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MONETARY SHOCKS IN THE G-6 COUNTRIES:

IS THERE A PUZZLE?

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ABSTRACT

This paper attempts to reduce the uncertainty about the dynamics of the monetary transmission mechanism. Central to this attempt is the identification of monetary policy shocks. Recently, VAR approaches that use over-identifying restrictions have shown success in isolating such shocks. This paper examines monetary shocks identified by long-run cointegration restrictions and the assumption of long-run money neutrality in exactly identified VAR models across six industrialized countries. The short-run dynamics corresponding to a monetary shock can be interpreted as a monetary policy shock. The results suggest that the stock of money has an active role in the transmission mechanism.

RÉSUMÉ

Les auteurs cherchent à clarifier la dynamique du mécanisme de transmission de la politique monétaire. L'identification des chocs de politique monétaire est au coeur de ce genre d'analyse. Divers chercheurs ont réussi récemment à isoler de tels chocs en imposant des restrictions de suridentification à des modèles vectoriels autorégressifs (VAR). Les auteurs de l'étude adoptent une démarche différente pour examiner les chocs monétaires survenus dans six pays industrialisés : afin d'isoler ces chocs, ils imposent des restrictions de cointégration à long terme à des modèles VAR et postulent que la monnaie est neutre en longue période. Selon les résultats qu'ils obtiennent, la dynamique à court terme qui caractérise un choc monétaire peut être interprétée comme un choc de politique monétaire. Le stock de monnaie serait par conséquent un rouage important du mécanisme de transmission de la politique monétaire.

CONTENTS

1. INTRODUCTION

Uncertainty about the monetary transmission mechanism plagues academics and central bankers alike. Econometric models that attempt to resolve this uncertainty have taken several forms. One of the primary issues of such models is the identification of monetary policy shocks. Since the seminal work of Sims (1980), vector autoregression (VAR) models have become an increasingly popular tool for empirical studies of the transmission mechanism. Unlike large-scale macroeconometric models, this methodology primarily uses necessary restrictions to achieve identification. In this regard, the initial VAR literature attempted to identify monetary policy shocks assuming a recursive-causal structure among key macro-variables such as money, interest rates, output and prices. Under such assumptions, however, VAR studies have been subject to two recurring puzzles. Immediately following an expansionary monetary policy shock, the interest rate increases rather than decreases (the “liquidity puzzle”). Second, also following an unanticipated monetary expansion, the price level initially decreases rather than increases (the “price puzzle”). These puzzles appear to depend critically on the restrictions used to identify monetary policy shocks.

The liquidity puzzle is often found in VAR models that measure monetary policy shocks by the orthogonalized innovation in conventional monetary aggregates (see Sims 1986). Two ways to resolve this puzzle have been suggested. First, conventional monetary aggregates may not be an appropriate measure of monetary policy, since their movements are the combination of both private and central banking behaviour. Second, measures of monetary policy that are under the direct influence of the central bank, such as short-term interest rates and non-borrowed reserves (NBR) in the United States, are more appropriate measures of policy (see Bernanke and Blinder 1992; Christiano and Eichenbaum 1992). Using short-term interest rates, several VAR studies find that a positive interest rate innovation leads to an initial fall in money, consistent with the liquidity effect.[1] Similarly, using bank reserves as the measure of policy, such as NBR, an expansionary policy shock initiates a fall in the interest rate (see Christiano and Eichenbaum 1992; Strongin 1995).

While measuring monetary policy using operational variables may resolve the liquidity puzzle, studies using both short-term interest rates and very narrow monetary aggregates remain susceptible to the price puzzle. For example Sims (1992), using a short-term interest rate across industrialized countries, and Christiano, Eichenbaum and Evans

1. For cross-country studies see Sims (1992) and Gerlach and Smets (1994). For a study of Canada, see Armour, Engert and Fung (1996).

