tightly identifies these shocks.

We also find that imposing invalid identifying restrictions may lead to misleading impulse-response and variance-decomposition results. In particular, when monetary policy shocks are identified with orthogonalized innovations to the federal funds rate, as is frequently assumed, the dynamic response of the nominal exchange rate to a monetary policy shock is counterfactually small and lacks the delayed overshooting pattern. The restrictions associated with the federal funds rate also result in a severe understatement of the importance of monetary policy shocks in accounting for the variability of the nominal exchange rate.

The rest of the paper is organized as follows. Section 2 describes the empirical methodology. Section 3 performs a preliminary analysis of the data. Section 4 discusses the estimated effects of monetary policy shocks on the nominal exchange rate, deviations from UIP, and exchange rate pass-through. Section 5 performs a robustness analysis. Section 6 concludes.

## 2. Empirical Methodology

### 2.1 Specification

The SVAR system (in innovation form) is:

$$A\nu_t = \epsilon_t, \tag{1}$$

where $\nu_t$ are the statistical innovations, $\epsilon_t$ are the structural innovations, and $A$ captures the interactions between current statistical innovations. The SVAR includes variables that belong to the goods market, reserve market, and foreign-exchange market. The goods variables are U.S. total output, $y_t$, the U.S. price index, $p_t$, and the world commodity-price index, $cp_t$. The reserve variables are the U.S. non-borrowed reserves, $nbr_t$, total reserves, $tr_t$, and the federal funds rate, $ff_t$. The foreign-exchange variables are the differential between the foreign and U.S. nominal short-term interest rates, $dr_t$, and the nominal exchange rate measuring the number of U.S. dollars needed to buy one unit of foreign currency, $e_t$.[2]

Following Bernanake and Mihov (1998), the market for U.S. bank reserves is further developed via the simple formulation:

$$\begin{aligned} \nu_{nbr,t} &= \phi_d\sigma_d\epsilon_{d,t} - \phi_b\sigma_b\epsilon_{b,t} + \sigma_s\epsilon_{s,t}, && (2)\\ \nu_{tr,t} &= -\alpha\nu_{ff,t} + \sigma_d\epsilon_{d,t}, && (3)\\ \nu_{tr,t} - \nu_{nbr,t} &= \beta\nu_{ff,t} - \phi_b\sigma_b\epsilon_{b,t}, && (4) \end{aligned}$$

where $\epsilon_{s,t}$ is a shock representing an exogenous policy action taken by the Fed, or monetary policy shock, while $\epsilon_{d,t}$ and $\epsilon_{b,t}$ denote respectively the shocks of demand for total reserves and for borrowed reserves by commercial banks. The parameters $\sigma_s$, $\sigma_d$, and $\sigma_b$ are the standard deviations scaling the structural innovations of interest, while $\phi_d$ and $\phi_b$ are unrestricted parameters, and $\alpha$ and $\beta$ are positive parameters. Equation (2) describes

[2] The choice of these variables is further discussed in Section 2.3.

the procedures which may be used by the Fed to select its monetary policy instruments. Equation (3) represents the banks' demand for total reserves in innovation form. Equation (4) is the banks' demand for borrowed reserves in innovation form, under the assumption of a zero discount-rate innovation.

Inserting the equilibrium solution of the model (2)–(4) in system (1) gives:

$$\begin{pmatrix} a_{11} & a_{12} & a_{13} & a_{14} & a_{15} & a_{16} & a_{17} & a_{18} \\ a_{21} & a_{22} & a_{23} & a_{24} & a_{25} & a_{26} & a_{27} & a_{28} \\ a_{31} & a_{32} & a_{33} & a_{34} & a_{35} & a_{36} & a_{37} & a_{38} \\ a_{41} & a_{42} & a_{43} & \frac{1+\phi_b}{\sigma_s} & -\frac{\phi_b+\phi_d}{\sigma_s} & \frac{\beta\phi_b-\alpha\phi_d}{\sigma_s} & a_{47} & a_{48} \\ a_{51} & a_{52} & a_{53} & 0 & \frac{1}{\sigma_d} & \frac{\alpha}{\sigma_d} & a_{57} & a_{58} \\ a_{61} & a_{62} & a_{63} & \frac{1}{\sigma_b} & -\frac{1}{\sigma_b} & \frac{\beta}{\sigma_b} & a_{67} & a_{68} \\ a_{71} & a_{72} & a_{73} & a_{74} & a_{75} & a_{76} & a_{77} & a_{78} \\ a_{81} & a_{82} & a_{83} & a_{84} & a_{85} & a_{86} & a_{87} & a_{88} \end{pmatrix} \begin{pmatrix} \nu_{y,t} \\ \nu_{p,t} \\ \nu_{cp,t} \\ \nu_{nbr,t} \\ \nu_{tr,t} \\ \nu_{ff,t} \\ \nu_{dr,t} \\ \nu_{e,t} \end{pmatrix} = \begin{pmatrix} \epsilon_{1,t} \\ \epsilon_{2,t} \\ \epsilon_{3,t} \\ \epsilon_{s,t} \\ \epsilon_{d,t} \\ \epsilon_{b,t} \\ \epsilon_{7,t} \\ \epsilon_{8,t} \end{pmatrix}, \tag{5}$$

where $a_{ij}$ $(i, j = 1, ...8)$ are unconstrained parameters. The system (5) allows interactions between the terms within and across the blocks of goods, reserve and foreign-exchange variables. As a result, all variables may be contemporaneously affected by the structural innovations and, in particular, by monetary policy shocks.

It is instructive to rewrite the fourth equation in (5) as

$$\nu_{s,t} = [\rho_{41}\nu_{y,t} + \rho_{42}\nu_{p,t} + \rho_{43}\nu_{cp,t} + \rho_{47}\nu_{dr,t} + \rho_{48}\nu_{e,t}] + \sigma_s\epsilon_{s,t}, \tag{6}$$

where $\rho_{4j} = -a_{4j}\sigma_s$ (for $j = 1, 2, 3, 7, 8$) and $\nu_{s,t} = (1 + \phi_b)\nu_{nbr,t} - (\phi_d + \phi_b)\nu_{tr,t} + (\beta\phi_b - \alpha\phi_d)\nu_{ff,t}$. Equation (6) is interpreted as the Fed's feedback rule. The term $\nu_{s,t}$ measures the statistical innovation of the monetary policy indicator. This indicator is expressed as a combination of the reserve variables, reflecting the notion that the Fed might be adopting a mixed procedure whereby it targets neither the interest rate nor a monetary aggregate exclusively. The expression in brackets in equation (6) captures the systematic response of the Fed to changes in the non-reserve variables. More precisely, the feedback rule implies that the Fed designs its policy by taking into account current values of output, the price level, commodity prices, the interest differential, and the exchange rate. Finally, the term $\sigma_s\epsilon_{s,t}$ is the scaled monetary policy shock.

The conditional scedastic structure of system (5) is:

$$A\Sigma_t A' = \Gamma_t. \tag{7}$$

The matrix $\Sigma_t = E_{t-1}(\nu_t \nu_t')$ measures the conditional non-diagonal covariance matrix of the non-orthogonal statistical innovations, $\Gamma_t = E_{t-1}(\epsilon_t \epsilon_t')$ is the conditional diagonal covariance matrix of the orthogonal structural innovations, while $I = E(\epsilon_t \epsilon_t')$ normalizes the unconditional variances of the structural innovations. The dynamics of the conditional variances of the structural innovations are determined by

$$\Gamma_t = (I - \Delta_1 - \Delta_2) + \Delta_1 \bullet (\epsilon_{t-1}\epsilon_{t-1}') + \Delta_2 \bullet \Gamma_{t-1}. \tag{8}$$

The operator $\bullet$ denotes the element-by-element matrix multiplication, while $\Delta_1$ and $\Delta_2$ are diagonal matrices of parameters. Equation (8) involves intercepts that are consistent with the normalization $I = E(\epsilon_t \epsilon_t')$. Also, (8) implies that all the structural innovations are conditionally homoscedastic if $\Delta_1$ and $\Delta_2$ are null. On the other hand, some structural innovations display time-varying conditional variances characterized by univariate generalized autoregressive conditional heteroscedastic [GARCH(1,1)] processes if $\Delta_1$ and $\Delta_2$ — which contain the ARCH and GARCH coefficients, respectively — are positive semi-definite and $(I - \Delta_1 - \Delta_2)$ is positive definite. Finally, all the conditional variances follow GARCH(1,1) processes if $\Delta_1$, $\Delta_2$, and $(I - \Delta_1 - \Delta_2)$ are positive definite.

## 2.2 Identification

Under conditional heteroscedasticity, system (5) and, in particular, the monetary policy shocks and their effects on macroeconomic variables can be identified. The sufficient (rank) condition for identification states that the conditional variances of the structural innovations are linearly independent. That is, $\lambda = 0$ is the only solution to $\Gamma\lambda = 0$, such that $(\Gamma'\Gamma)$ is invertible — where $\Gamma$ stacks by column the conditional volatilities associated with each structural innovation. The necessary (order) condition requires that the conditional variances of (at least) all but one structural innovations are time-varying. In practice, the rank and order conditions lead to similar conclusions, given that the conditional variances are parameterized by GARCH(1,1) processes (see Sentana and Fiorentini 2001).[3]

[3] Rigobon (2003) develops an alternative identification method which also exploits the conditional heteroscedasticity of the structural innovations. His method and the procedure used in this paper share the same intuition: the conditional heteroscedasticity adds equations to the system, allowing the number of unknown parameters to match the number of equations. However, the estimation strategies are different: In Rigobon's paper, the conditional variances of the structural innovations are estimated for each of the pre-selected volatility regimes, rather than from a parametric specification such as a GARCH process.

Under conditional homoscedasticity, system (5) is not identified. In this environment, certain targeting restrictions and orthogonality conditions must be imposed to identify the monetary policy shocks (rather than the entire system). The usual targeting restrictions, which define the monetary policy indicator, are the following:

- *Non-borrowed reserve (NBR) indicator:* $\phi_d = \phi_b = 0$. The Fed targets the non-borrowed reserves. Thus, the non-borrowed reserves are the single policy variable and $\nu_{s,t} = \nu_{nbr,t}$.
- *Borrowed reserve (BR) indicator:* $\phi_d = 1$ and $\phi_b = \alpha/\beta$. The Fed targets borrowed reserves. The policy variables are the non-borrowed reserves and total reserves, and the policy indicator is $\nu_{s,t} = -(1 + \alpha/\beta)(\nu_{tr,t} - \nu_{nbr,t})$.
- *Adjusted non-borrowed reserve (ANBR) indicator:* $\alpha = \phi_b = 0$. In this case, shocks to total reserves are purely demand shocks which are fully accommodated by the Fed in the short run. The policy indicator is the adjusted non-borrowed reserves or the portion of non-borrowed reserves which is orthogonal to total reserves: $\nu_{s,t} = \nu_{nbr,t} - \phi_d \nu_{tr,t}$.
- *Federal funds rate (FFR) indicator:* $\phi_d = 1$ and $\phi_b = -1$. The Fed targets the federal funds rate and decides to fully offset shocks to total reserves and borrowing demand. The federal funds rate is the single policy variable and $\nu_{s,t} = -(\beta + \alpha)\nu_{ff,t}$.

The usual orthogonality conditions are the following:

- *No impact effects:* $a_{ij} = 0$ (for $i = 1, \ldots, 3$ and $j = 4, \ldots, 8$). These restrictions impose the absence of contemporaneous effects of policy variables on goods variables. In this case, the policy shocks in (6) are orthogonal to the goods variables. Note that these effects can be direct or indirect. The direct effects are captured, for example, by $a_{i,4}$ (for $i = 1, \ldots, 3$) for the NBR indicator. The policy variables can indirectly affect the goods variables through their current effects on non-policy reserve variables and on foreign-exchange variables. For example, these indirect effects are captured by $a_{i,j}$ (for $i = 1, \ldots, 3$ and $j = 5, \ldots, 8$) for the NBR indicator.
- *No feedback effects:* $a_{ij} = 0$ (for $i = 4, \ldots, 6$ and $j = 7, 8$). These restrictions imply that $\rho_{4j} = 0$ (for $j = 7, 8$). This means that the Fed does not adjust endogenously

the monetary policy to current changes of interest differential and exchange rate. In this context, there is no systematic response of the monetary authority to movements in foreign-exchange variables.

Under conditional homoscedasticity, combining one set of targeting restrictions with all the sets of orthogonality conditions presented above ensures exact identification of monetary policy shocks. In this case, the shocks are generated from a standard VAR-based procedure that imposes the conditional homoscedasticity of the VAR residuals as well as non-testable identifying restrictions. Specifically, with a single policy variable, the shocks are computed from a Choleski decomposition of the VAR-residual covariance matrix. This decomposition is typically obtained by ordering the goods variables first, followed by the policy variable, by the other reserve variables, and by the foreign-exchange variables. This strategy has been used, for example, by Eichenbaum and Evans (1995), who combine the restrictions associated with either the FFR or NBR indicator along with the orthogonality conditions in order to achieve identification. Faust and Rogers (2003), on the other hand, impose the FFR targeting restrictions, but relax some of the orthogonality conditions, which leads to under-identification.

### 2.3 Estimation strategy and data

We estimate system (5) using a two-step procedure. The first step consists in estimating, by ordinary least squares (OLS), a $\tau$-order VAR that includes output, consumer and commodity prices, non-borrowing and total reserves, the federal funds rate, the interest rate differential, and the nominal exchange rate. From this regression, we extract estimates of the statistical innovations $\nu_t$ for $t = \tau+1, ..., T$. For given values of the (non-zero) elements of the matrices $A, \Delta_1$, and $\Delta_2$, it is then possible to construct an estimate of the conditional covariance matrix $\Sigma_t$ recursively, using equations (7) and (8) and the initialization $\Gamma_\tau = I$. The second step consists in estimating the (non-zero) elements of the matrices $A, \Delta_1$, and $\Delta_2$ using full-information maximum likelihood (FIML). This method assumes that the statistical innovations are conditionally normally distributed.

Our empirical analysis is based on monthly data taken from the International Financial Statistics, the Board of Governors of the Federal Reserve System, the Federal Reserve Bank of Saint-Louis, and the Bureau of Labor Statistics databases. Following Eichenbaum and Evans (1995), we study the effects of U.S. monetary policy shocks on several exchange rates vis-à-vis the U.S. dollar. Therefore, six different SVARs are estimated, corresponding to the currencies of the remaining G7 countries: Canada (Canadian Dollar), France (French Franc), Germany (Deutschmark), Italy (Lira), Japan (Yen), and the U.K. (British Pound). Each of these SVARs also includes the interest rate differential between the U.S. and one of the afore-mentioned countries. Since France, Italy, and Germany joined the European Monetary Union in 1999, results for these countries are based on the sample period 1982:11 to 1998:12. For the remaining countries, the sample extends to 2004:10. The starting date of the sample is widely believed to have marked the beginning of an era of stable monetary policy in the U.S.

The variables used in the analysis are constructed as follows: $y_t$ is measured by the U.S. industrial-production index; $p_t$ is measured by the U.S. consumer-price index, $cp_t$ is measured by the world-export commodity-price index; $nbr_t$ and $tr_t$ correspond to the U.S. non-borrowed and total reserves, respectively, and $ff_t$ is measured by the federal funds rate. The interest rate differential, $dr_t$, is constructed as the difference between the U.S. three-month Treasury Bill rate and its foreign counterpart.[4] The exchange rate, $e_t$, is defined as the number of U.S. dollars required to purchase one unit of foreign currency. All series, except the federal funds rate and the interest rate differential are seasonally adjusted and expressed in logarithm.[5]

Our empirical analysis does not take into account the foreign counterparts of $y_t$ and $p_t$. While, admittedly, this omission implies an asymmetric treatment of the U.S. and the foreign country, it is largely dictated by the computational burden involved in estimating GARCH processes with a large number of variables. When we added foreign output and the foreign price index to the list of variables used in estimation, the extended SVAR proved to be difficult to estimate by FIML. To circumvent this problem, and as a robustness check of our results, we estimated the extended SVAR using a conditional (limited-information) maximum-likelihood procedure.[6] The details and results of this estimation are discussed in Section 5.

[4]The short-term interest rate is measured by the three-month Treasury Bill rate for Canada, France, and the U.K., by the Call money rate for Germany and Japan, and by the money market rate for Italy.

[5]Exchange-rate series are not seasonally adjusted but are expressed in logarithm.

[6]An alternative strategy to treat the U.S. and the foreign country in a more symmetric manner would be to express output and prices in terms of differences between the U.S. and each of the remaining G7 countries. This approach would not allow us, however, to compute the response of U.S. variables to a U.S. monetary policy shock, and therefore to compare our results to those reported in existing studies.

## 3. Preliminary Analysis

### 3.1 Specification tests

For each system, the VAR process includes six lags ($\tau = 6$).[7] The McLeod and LM test statistics are often significant for $p$-order autocorrelations of the squared VAR residuals and ARCH($p$) effects (with $p = 1, 3, 6$). These results suggest the presence of conditional heteroscedasticity in some statistical innovations, which is likely to translate into time-varying conditional variances of the structural innovations.

Table 1 reports estimates of the GARCH(1,1) parameters. These estimates indicate that the U.S. monetary policy shock has a highly persistent conditional variance, as the sum of the ARCH and GARCH coefficients exceeds 0.95 in all cases. More importantly, the estimates imply that the order condition for identification is always satisfied, since seven of the eight structural innovations are heteroscedastic. Likewise, the rank condition is always satisfied given that $(\Gamma'\Gamma)$ is systematically invertible. These results mean that monetary policy shocks can be identified through the conditional heteroscedasticity of the structural innovations.[8]

Table 2 reports estimates of the reserve-market parameters. With few exceptions, the estimates are fairly similar across the six systems.[9] This is somewhat reassuring given that the six systems share the same U.S. reserve-market variables. The slope of demand for total reserves, $\alpha$, has the correct sign in all cases, while that of the supply of borrowed reserves, $\beta$, has the predicted sign in all but one case. Both parameters, however, are often imprecisely estimated.

Recall that the specification (2)–(4) leads to the following linear restrictions in (5): $a_{54} = 0$ and $a_{64} = -a_{65}$. These restrictions are tested individually and jointly using a Wald test. The results, reported in Panel A of Table 3, indicate that the restrictions cannot be rejected at any conventional level of significance, which suggests that the specification (2)–

[7] In all cases, the Ljung-Box and heteroscedastic-robust Lagrange-Multiplier (LM) test statistics are never significant (at the 1% level) for $p$-order autocorrelations and AR($p$) processes of the VAR residuals (with $p = 1, 3, 6$).

[8] In all cases, the McLoed and LM test statistics are never significant for $p$-order autocorrelations of the squared structural innovations (relative to their conditional variances) and GARCH($p$,$q$) effects (with $p = 3$, and 6 and $q = 1$). This suggests that the estimates of the GARCH(1,1) coefficients provide an adequate description of the conditional heteroscedasticity of all structural innovations, including monetary policy shocks.

[9] The only notable exceptions are the estimate of $\sigma_b$ in the system involving Japan and those of $\beta$ in the systems involving Germany and Japan.

(4) provides an empirically plausible description of the U.S. reserve market.[10] We therefore refer to the SVAR (5) as the unrestricted system, and to the dynamic responses and variance decompositions it generates as the valid ones.

Panel B of Table 3 reports for each estimated system the p-value of the Wald statistic for the joint test of the targeting and orthogonality restrictions associated with each monetary policy indicator. The results indicate that these restrictions are soundly rejected in all cases, thus implying that it is inappropriate to account for the stance of U.S. monetary policy using a framework in which the Fed focuses on a single reserve variable, and which imposes the orthogonality of the policy shock to a set of economic variables.

### 3.2 Monetary policy shocks

From the monetary authority's feedback rule (6), it is possible to extract the sequence of U.S. monetary policy shocks, $\epsilon_{s,t}$, implied by the unrestricted system (5). Smoothed measures of these shocks (computed as five-month centered moving averages) are depicted in Figure 1. A negative (positive) value of the smoothed measure represents an unanticipated contractionary (expansionary) policy by the Fed. The shaded vertical areas represent the 1991 and 2001 contractions as identified by the National Bureau of Economic Research.

The smoothed policy shocks exhibit a similar pattern across the six estimated systems. In particular, there is an episode of very large negative policy shocks in the mid-1980s, which coincides with the period in which the Fed decreased substantially its non-borrowed reserves in order to sterilize the effects of its extensive lending to the Continental Illinois Bank on total reserves (see Benston et al. 1986). Similarly, there is a sequence of contractionary policy shocks in 1988, which coincides with the exogenous and unpredictable policy tightening in December 1988 identified by Romer and Romer (1994) based on their narrative analysis of the minutes of the Federal Open Market Committee meetings. The figure also shows that the 1991 contraction was preceded by contractionary policy shocks, while the 2001 contraction was preceded by expansionary ones. During both contractions, however, monetary policy shocks were fairly small. More generally, policy shocks appear to be much less volatile in 1990s and 2000s than in the 1980s. This observation is consistent with the view that U.S. monetary policy has become more effective in stabilizing the economy in recent years.

[10] These restrictions were also tested using a likelihood-ratio test and were not rejected.