(1994), using NBR in the United States, find the price puzzle. For Canada, Fung and Gupta (1994) encounter the price puzzle using excess cash reserves and a recursive Wold ordering to identify a monetary policy shock.

The deeper root of the puzzles in the initial VAR literature relates to the difficulty of isolating a monetary policy shock. One source of this difficulty is the recursive-causal assumptions, criticized by Cooley and Leroy (1985), among others. A second source is that the four or five variable central-bank reaction functions are over-simplified. In response to these difficulties, recent studies have relaxed the recursive-causal assumptions and included additional information in the reaction function, thereby increasing the likelihood of distinguishing a monetary policy shock from a central-bank reaction to movements in other economic variables. Gordon and Leeper (1994), for example, use contemporaneous restrictions to separate money demand and monetary policy shocks, and include additional “world-economy shocks,” such as commodity price shocks, to resolve both the liquidity and price puzzles for the United States. For the G-7 countries, Grilli and Roubini (1996) find similar results.[2] The main difference of the approach used in these recent papers from the previous literature is its reliance on over-identifying restrictions.

One commonality among the papers discussed above is the use of contemporaneous restrictions. An alternative approach to the problem of identification uses long-run propositions of economic theory, which are arguably less controversial than their contemporaneous counterpart.[3] Moreover, this branch of literature has not relied on over-identifying restrictions. For example, Lastrapes and Selgin (1995) and Kasumovich (1996) identify monetary policy shocks by relying on the assumption of long-run money neutrality without introducing over-identifying restrictions. In their simple four-variable models, which include conventional monetary aggregates, an expansionary monetary shock leads to an initial decrease in the interest rate and an increase in the price level.

In this paper, simple cointegrated VAR models for each country consisting of a conventional monetary aggregate, a price level, output and a short-term interest rate are estimated. A monetary shock is identified by the assumption that a permanent change in the nominal stock of money has a proportionate effect on the price level with no long-run real economic consequences; that is, a monetary shock is neutral in the long run. This

2. For additional detail, see Kim and Roubini (1995), Kim (1995), and Grilli and Roubini (1995).

3. In contrast, Faust and Leeper (1994) present an extensive discussion concerning the potential problems of using long-run restrictions.

differs from previous cross-country studies that examine the effects of monetary policy isolated using over-identifying contemporaneous restrictions.[4]

To preview the results, we find that the short-run dynamics of a monetary shock can be interpreted as a monetary policy shock across countries. We consistently find liquidity effects and no price puzzles. An expansionary shock leads to an increase in the money stock, a temporary decline in the interest rate, a temporary rise in real output, and an impact increase in the price level. These results give support to the appropriateness of identifying monetary policy shocks using long-run restrictions and suggest that focussing on the operational aspects of monetary policy may not be necessary in a VAR context.

The paper is organized as follows. Section 2 outlines the empirical methodology for identification of the monetary shock. Section 3 summarizes the data and their cointegration properties. Next, the parameters of a single cointegration vector are estimated for each country. This vector is interpreted as a long-run demand-for-money function. In Section 4, we present impulse-response functions for the monetary shocks and examine the speed at which the models reach equilibrium. Section 5 summarizes the key results and suggests possible extensions for future research.

4. For the case of New Zealand, Fischer, Fackler and Orden (1995) also use long-run cointegration restrictions to identify the monetary policy component in a broad aggregate. Unlike these authors, we attribute the fluctuations in output and prices to interest rates and examine monetary shocks across industrialized countries.