### 3.3 Dynamic responses of selected U.S. variables

In this section, we analyze the dynamic effects of U.S. monetary policy shocks on output, the price level, and the nominal interest rate. The aim of this exercise is to demonstrate that our approach yields sensible results regarding the effects of monetary policy on U.S. economic activity.

Figure 2 depicts the unrestricted response of output (first row), the price index (second row), and the federal funds rate (sixth row) to an expansionary monetary policy shock, along with their corresponding confidence intervals.[11] The size of the shock is normalized to its unconditional standard deviation. Following a monetary expansion, U.S. output increases in all of the estimated systems. In four of the six cases, the response is hump shaped, reaching its peak between 15 and 20 months after the shock. This non-monotonic pattern has been documented by several empirical VAR-based studies using various identification schemes (e.g., Christiano, Eichenbaum, and Evans 1999). The shock also triggers an increase in the price level in all but one system (the United Kingdom). While the response is relatively muted in the first months following the shock, the price index converges to a higher long-run level. The initial inertia of the price level in response to a monetary policy shock has also been observed by Christiano et al. (1999).

The monetary expansion leads to an initial decline in the federal funds rate, which persists for five months in most of the systems, followed by a sharp increase in the subsequent months. That is, the unrestricted system generates a short-lived liquidity effect. This result tends to support the so-called vanishing-liquidity-effect hypothesis, according to which the fall in the nominal interest rate following an expansionary monetary policy shock has become smaller in the post-1982 period (see also Pagan and Robertson 1995 and Christiano 1995).

In sum, the unrestricted system (5) generates empirically plausible dynamic responses of output, the price level, and the federal funds rate to a U.S. monetary policy shock. In particular, it does not lead to a price puzzle, namely, a negative response of the price level to an expansionary monetary policy shock, as is the case with some earlier studies.[12]

[11] These (possibly asymmetric) 68% confidence intervals are computed using the Bayesian procedure suggested by Sims and Zha (1999).

[12] Grilli and Roubini (1995) find a sizable price puzzle in the G7 countries. Furthermore, Scholl and Uhlig (2005) show that applying the recursive identification scheme adopted by Eichenbaum and Evans (1995) to an updated dataset leads to a price puzzle, although there was none in the original study.

# 4. Results

This section studies the predictions of system (5) regarding the dynamic effects of U.S. monetary policy shocks on the nominal exchange rate, deviations from UIP, and the degree of exchange rate pass-through. It also evaluates the relative importance of monetary policy shocks in accounting for exchange-rate and excess-return variability. In each case, we discuss the consequences of imposing counterfactual identifying assumptions. More specifically, we compare the valid dynamic responses and variance decompositions with those obtained upon imposing various sets of recursive zero restrictions.

## 4.1 The nominal exchange rate

The dynamic responses of the nominal exchange rate are depicted in Figure 3. The first row shows the unrestricted response of the U.S. dollar (vis-à-vis the currencies of the G7 countries) to an unanticipated U.S. monetary expansion and its confidence interval. The remaining rows superpose on the unrestricted response those obtained by imposing the restrictions associated with the ANBR, BR, FFR, and NBR indicators along with the orthogonality conditions. The unrestricted system implies that following a positive monetary policy shock, the U.S. dollar depreciates significantly and persistently against all the six currencies, as theory predicts. In general, the exchange rate exhibits delayed overshooting, attaining its maximum response between 8 and 11 months after the shock. This result violates Dornbusch's overshooting model according to which the nominal exchange rate should immediately exceed its long-run level in response to a monetary policy shock. On the other hand, our findings suggest that the peak response of the nominal exchange rate occurs much earlier than reported by Eichenbaum and Evans (1995).[13] The only exception occurs in the system involving Canada, where the response of the nominal exchange rate is muted at all horizons.

What are the implications of imposing invalid identifying restrictions on the system (5) for the dynamic response of the nominal exchange rate? The second, third, and fourth rows of Figure 3 show that imposing the restrictions associated with the NBR, BR, and ANBR indicators along with the orthogonality conditions yields qualitatively similar predictions to

[13] Eichenbaum and Evans (1995) find that in response to a contractionary shock to U.S. monetary policy, the U.S. dollar reaches its maximum appreciation between 24 and 39 months after the shock, before starting to depreciate.

those implied by the unrestricted specification. The three sets of restrictions imply that the nominal exchange rate reaches its peak with a delay in five of the six estimated systems. In each of these cases, however, the magnitude of the exchange rate response is relatively small compared with its unrestricted counterpart. On the other hand, imposing the restrictions associated with the FFR indicator and the orthogonality conditions generates results that deviate significantly from the valid ones. Under these restrictions, the response of the U.S. dollar is essentially flat and, with one exception (Japan), does not exhibit any clear delayed overshooting. In addition, the estimated response of the U.S.–Canada bilateral exchange rate is negative on impact, an anomaly often called the exchange rate puzzle (see, for example, Grilli and Roubini 1995, 1996).

These results suggest that the delayed overshooting of the exchange rate is robust to imposing invalid identifying restrictions, except when the monetary policy shock is identified with orthogonalized innovations to the federal funds rate. These restrictions, however, generate counterfactually weak effects of monetary policy shocks on the nominal exchange rate. Interestingly, our results are quantitatively close to those reported by Scholl and Uhlig (2005) who adopt an agnostic identification procedure based on sign restrictions. Consistently with our results, they find that the maximum response of the nominal exchange rate to a monetary policy shock takes place within a year. In contrast to their methodology which, by construction, allows them to avoid the price puzzle discussed above, our procedure resolves this puzzle without imposing any *a priori* restrictions on the response of relevant variables.

Variance-decomposition results are shown in Figure 4, which reports the fraction of the variance of the k-step ahead forecast error of the nominal exchange rate that is attributed to monetary policy shocks. The figure indicates that for very short horizons (of less than three months), monetary policy shocks explain less than 10 percent of exchange rate variability. At the one-year horizon, this fraction increases significantly, reaching 40 percent in four of the six cases. For longer horizons, it decreases slightly, but remains larger than 30 percent and statistically significant. The only case in which monetary policy shocks seem to be irrelevant for exchange rate fluctuations is the U.S.–Canada bilateral exchange rate. Overall, these findings suggest that monetary policy shocks are relatively important in accounting for exchange-rate variability. This conclusion is in line with the evidence reported by Eichenbaum and Evans (1995), but conflicts with the conclusion reached by Scholl and

Uhlig (2005).

The remaining rows of Figure 4 show that imposing the restrictions associated with the NBR, BR, ANBR, and FFR indicators leads to an important understatement of the importance of monetary policy shocks in accounting for exchange rate movements. In particular, identifying monetary policy shocks with orthogonalized innovations to the federal funds rate conveys the impression that these shocks play little role in accounting for exchange-rate variability. Among the four sets of restrictions, those associated with the NBR indicator seem to be the least harmful, although they still lead to sizable departures from the valid variance-decomposition results in some cases. The BR and ANBR indicators, on the other hand, yield virtually identical results.

### 4.2 Deviations from UIP

The last row of Figure 2 shows that, in four of the six systems, an unanticipated U.S. monetary expansion drives a positive wedge between the foreign interest rate and its U.S. counterpart, at least in the first few months following the shock. Together with the delayed overshooting of the nominal exchange rate, this positive interest rate differential implies a violation of UIP. Indeed, UIP requires that a positive interest rate differential be associated with an appreciating exchange rate, meaning that the dollar should have immediately depreciated following the shock, before subsequently appreciating towards its long-run equilibrium. Deviations from UIP imply that there are predictable excess returns in the foreign-exchange market. Following Eichenbaum and Evans (1995), we construct a measure of excess return, $\Psi_t$, as

$$\Psi_t = dr_t + e_{t+1} - e_t.$$

Hence, if UIP holds, the conditional expectation $E_t\Psi_{t+k}$ must be identically zero at any horizon $k$. The effect of a monetary policy shock on this conditional expectation is straightforward to compute from the impulse responses of the interest rate differential and the nominal exchange rate. The dynamic path of expected excess return obtained from the unrestricted and restricted systems are shown in Figure 5.

The unrestricted system predicts that a positive monetary policy shock leads to sizable and persistent departures from UIP. In some cases (France, Italy, and Japan), the expected excess return exceeds 20 basis points on impact. In other cases (Germany and the U.K.), deviations from UIP take more than three years to die away. These results are consistent with the evidence reported by Eichenbaum and Evans (1995) and Faust and Rogers (2003), which points to large conditional deviations from UIP, thus questioning the validity of models of exchange rate determination that embed the UIP condition. For Canada, however, expected excess return is small and statistically indistinguishable from zero. This outcome is a direct consequence of the muted response of the nominal exchange rate and the interest rate differential between the U.S. and Canada. A plausible explanation for this result is the high degree of financial integration between the U.S. and Canada.[14]

Figure 5 shows that imposing the restrictions associated with the different policy indicators and orthogonality conditions leads to roughly similar deviations from UIP on impact. For longer horizons, however, imposing these invalid restrictions is less innocuous, as they underestimate excess return in most cases. When monetary policy shocks are identified with orthogonalized innovations to the federal funds rate, even the sign of the implied excess return is sometimes incorrect.

How important are monetary policy shocks in accounting for deviations from UIP? Variance-decomposition results, displayed in Figure 6 indicate that the contribution of monetary policy shocks to the variability of excess return is less than 10 percent in the short run. For longer horizons, however, the evidence is rather mixed: In two cases (Germany and the United Kingdom), the fraction of variance attributed to monetary policy shocks reaches roughly 50 percent at the 36-month horizon. In three other cases, this fraction is less than 10 percent. Based on these observations, it is unclear whether the well-documented empirical failure of UIP (see surveys by Froot and Thaler 1990 and Isard 1995) is driven by monetary policy shocks.

A much more conclusive (but incorrect) evidence is obtained by imposing the restrictions associated with the different policy indicators and orthogonality conditions. The results obtained upon imposing these restrictions clearly suggest that monetary policy shocks cannot be responsible for deviations from UIP. In other words, these restrictions lead to underestimating the contribution of monetary policy shocks to the variability of excess return.

[14]See Normandin (2004) for a discussion of the factors that could have contributed to the high degree of integration between Canadian and U.S. financial markets.

### 4.3 Exchange rate pass-through

A question that has received considerable attention in recent years is the extent to which exchange rate movements are passed through to domestic prices. Traditionally, exchange rate pass-through is measured as the regression coefficient of domestic inflation on changes in the nominal exchange rate. This reduced-form approach treats pass-through as an unconditional phenomenon. Bouakez and Rebei (2008) argue, however, that the degree of exchange rate pass-through depends on the nature of the structural shocks impinging on the economy, and suggest instead to compute conditional estimates of pass-through. One obvious advantage of this approach is that it does not treat exchange rate movements as being exogenous, which would be an inappropriate assumption for a large economy like the United States. Bouakez and Rebei define pass-through conditional on a given shock as the ratio of the response of the price level to that of the nominal exchange rate. In this paper, we find it convenient to compute exchange rate pass-through conditional on a monetary policy shock, $\rho_t$, as the dynamic response of the following variable:

$$\rho_t = p_t - e_t.$$

With this definition, exchange rate pass-through is incomplete when the response of $\rho_t$ is negative and complete when the response is equal to zero. One advantage of this measure of pass-through compared with the ratio-based measure is that it is less erratic in regions where the response of the nominal exchange rate is close to zero. The ratio-based measure implies that pass-through is infinite when the response of the exchange rate changes sign.

The unrestricted and restricted measures of pass-through conditional on a U.S. monetary policy shock are depicted in Figure 7.[15] The unrestricted system implies that the monetary policy shock is followed by a prolonged period of incomplete pass-through, which lasts for more than 15 months in five of the six systems. Only in the system involving Canada, does exchange rate pass-through appear to be complete at all horizons. It is worth noting that because the response of the price index is relatively muted, the shape of exchange rate pass-through is largely determined by the response of the nominal exchange rate. In cases where the exchange rate exhibits delayed overshooting, pass-through has an inverted hump-shaped pattern, with a trough occurring between 8 and 11 months after the shock. This

[15] Since pass-through is measured as a dynamic response (which is constant for a given horizon), its variance decomposition is not defined.

observation suggests that exchange rate delayed overshooting and incomplete pass-through are two different manifestations of the same effect.

Results obtained under the restrictions associated with the NBR, BR, and ANBR indicators yield similar conclusions regarding the length of the period of incomplete pass-through. Imposing these restrictions, however, lead to overestimating the degree of pass-through during that period, especially in the systems involving Italy and Japan. Identifying monetary policy shocks with orthogonalized innovations to the federal funds rate, on the other hand, leads to a totally different pattern of exchange rate pass-through, indicating that it is nearly complete at all horizons.

## 5. Robustness Analysis

The purpose of this section is to study the robustness of our results to an alternative specification of the SVAR that treats the U.S. and the foreign country more symmetrically, as in Grilli and Roubini (1995, 1996). More specifically, we extend the baseline system (5) to an 11-variable system by adding foreign output and the foreign price index to the list of variables used in estimation and by including the U.S. and the foreign short-term interest rates separately (rather than their difference).[16] Given the difficulty to estimate the extended system by FIML, we instead use a limited-information maximum-likelihood procedure that consists of three steps. The first step is similar to that described in Section 2.3. That is, we estimate by OLS a reduced-form VAR with the 11 variables discussed above and six lags to extract the associated statistical innovations and construct an estimate of their conditional covariance matrix. In the second step, we estimate by maximum likelihood the elements of $A$, $\Delta_1$, and $\Delta_2$ that are related to the variables included in the baseline SVAR (where the interest rate differential is replaced with the U.S. interest rate). Finally, conditional on these estimates, we select the remaining elements of $A$, $\Delta_1$, and $\Delta_2$ that maximize the log-likelihood of the sample.

As in the baseline case, estimates of the GARCH(1,1) parameters obtained in the extended case imply that the order and rank conditions for identification are fulfilled. Estimates of the reserve-market parameters are also similar to those reported in Table 2.

[16] Foreign output and the foreign price index are respectively measured by the industrial-production index and the consumer-price index in each of the non-U.S. G7 countries. Both series are seasonally adjusted and expressed in logarithm.

Moreover, test results indicate that the restrictions $a_{54} = 0$ and $a_{64} = -a_{65}$ cannot be rejected at conventional significance levels whereas the targeting and orthogonality restrictions discussed in Section 2.2 are soundly rejected by the data.[17] Below, we discuss the predictions of the extended SVAR regarding the effects of monetary policy shocks on the nominal exchange rate, excess return, and exchange rate pass-through. To conserve space, we only report the unrestricted results and those obtained upon imposing the restrictions associated with the FFR indicator and the orthogonality conditions, since those restrictions yield the largest departures from the unrestricted results.

The top panels of Figure 8 show the unrestricted response of the nominal exchange rate to an expansionary monetary policy shock. As before, the nominal exchange rate depreciates and exhibits delayed overshooting in five of the six estimated systems, with a peak occurring between 8 to 12 months after the shock. Canada is again an outlier in that the response of its bilateral exchange rate with the U.S. is muted and mostly statistically insignificant. When identification is achieved by imposing the restrictions associated with the FFR indicator, the resulting dynamic responses deviate substantially from their unrestricted counterparts, especially in the case of France, Germany, and Italy (see the bottom panels of Figure 8). Variance-decomposition results for the nominal exchange rate are shown in Figure 9. Because the extended SVAR involves more shocks than the baseline SVAR, the contribution of monetary policy shocks to exchange rate variability is, by construction, smaller in the former than in the latter.[18] Nonetheless, the top panels of Figure 9 indicate that monetary policy shocks still explain a non-trivial share of exchange-rate movements. For example, the contribution of these shocks to the conditional variance of the one-year step ahead forecast error of the nominal exchange rate exceeds 10 percent in most cases and reaches 30 percent in the case of France and Italy.

The unrestricted dynamic responses of excess return, depicted in the top panels of Figure 10, are similar in shape to those obtained in the baseline SVAR, but are smaller in magnitude. Excess return is statistically significant on impact only in two of the six estimated systems (France and Japan). In the systems involving Germany, Japan and the U.K., there is a sizable and statistically significant excess return (exceeding 20 basis points in the latter case) at horizons between one and two years after the shock. Hence, there is

[17]Estimation and test results are not reported but are available upon request.

[18]This can easily be verified by comparing the top panels of Figures 4 and 9.

clear evidence of deviations from UIP conditional on U.S. monetary policy shocks, although not to the extent suggested by the baseline SVAR. The bottom panels of Figure 10 show that the size and, perhaps more importantly, the sign of excess return are often incorrect when predicted by a restricted SVAR in which monetary policy shocks are identified with orthogonalized innovations to the federal funds rate. Variance-decomposition results, shown in Figure 11, indicate that monetary policy shocks explain very little (less than 5 percent) of the variability of excess return at all horizons, thus reinforcing the view that deviations from UIP are unlikely to be driven by monetary policy shocks.

Exchange rate pass-through conditional on a U.S. monetary policy shock is reported in Figure 12. As in the baseline case, conditional pass-through is a mirror image of the dynamic response of the nominal exchange rate. The unrestricted SVAR predicts that a monetary policy shock triggers a significant departure from complete pass-through (except for Canada), which attains its maximum between 8 and 12 months after the shock and which persists for several months. In contrast, the results obtained by imposing the restrictions associated with the FFR indicator convey the misleading impression that pass-through is much more complete.

Overall, these results closely resemble those implied by the baseline SVAR. In particular, our main findings regarding the delayed overshooting of the nominal exchange rate, the existence of excess returns and the delayed pass-through conditional on a monetary policy shock appear to be very robust.

## 6. Conclusion

This paper has estimated the effects of monetary policy shocks on the nominal exchange rate, excess return, and exchange rate pass-through using a flexible SVAR approach that relaxes arbitrary identifying assumptions associated with the choice of the policy indicator and orthogonality conditions. Our approach identifies monetary policy shocks through the conditional heteroscedasticity of the structural innovations, thus enabling us to test the restrictions commonly imposed in the literature in order to achieve identification.

Our results indicate that a monetary policy expansion leads to a delayed overshooting of the nominal exchange rate, with a peak occurring at around 10 months after the shock, to large deviations from UIP, and to several months of incomplete pass-through. We also show that imposing invalid identifying restrictions may deliver misleading results regarding the effects of monetary policy shocks. Finally, we find that monetary policy shocks account for a non-trivial proportion of the variability of the exchange rate and exchange rate pass-through. While the latter observation lends support to standard sticky-price models of exchange rate determination, the delayed overshooting of the nominal exchange rate and the existence of predictable excess returns on the foreign exchange market are clearly at odds with the predictions of these models. This paper's results, therefore, provide guidance about the direction in which models of exchange rate determination should be amended to better fit the data.

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**Table 1. Estimates of the GARCH(1,1) Parameters**

| | Canada | France | Germany | Italy | Japan | U.K. |
|---|---|---|---|---|---|---|
| $\epsilon_{1,t}$ | – | – | – | – | – | – |
| | – | – | – | – | – | – |
| $\epsilon_{2,t}$ | 0.331 | 0.126 | 0.091 | 0.133 | 0.043 | 0.067 |
| | (0.122) | (0.079) | (0.042) | (0.114) | (0.022) | (0.028) |
| | 0.569 | 0.844 | 0.905 | 0.776 | 0.950 | 0.932 |
| | (0.107) | (0.088) | (0.047) | (0.205) | (0.023) | (0.031) |
| $\epsilon_{3,t}$ | 0.222 | 0.472 | 0.073 | 0.276 | 0.460 | 0.288 |
| | (0.122) | (0.205) | (0.061) | (0.214) | (0.142) | (0.124) |
| | 0.434 | 0.279 | 0.890 | 0.519 | 0.427 | 0.375 |
| | (0.289) | (0.183) | (0.107) | (0.306) | (0.094) | (0.253) |
| $\epsilon_{s,t}$ | 0.193 | 0.540 | 0.600 | 0.207 | 0.211 | 0.253 |
| | (0.037) | (0.163) | (0.189) | (0.044) | (0.050) | (0.047) |
| | 0.805 | 0.424 | 0.357 | 0.790 | 0.788 | 0.744 |
| | (0.037) | (0.147) | (0.130) | (0.044) | (0.050) | (0.046) |
| $\epsilon_{d,t}$ | 0.279 | 0.350 | 0.302 | 0.361 | 0.190 | 0.360 |
| | (0.123) | (0.215) | (0.182) | (0.236) | (0.124) | (0.136) |
| | – | – | – | – | – | – |
| $\epsilon_{b,t}$ | 0.367 | 0.144 | 0.309 | 0.263 | 0.143 | 0.235 |
| | (0.174) | (0.172) | (0.254) | (0.207) | (0.068) | (0.128) |
| | – | – | – | – | – | – |
| $\epsilon_{7,t}$ | 0.589 | 0.688 | 0.129 | 0.623 | 0.108 | 0.458 |
| | (0.159) | (0.200) | (0.164) | (0.267) | (0.048) | (0.129) |
| | 0.358 | 0.229 | 0.711 | 0.265 | 0.878 | – |
| | (0.154) | (0.177) | (0.384) | (0.202) | (0.051) | – |
| $\epsilon_{8,t}$ | 0.369 | 0.181 | 0.362 | 0.093 | 0.474 | 0.203 |
| | (0.150) | (0.187) | (0.211) | (0.187) | (0.177) | (0.137) |
| | – | – | – | – | – | – |

Notes: Entries are the estimates (standard errors) of the parameters of the GARCH(1,1) processes. For each structural innovation, the first and second rows refer to the ARCH and GARCH coefficients, respectively. A dash ($-$) indicates that zero-restrictions are imposed to ensure that $\Delta_1$ and $\Delta_2$ are non-negative definite.