2. EMPIRICAL METHODOLOGY

In this section, we outline the empirical methodology and show how the long-run propositions of economic theory are used to identify economic shocks.[5]

2.1 THE IDENTIFIED VAR

After estimating a cointegrated VAR model (an error-correction model), we represent the time series of economic variables as a function of past disturbance terms:

$$\Delta X_t = I_n e_t + G_1 e_{t-1} + G_2 e_{t-2} + \ldots + G_\infty e_{t-\infty} \qquad (1)$$

where X_t is an nx1 vector of time-t economic variables, Δ is the first-difference operator, G_i is an nxn matrix of known parameters, and e_t is an nx1 vector of time-t reduced-form residuals (one-period forecast errors). Model (1) is referred to as the reduced-form moving-average representation (MAR) of the cointegrated VAR model. In order to give a structural interpretation to the historical errors in the economic variables, we specify the following structural moving-average representation (SMAR):

$$\Delta X_t = \Phi_0 \varepsilon_t + \Phi_1 \varepsilon_{t-1} + \Phi_2 \varepsilon_{t-2} + \ldots + \Phi_\infty \varepsilon_{t-\infty} \qquad (2)$$

where Φ_i is an nxn unknown polynomial matrix, and ε_t is an unknown nx1 vector of time-t structural (or economic) shocks. The MAR and SMAR are related by

$$\Phi(L)\Phi_0^{-1} = G(L) \qquad (3)$$

$$\Phi_0 \varepsilon_t = e_t \qquad (4)$$

defining $G(L)$ equal to $I_n L^0 + G_1 L^1 + \ldots + G_l L^l$ and $\Phi(L)$ equal to $\Phi_0 L^0 + \Phi_1 L^1 + \ldots + \Phi_l L^l$, where L is the lag operator and l is the truncated lag length.[6] Equation (3) relates the parameters of the structural and reduced-form representations through the matrix of contemporaneous relationships, Φ_0. From equation (4), a reduced-form residual can be expressed as a combination of underlying economic shocks. By identifying the elements of Φ_0, the dynamic relationship between the structural shocks and the economic variables can be examined through model (2). We distinguish between

5. A more detailed description of this methodology can be found in King, Plosser, Stock and Watson (1991), Fischer, Fackler and Orden (1995), and Gonzalo and Ng (1996).

6. In the empirical results, an arbitrarily large truncation is used (300 quarters).

temporary and permanent shocks. Unlike a temporary shock, a permanent shock has a long-run impact on at least one of the economic variables (for example, a monetary policy shock that has a permanent effect on the money stock and the price level).

To identify the structural model from the estimated model, restrictions must be imposed on the structural model. From equation (4), $\Phi_0\Sigma_\varepsilon\Phi_0' = \Sigma_e$. There are $n(n+1)/2$ unique known elements in Σ_e, estimated from model (1). However, there are n^2 elements in Φ_0 and $n(n+1)/2$ unique elements in Σ_ε that are unknown. Thus, an additional n^2 restrictions are sufficient to identify model (2).

The first component of the identification strategy assumes that permanent and temporary shocks originate from independent sources. The covariance matrix of the structural shocks is therefore

$$\Sigma_\varepsilon = E(\varepsilon_t\varepsilon_t') = \begin{bmatrix} \Sigma_{\varepsilon^P} & 0 \\ 0 & \Sigma_{\varepsilon^T} \end{bmatrix} \qquad (5)$$

and Σ_ε is partitioned conformably with $\varepsilon_t = (\varepsilon_t^{P}{}' \ \ \varepsilon_t^{T}{}')'$, where ε_t^P is the kx1 vector of permanent shocks, ε_t^T is the rx1 vector of temporary shocks, and $k+r$ is equal to n. Assuming that the structural shocks with permanent effects are mutually orthogonal and the structural shocks with temporary effects are mutually orthogonal, the assumptions associated with (5) generate $n(n-1)/2$ unique identification restrictions.