**Table 2. Estimates of the Reserve-Market Parameters**

| | Canada | France | Germany | Italy | Japan | U.K. |
|---|---|---|---|---|---|---|
| $\sigma_s$ | 0.042 | 0.022 | 0.019 | 0.035 | 0.027 | 0.022 |
| | (0.032) | (0.084) | (0.124) | (0.141) | (0.215) | (0.016) |
| $\sigma_d$ | 0.119 | 0.012 | 0.011 | 0.059 | 0.100 | 0.014 |
| | (0.132) | (0.004) | (0.003) | (0.104) | (0.086) | (0.002) |
| $\sigma_b$ | 0.035 | 0.0113 | $-0.243$ | 0.088 | 4.144 | 0.098 |
| | (0.010) | (0.145) | (0.654) | (0.073) | (25.539) | (0.088) |
| $\alpha$ | 0.083 | 0.045 | 0.035 | 0.206 | 0.045 | 0.001 |
| | (0.137) | (0.022) | (0.025) | (0.105) | (0.098) | (0.008) |
| $\beta$ | 0.181 | 0.366 | $-0.897$ | 0.357 | 3.458 | 0.480 |
| | (0.051) | (0.517) | (2.415) | (0.304) | (19.155) | (0.426) |
| $\phi_d$ | 0.883 | 0.978 | 0.868 | 0.762 | 0.902 | 0.896 |
| | (0.269) | (0.446) | (0.760) | (1.058) | (7.644) | (0.250) |
| $\phi_b$ | 0.375 | 0.152 | $-0.046$ | 0.452 | 0.011 | 0.002 |
| | (0.595) | (0.810) | (0.780) | (0.983) | (7.923) | (0.038) |

Notes: Entries are the estimates of the structural parameters of the U.S. reserve market. Numbers in parentheses are standard errors.

**Table 3. Test Results**

| | Canada | France | Germany | Italy | Japan | U.K. |
|---|---|---|---|---|---|---|
| | *Panel A: Tests of theoretical restrictions* | | | | | |
| $a_{54} = 0$ | [0.868] | [0.982] | [0.868] | [0.535] | [0.903] | [0.749] |
| $a_{64} = -a_{65}$ | [0.967] | [0.317] | [0.880] | [0.386] | [0.912] | [0.649] |
| *Joint* | [0.985] | [0.590] | [0.978] | [0.529] | [0.986] | [0.858] |
| | *Panel B: Tests of identification conditions* | | | | | |
| Targeting and Orthogonality | [0.000] | [0.000] | [0.000] | [0.000] | [0.000] | [0.000] |
| Targeting, Orthogonality, and Homoscedasticity | [0.000] | [0.000] | [0.000] | [0.000] | [0.000] | [0.000] |

Note: Entries are the p-values of the $\chi^2$-distributed Wald test statistics. Targeting restrictions refer to the sets of restrictions associated with the NBR, BR, ANBR, or FFR indicator. Identical p-values are obtained when each of these sets is combined with (*i*) orthogonality conditions and (*ii*) orthogonality conditions and homoscedasticity.

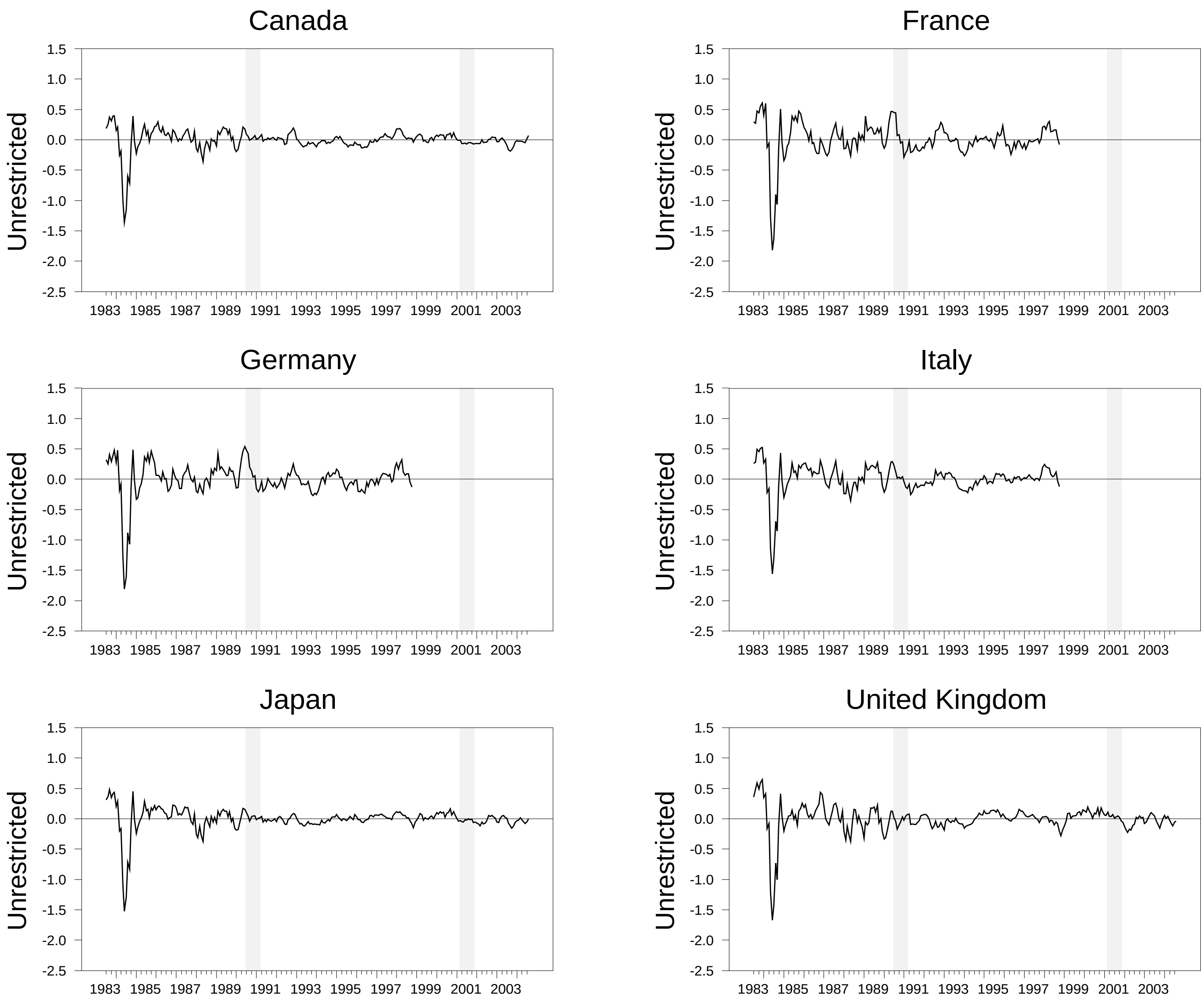


Figure 1: Monetary Policy Shocks

Solid lines: unrestricted monetary policy shocks, shaded areas: NBER contractions

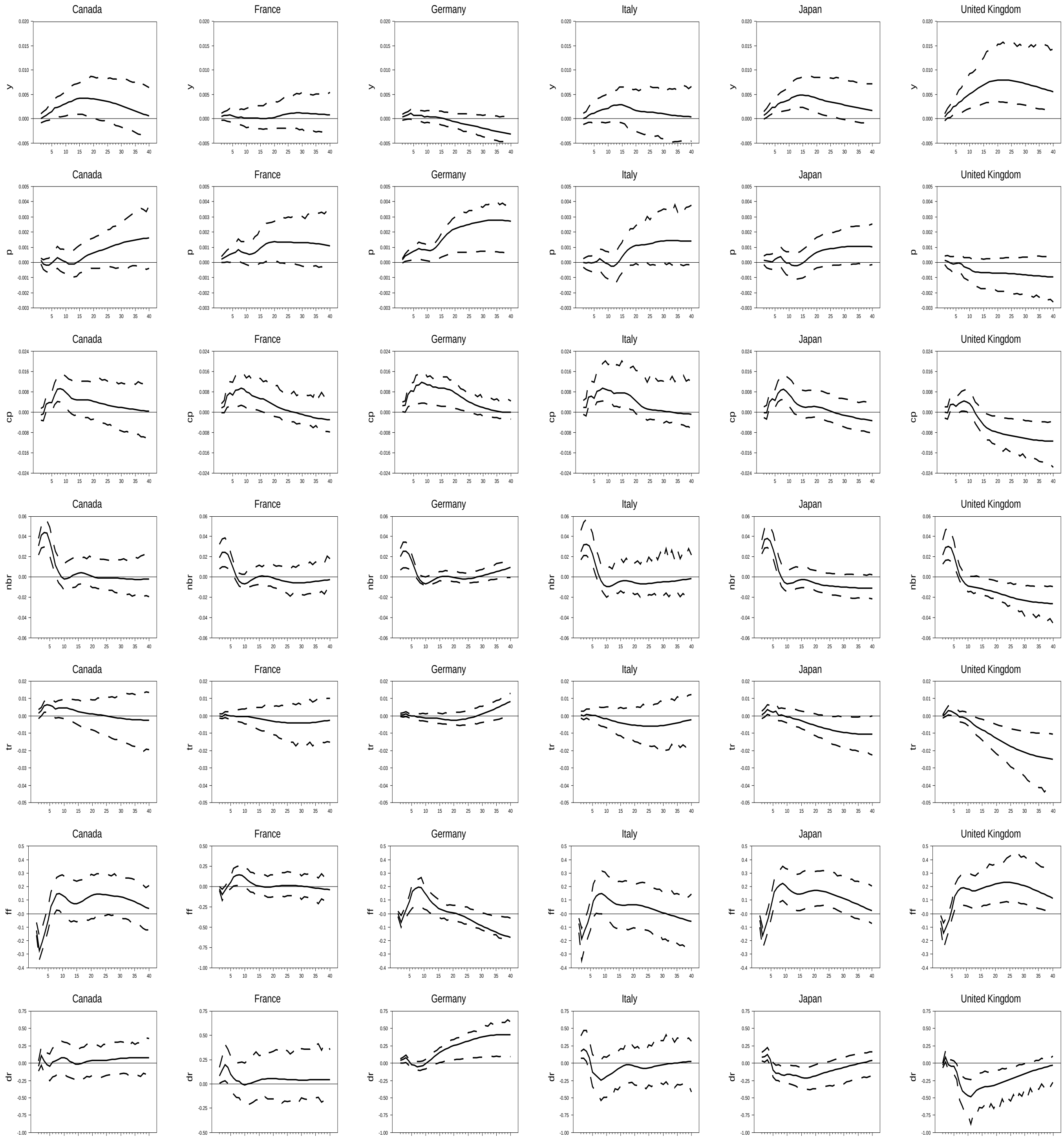


Figure 2: Unrestricted Dynamic Responses to an Expansionary Monetary Policy Shock