The second component of the identification strategy imposes cointegration constraints on the matrix of long-run multipliers $\Phi(1)$. Note that from the condition that the nxr matrix of cointegration relationships (β) is orthogonal to the matrix of long-run multipliers ($\beta'\Phi(1) = 0$; see Engle and Granger 1987), there are only nxk independent unknown elements in $\Phi(1)$. For simplicity, we partition the matrix of long-run multipliers by the number of permanent shocks in the model. The first k columns of the matrix represent the long-run response of the change in X_t to the k permanent shocks. The long-run response of the change in X_t to the temporary shocks is represented by the last r columns and is equal to zero by definition.[7] Specifically, setting $L = 1$, model (2) is rewritten as

7. Since there are r stationary linear combinations among the non-stationary economic variables, the structural model has r temporary shocks.

$$\Delta X_t = [A \;\; 0]\varepsilon_t \tag{6}$$

where $\Phi(1) = [A \;\; 0]$, the long-run multipliers of the permanent shocks are summarized by the $n \times k$ matrix A, and the long-run multipliers of the temporary shocks are summarized by the $n \times r$ matrix of zeros. The zero restrictions associated with this partition generate $r \times k$ unique identification restrictions.

The parameters of the cointegration vectors are used to restrict the long-run multipliers of the permanent shocks. Specifically, $A = \tilde{A}\Pi$, where the $n \times k$ matrix $\tilde{A}$ is a known matrix of "cointegration structure" whose components are determined by the condition that its columns are orthogonal to the matrix of cointegration relations ($\beta'\tilde{A} = 0$). The matrix Π is a $k \times k$ matrix with full column rank. We assume that Π is lower triangular and normalize the diagonal elements to 1's (a long-run Wold ordering). As a result, we obtain the additional $k(k+1)/2$ identification restrictions necessary to uniquely identify the permanent shocks in the model.[8] In this sub-section, we illustrate the identification scheme for a simple monetary model.

2.2 A SIMPLE MONETARY MODEL

Consider a four-variable model consisting of the nominal stock of money (in logs, m), the price level (in logs, p), a nominal interest rate (R), and real income (in logs, y). Suppose that, in their univariate representations, the variables are subject to a stochastic (non-stationary) trend. According to standard monetary models, we can express a long-run demand-for-money function as $m_t^d = \mu + p_t + \beta_1 y_t - \beta_2 R_t + \varepsilon_t^T$, where μ is a constant term, β_1 is the income elasticity of the demand for money, β_2 is the interest rate semi-elasticity, and ε_t^T is a money-demand shock.[9] If the money-demand shock is stationary, then the variables are cointegrated. In this case, since the variables share a stochastic trend, in a multivariate representation there is one temporary shock (the money-demand shock) and three permanent shocks (three independent stochastic trends). We interpret the permanent shocks below.

Defining $X = \begin{bmatrix} y & R & p & m \end{bmatrix}'$, the 4×3 matrix A can be expressed as

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8. In the case where there is one permanent shock ($k = 1$), Π is a scalar and therefore redundant. However, when more than one permanent shock is present ($k > 1$), the permanent shocks are non-unique. In other words, for any non-singular matrix P, $(AP)(P^{-1}\varepsilon_t) = A\varepsilon_t$.

9. More recently, such a function has been derived from general-equilibrium models that include money in a representative agent's utility function (see Kim 1995).

$$A = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \\ \beta_1 & -\beta_2 & 1 \end{bmatrix} \begin{bmatrix} 1 & 0 & 0 \\ \pi_{21} & 1 & 0 \\ \pi_{31} & \pi_{32} & 1 \end{bmatrix} \tag{7}$$

where $\beta' = \begin{bmatrix} -\beta_1 & \beta_2 & -1 & 1 \end{bmatrix}$. As discussed above, in conjunction with Π, the matrix of cointegration structure $\tilde{A}$ is used to interpret the permanent shocks. According to the first column of $\tilde{A}$, a 1 per cent increase in level of output has a positive effect on the demand for money of β_1. Combined with Π, a permanent output shock can have a permanent effect on all (real) variables. This shock can be interpreted as a productivity shock. The second column of $\tilde{A}$ represents a 1 per cent permanent real interest rate shock. In other words, a 1 per cent increase in real interest rates has a negative effect on the demand for money of β_2. Combined with Π, a permanent real interest rate shock also can have a permanent effect on all variables other than output. This shock can be interpreted as either a foreign interest rate shock or a risk-premium shock.

Our focus is on the third column, which can be interpreted as a monetary policy shock. According to this column, a monetary shock is one that leads to a proportionate change in the stock of money and the price level with no long-run real economic effects.[10] In other words, the four-variable system has three independent stochastic trends as represented by the demand-for-money function. Abstracting from the output and interest rate trends, “monetary shocks” are defined as the innovation in the common trend of money and prices.

10. Note that with respect to the monetary shock, specifying $X' = \begin{bmatrix} y & R & m & p \end{bmatrix}$ is equivalent to $X' = \begin{bmatrix} y & R & p & m \end{bmatrix}$ with the appropriate adjustments to the matrix $\tilde{A}$. That is, inflation targeting is equivalent to monetary targeting in the long run.

3. PROPERTIES OF THE DATA

3.1 DESCRIPTION OF THE DATA

The six industrialized countries examined in this study are Canada, France, Germany, Japan, the United Kingdom, and the United States. The data used are quarterly observations and are summarized in Table 1.[11] The measure of the nominal stock of money is defined as the currency outside banks plus demand deposits, the conventional definition of M1. The measure of the nominal interest rate is the money market rate of interest. The measure of the price level is the consumer price index. Depending on the availability of data, the level of real income is measured by either GNP, GDP or Industrial Production.

Table 1. Summary of the data set

Country	*m*	*p*	*y*	*R*	Sample
Canada	M1	CPI	GDP	overnight	1954Q4-1995Q4
France	M1	CPI	IP	overnight	1957Q1-1995Q1
Germany	M1	CPI	GNP	overnight	1960Q1-1988Q4
Japan	M1	CPI	GNP	overnight	1960Q1-1993Q1
United Kingdom	M1	CPI	GDP	90-day bill	1957Q1-1994Q1
United States	M1	CPI	IP [a]	overnight	1957Q1-1995Q2

a. For the United States, we also considered GDP and GNP. However, in both cases the impulse-response functions were anomalous. Although the initial responses were consistent with a monetary policy shock, estimates of their variances were unusually large.

There are at least two reasons that lead us to use a conventional monetary aggregate in our study. The first reason relates to our identification strategy. In this paper, shocks are interpreted from the cointegration relationship among M1, a price level, output and a short-term interest rate. It is possible to interpret the estimated cointegration relationship as a long-run demand-for-money function. A key issue of such an analysis is

11. Unless noted otherwise, the data are constructed using *International Financial Statistics*, published by the International Monetary Fund. The data are seasonally adjusted in RATS version 4.1. Note the following exceptions in the data. The data for Canada are from the Cansim database. The overnight interest rate in Canada is constructed as the average of the day-to-day loan rate prior to 1971Q2, and the average Canadian overnight financing rate after 1971Q2. For the United Kingdom, the definition of money includes building societies from 1987Q1. Italy was excluded from the sample of countries primarily because a money-market interest rate was not readily available prior to 1971Q1.

the temporal stability of the parameters of the demand-for-money function. A stable function implies that while the stock of money may not be equal to its long-run desired level at any moment in time, agents alter their expenditure decisions to attain a unique level of money holdings. That is, deviations of money from its long-run demand are inherently transitory, and the monetary authority can achieve a stable path for prices by targeting a stable path of monetary expansion.

The second reason for using a conventional monetary aggregate in our study relates to cross-country comparisons. Although it may be more desirable to use a monetary instrument, such as NBR in the United States, the fact that there have been considerable changes in reserve requirements in most countries is inherently problematic. In addition, although there are considerable differences in the institutional setting and implementation of monetary policy across countries, we attempt to determine whether the effects of monetary policy shocks are common across countries. Since we are interested in using a common framework to study the effects of unanticipated monetary policy across industrialized countries, we try to use a similar data set, sample period, and framework for all countries, as far as allowed by the availability of data.