Solid lines: unrestricted, dashed lines: confidence intervals

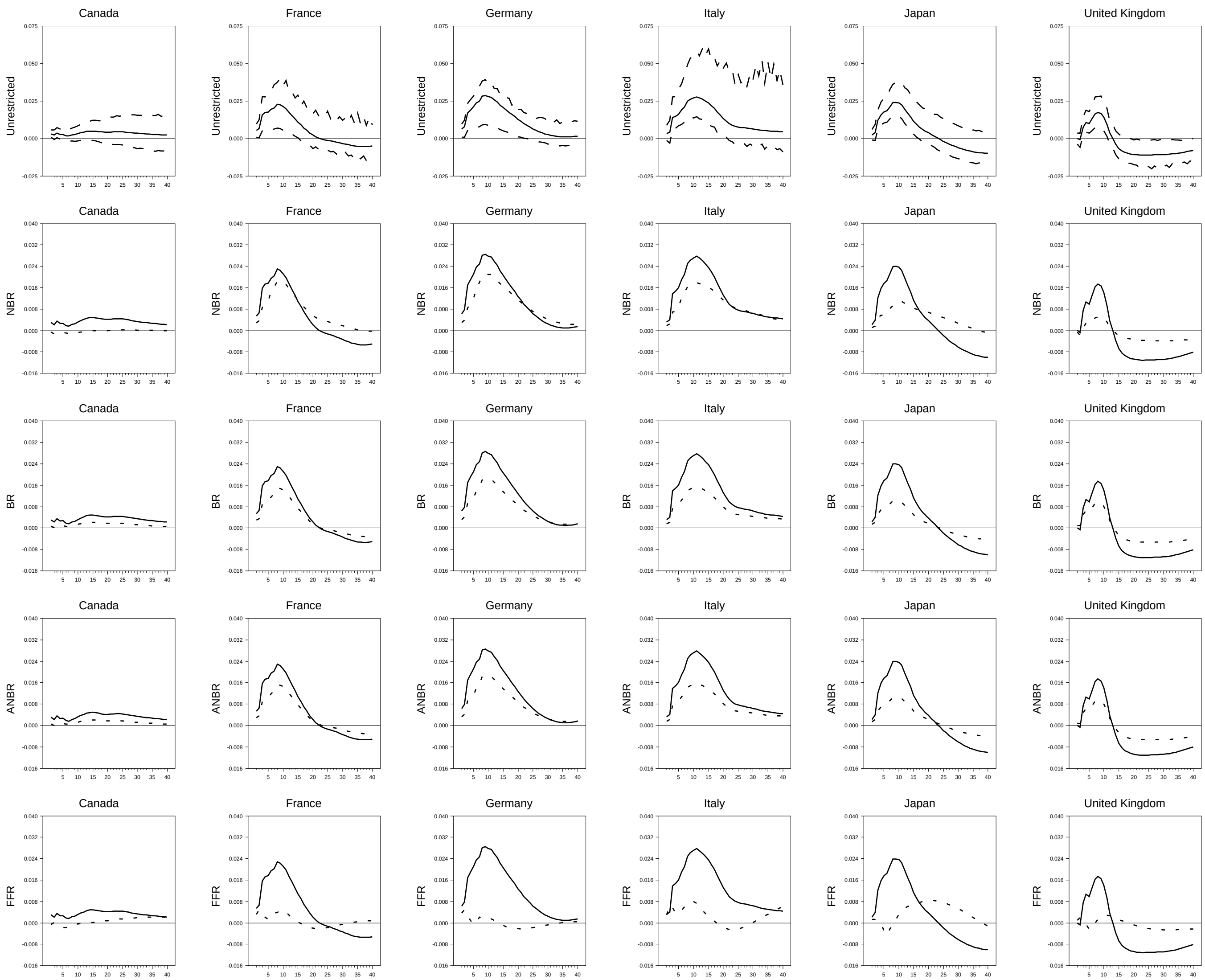


Figure 3: Dynamic Responses of the Nominal Exchange Rate to an Expansionary Monetary Policy Shock

Solid lines: unrestricted, dashed lines: confidence intervals, dotted lines: restricted (targeting+orthogonality)

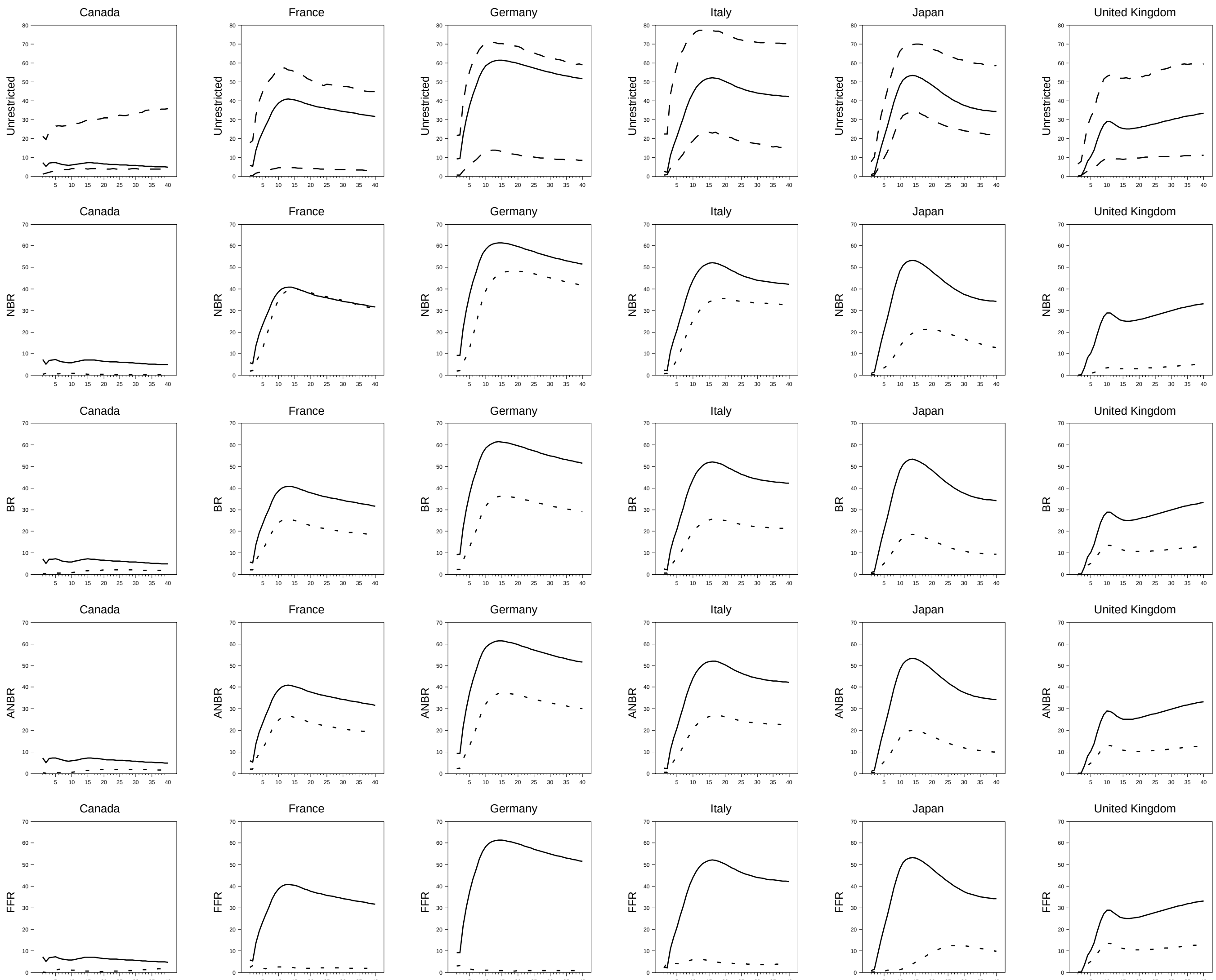


Figure 4: Contribution of Monetary Policy Shocks to the Conditional Variance of the k-Step Ahead Forecast Error of the Nominal Exchange Rate

Solid lines: unrestricted, dashed lines: confidence intervals, dotted lines: restricted (targeting+orthogonality)