3.2 EVIDENCE OF COINTEGRATION

In a multivariate VAR model, we conduct tests to determine the presence of unit roots.[12] If in a four-variable model the variables are not cointegrated, then four unit roots (independent stochastic trends) should emerge from the data. In this regard, two tests for cointegration are considered: the trace test and the lambda-max test (see Johansen and Juselius 1990 for a description of the tests).

When the number of variables combined with the lag specification is “large” relative to the size of the data set, there can be a substantial small-sample bias toward finding cointegration (for example, see Gonzalo 1994). To address this issue, we based our conclusion of cointegration on finite-sample critical values, which are reported in footnote

12. Two univariate unit-root tests were examined: the augmented Dickey-Fuller test and the $MZ\alpha$ test developed by Stock (1990). Other than the interest rate in Japan, the results strongly suggested that the variables are non-stationary across countries. However, the null hypothesis of a unit root could not be rejected for either the price level (in logs) or the first difference of the price level (inflation). The unit-root properties of the money supply (for France and Japan) were similar to those of the price level.

a of Table 2.[13] Table 2 presents the test statistics for the null hypothesis of no cointegration. For all six countries, evidence of at least one cointegration vector emerges from the data. That is, among the four economic variables, there is evidence of a common stochastic trend. Given the evidence in favor of cointegration, the following section estimates the parameters of the cointegration vectors and tests the temporal stability of these parameters.

Table 2. Test statistics for the trace and lambda-max multivariate unit-root tests.

Countries	Test statistics[a]	
	λ max	Trace
Canada	30.25	73.22
France	39.53	86.70
Germany	32.40	66.96
Japan	31.79	88.65
United Kingdom	34.76	67.68
United States	59.32	95.99

a. The null hypothesis is no cointegration ($r = 0$, $k = 4$). The finite-sample critical values for the case of a restricted constant are 29.17 (5 per cent significance level) and 34.66 (1 per cent significance level) for the λ-max test and 55.00 (5 per cent significance level) and 62.15 (1 per cent significance level) for the trace test. For the case of an unrestricted constant, the finite-sample critical values are 27.97 (5 per cent significance level) and 33.15 (1 per cent significance level) for the λ-max test, and 48.42 (5 per cent significance level) and 55.52 (1 per cent significance level) for the trace test. (See footnote 13 for a discussion of the methodology used to generate the finite-sample critical values.)

Notes: The models are estimated with three centered seasonal dummy variables. Depending on which specification supports fewer cointegration vectors, the constant term is estimated as either restricted (in the cointegration space) or unrestricted (in the conditional mean). A sequential lag-length test (likelihood-ratio test) was used to determine the optimal number of lags for the VAR in levels. The specifications of the models for each country are as follows: Canada is estimated with six lags and a restricted constant; France is estimated with six lags and a restricted constant; Germany is estimated with eight lags and an unrestricted constant; Japan is estimated with six lags and a restricted constant; the United Kingdom is estimated with nine lags and a restricted constant; and the United States is estimated with six lags and a restricted constant.

3.3 INTERPRETATION OF THE COINTEGRATION VECTORS

This sub-section describes the parameters of a single cointegration vector estimated for each country with the multivariate cointegration techniques developed by

13. The following methodology was used to generate these critical values. For a sample size of 150 observations, the non-deterministic parts of the data-generating process were drawn from a mean zero Gaussian distribution with variance one. The constant term was drawn from a mean ten Gaussian distribution with variance one. Then, the null hypothesis of no cointegration was tested according to the Johansen methodology ($n - r$ stochastic trends). Upon replicating this procedure 12,000 times, the finite-sample quantiles were calculated. This methodology is simply a finite-sample version of Osterwald-Lenum (1992) in which asymptotic critical values were generated.