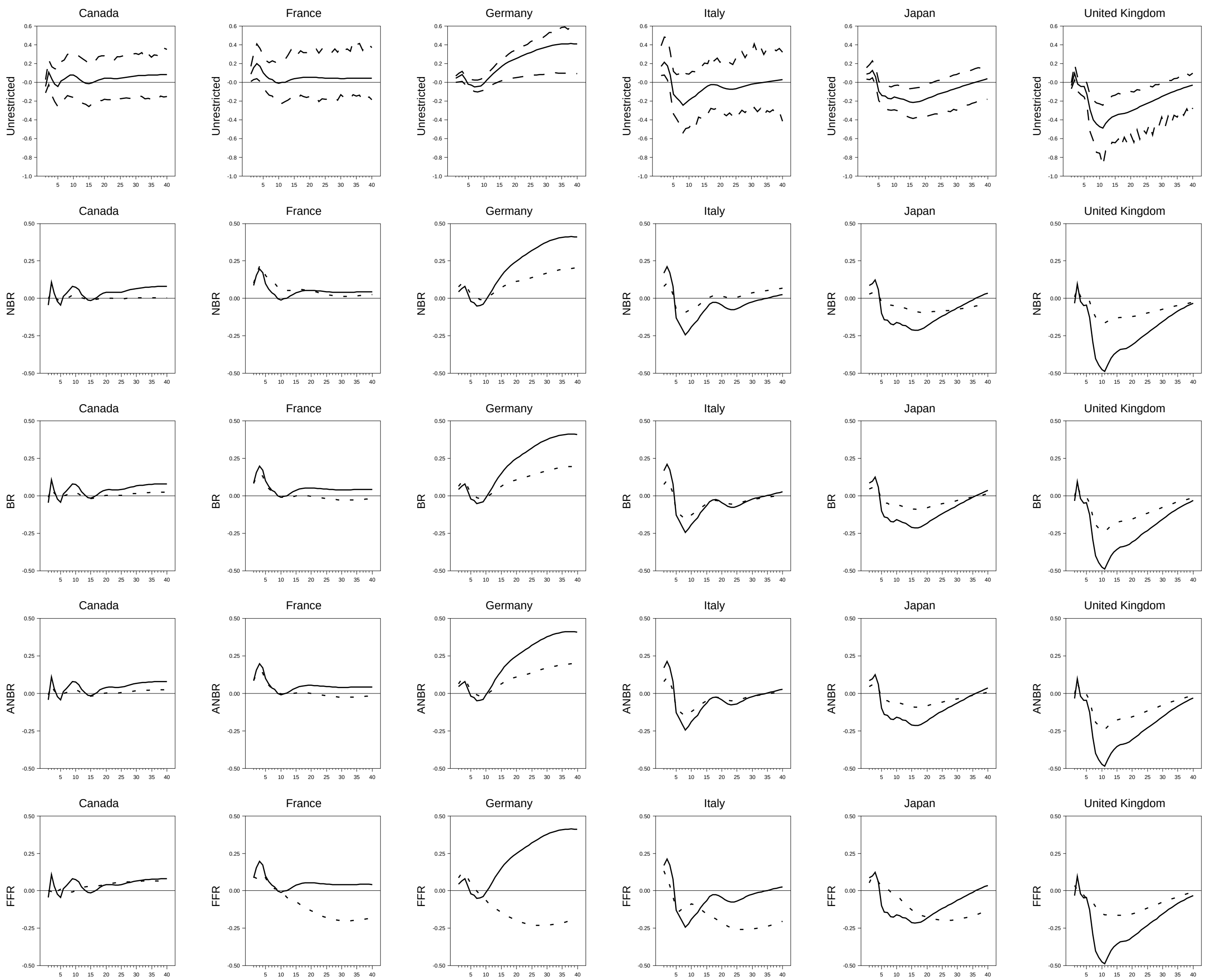


Figure 5: Dynamic Responses of Excess Return to an Expansionary Monetary Policy Shock

Solid lines: unrestricted, dashed lines: confidence intervals, dotted lines: restricted (targeting+orthogonality)

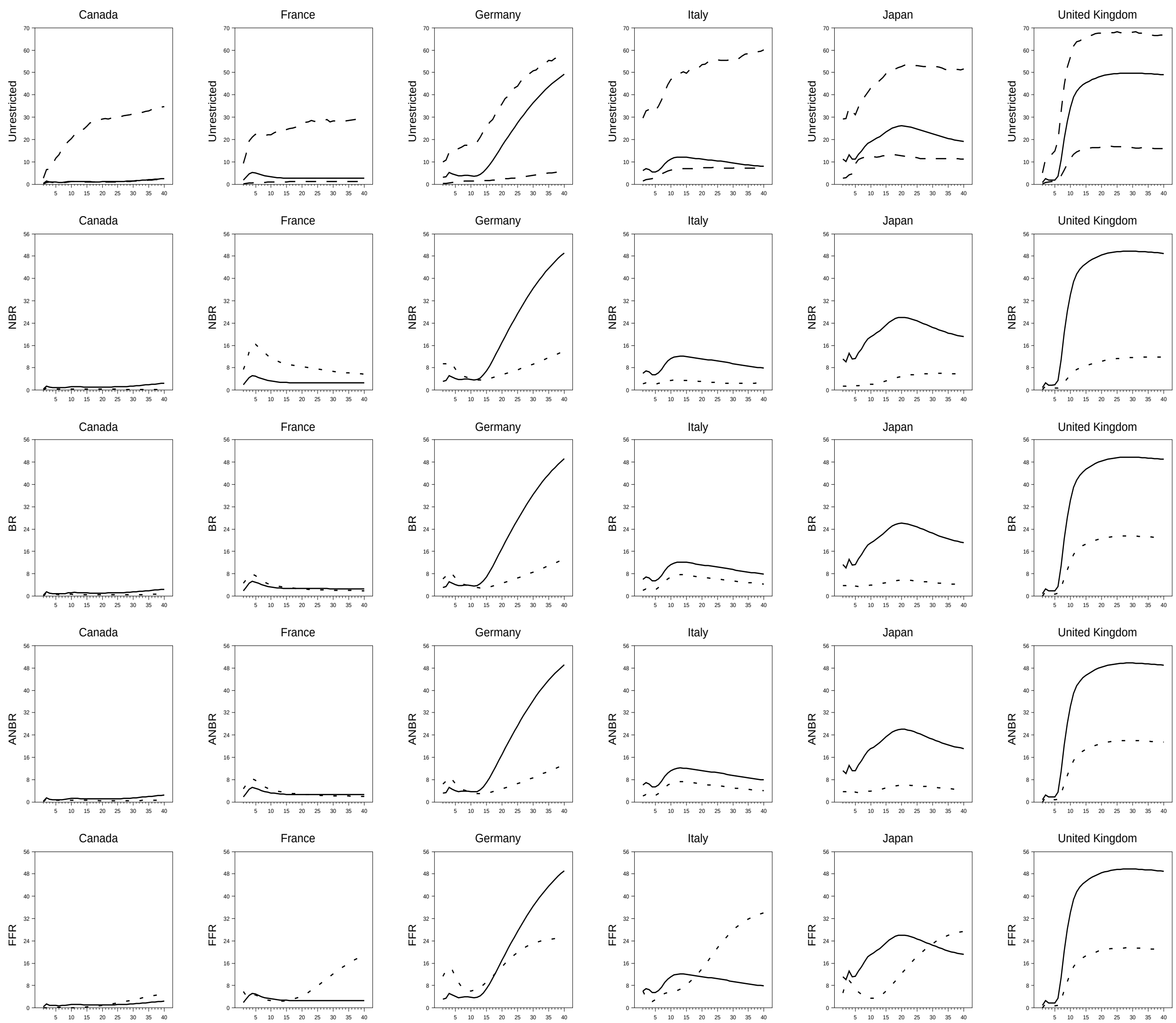


Figure 6: Contribution of Monetary Policy Shocks to the Conditional Variance of the k-Step Ahead Forecast Error of Excess Return

Solid lines: unrestricted, dashed lines: confidence intervals, dotted lines: restricted (targeting+orthogonality)

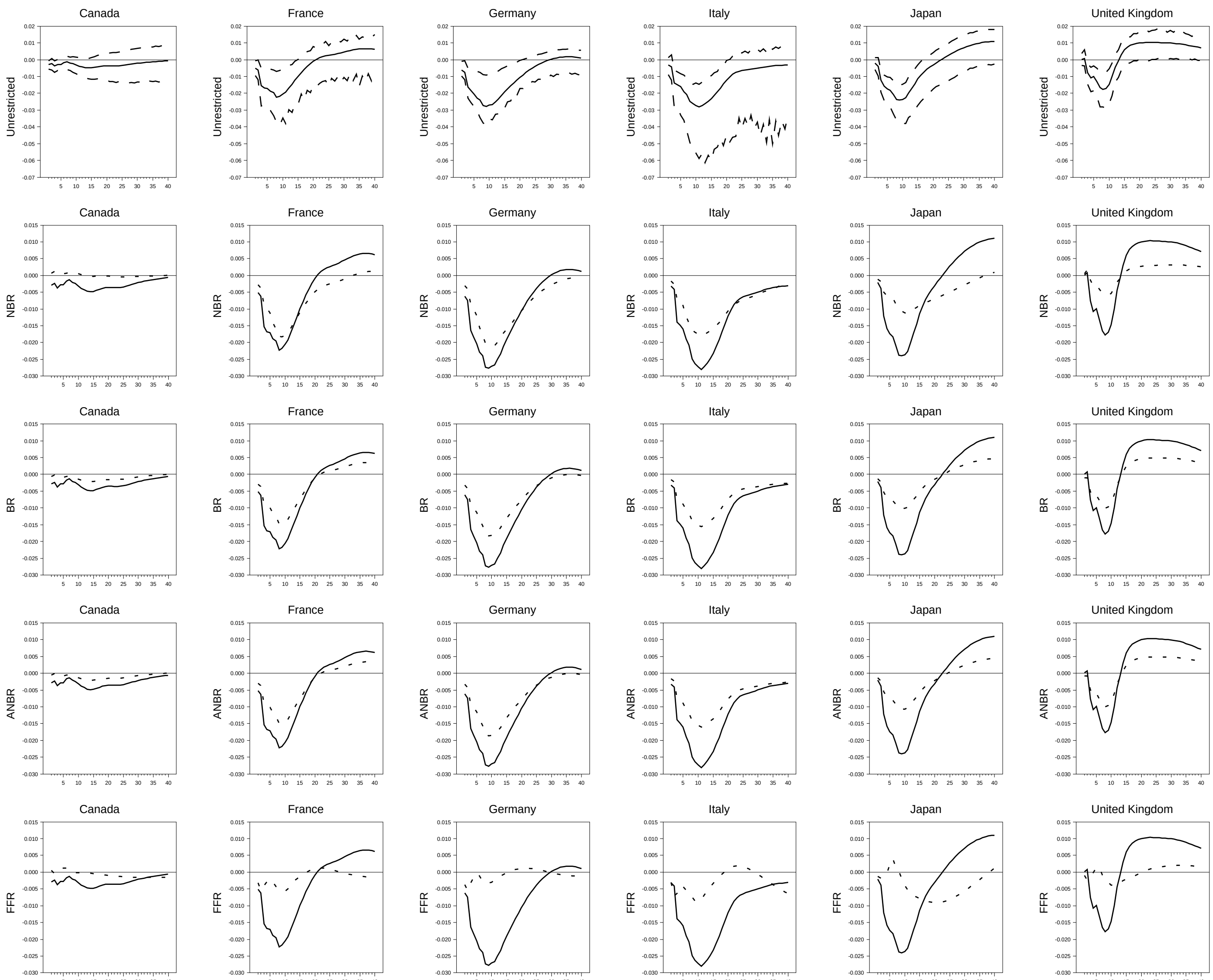


Figure 7: Exchange Rate Pass-Through Conditional on an Expansionary Monetary Policy Shock

Solid lines: unrestricted, dashed lines: confidence intervals, dotted lines: restricted (targeting+orthogonality)

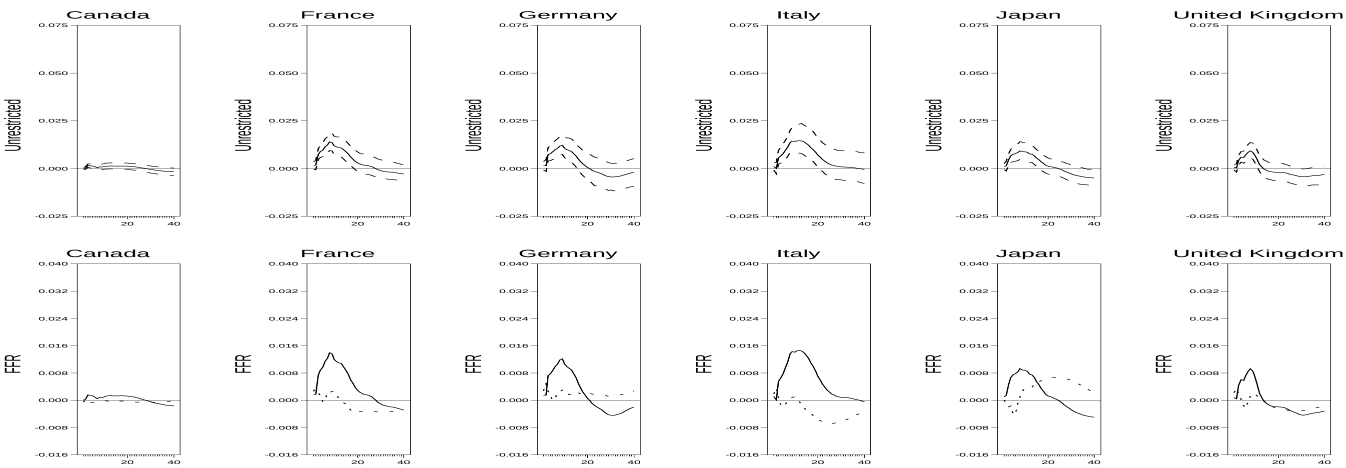


Figure 8: Dynamic Responses of the Nominal Exchange Rate to an Expansionary Monetary Policy Shock in the Extended SVAR

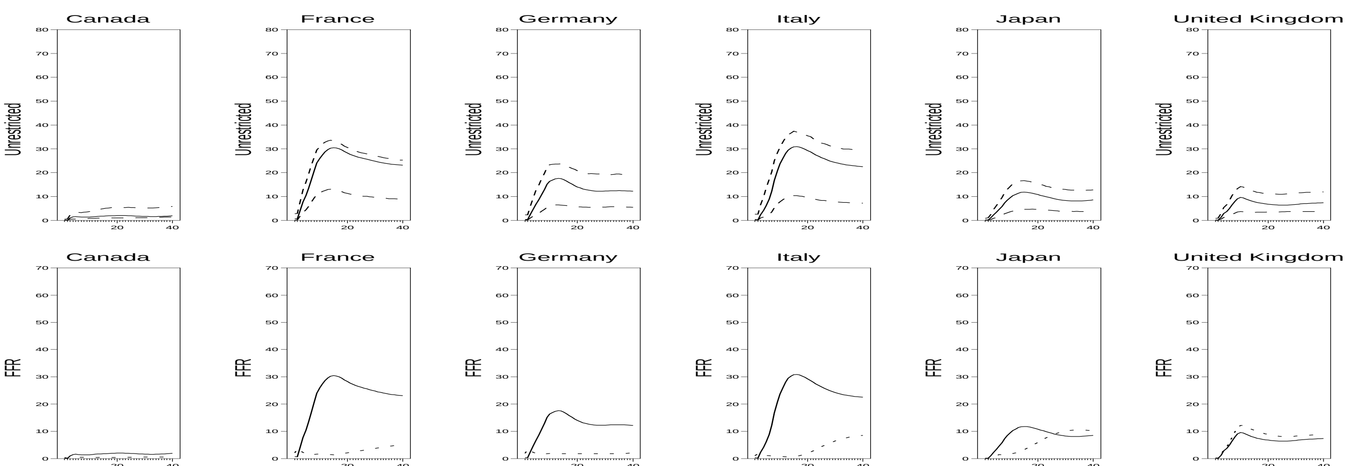


Figure 9: Contribution of Monetary Policy Shocks to the Conditional Variance of the k-Step Ahead Forecast Error of the Nominal Exchange Rate in the Extended SVAR

Solid lines: unrestricted, dashed lines: confidence intervals, dotted lines: restricted (targeting+orthogonality)

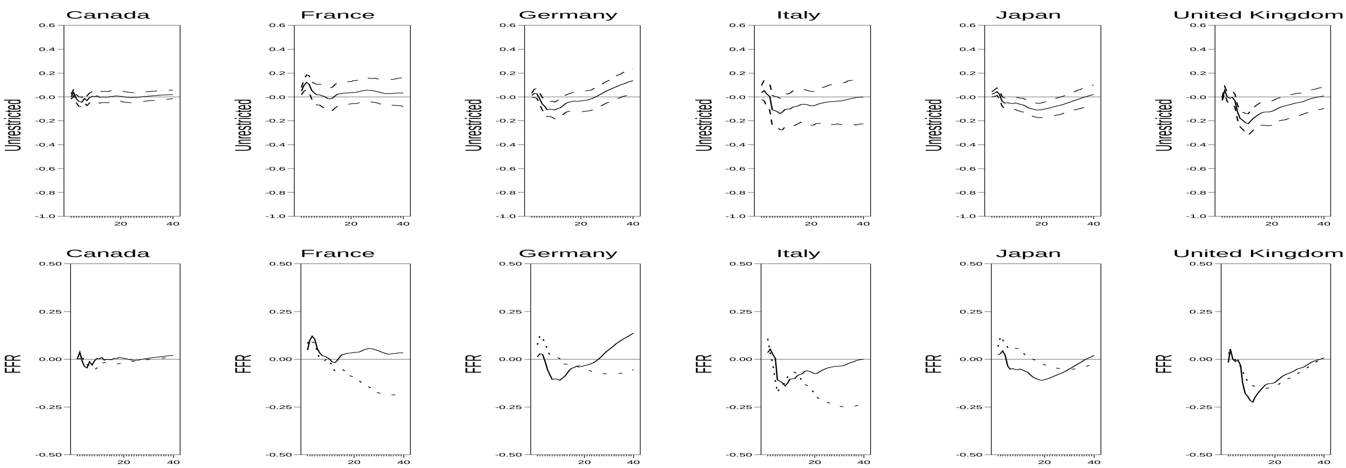


Figure 10: Dynamic Responses of Excess Return to an Expansionary Monetary Policy Shock in the Extended SVAR

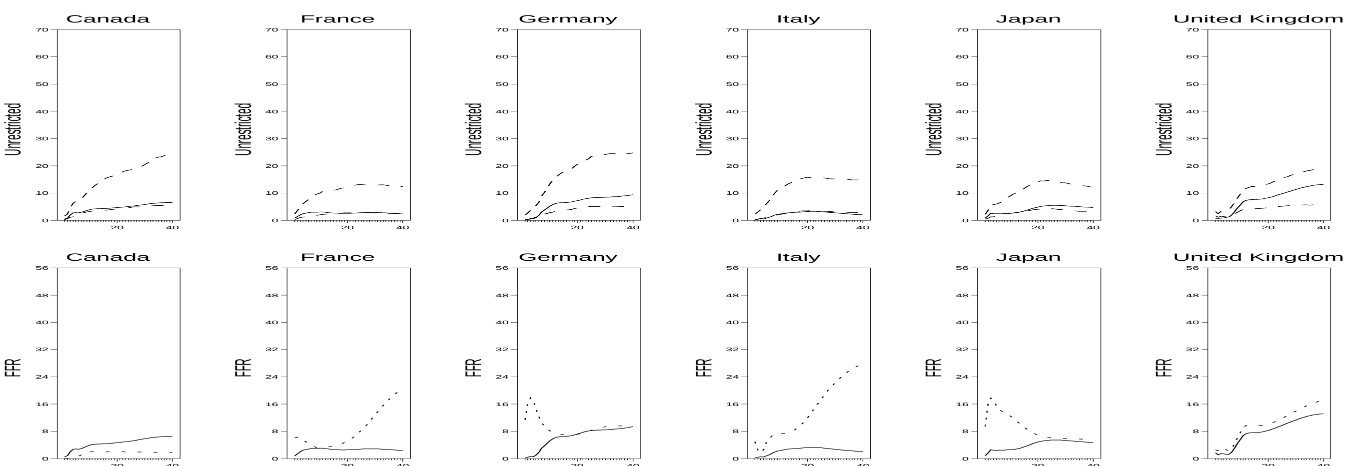


Figure 11: Contribution of Monetary Policy Shocks to the Conditional Variance of the k-Step Ahead Forecast Error of Excess Return in the Extended SVAR

Solid lines: unrestricted, dashed lines: confidence intervals, dotted lines: restricted (targeting+orthogonality)

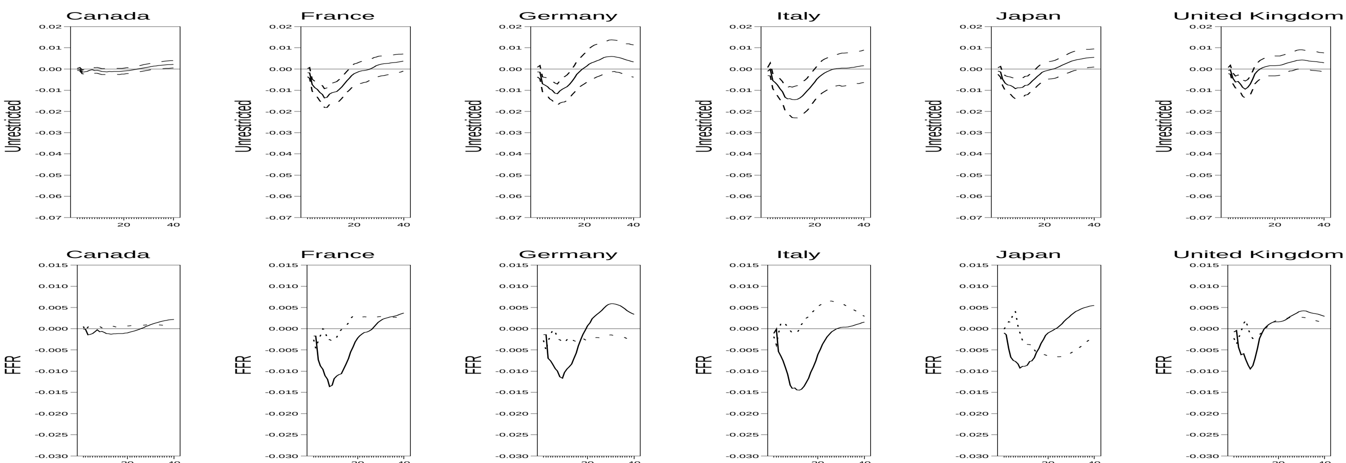


Figure 12: Exchange Rate Pass-Through Conditional on an Expansionary Monetary Policy Shock in the Extended SVAR

Solid lines: unrestricted, dashed lines: confidence intervals, dotted lines: restricted (targeting+orthogonality)