Johansen and Juselius (1990). This vector is interpreted as a long-run demand-for-money function. Then we test for the temporal stability of the demand these parameters, since this has important implications for the conduct of monetary policy.

After estimating a vector error-correction model, the cointegration space is restricted to one cointegration vector. Then the joint hypothesis of unitary price and income elasticity is tested. If this hypothesis is not rejected, the velocity of M1 is cointegrated with the nominal interest rate. If this hypothesis is rejected, the hypothesis of unitary price elasticity is then tested. In any case, unitary price elasticity will be maintained for those countries in which the hypothesis is rejected. (In terms of a monetary shock, which will be analysed in the next section, unitary price elasticity is a component of the assumption that monetary policy is neutral in the long run.)

Table 3. Parameter estimates of the cointegration vectors

Countries	P-values (for null hypothesis)[a]		Estimated cointegration vectors
	unitary p and y elasticity	unitary p elasticity	
Canada	0.120	----	$m = p + 1.00y - 0.12R + 1.75$
France	0.013	0.005	$m = p + 0.40y - 0.05R - 1.54$
Germany	0.275	----	$m = p + 1.00y - 0.27R$
Japan	0.012	0.483	$m = p + 0.85y - 0.03R + 10.67$
United Kingdom	0.341	----	$m = p + 1.00y - 0.57R - 7.48$
United States	0.000	0.001	$m = p + 0.50y - 0.03R + 0.33$

a. Johansen (1991) shows that likelihood-ratio tests for linear restrictions on the cointegration vector are asymptotically distributed chi-squared where the degrees of freedom are equal to the number of restrictions.

Notes: The regressions are described in the notes to Table 2.

Table 3 summarizes the cointegration results. For Canada, Germany and the United Kingdom, the joint hypothesis of unitary price and income elasticity cannot be rejected at reasonable significance levels. For Japan, the joint hypothesis can be rejected, but unitary price elasticity cannot. Both hypotheses are rejected for France and the United States. In terms of identifying monetary policy shocks in these two countries, the restriction that a long-run change in the stock of money has a proportionate effect on the price level should be used with caution. The parameter estimates of the cointegration vectors are, however, broadly consistent with a demand-for-money function. The interest rate semi-elasticities are all negative and range in value from -0.03 (Japan) to -0.57 (United Kingdom). All of the countries that reject unitary income elasticity (France, Japan and the United States) have elasticities of less than one.

Next we examine the temporal stability of the parameters of the cointegration vectors. The test used is as follows. First, we define the full sample estimation as $[ef, el]$. Holding ef fixed, all of the parameters of the vector error-correction model are re-estimated for each quarter between the subsamples $[ef, 1975Q1]$ and $[ef, el]$. Note that the cointegration space is restricted to one vector for each subsample. Then we test whether the parameters of the cointegration vector estimated in the subsample are statistically different from those estimated in the full sample.[14] The null hypothesis is that the subsample parameter estimates of the cointegration vector are equal to the full-sample parameter estimates. This is in contrast to Hoffman, Rasche and Tieslau (1995), who focus on the stability of the individual estimated money-demand parameters. We then compare the test statistics with the critical values according to Andrews (1993).[15]

For the first four to twelve quarters, the constant-parameter null hypothesis was rejected (not shown). We attribute this rejection to the small sample. After this period, for Germany, Japan and the United Kingdom, the subsample parameters of the cointegration vector were not statistically different from the full-sample parameters. For Canada, the null hypothesis could not be rejected at the 10 per cent significance level. On the other hand, for both France and the United States, the constant-parameter null hypothesis was strongly rejected over time. However, for the United States, the constant-parameter null hypothesis could not be rejected when the restriction of unitary price elasticity was relaxed. For France, relaxing the assumption of unitary price elasticity, the null hypothesis was rejected from 1980Q1 to 1981Q1. These results suggest that the restriction of unitary price elasticity could be the source of money-demand instability for France and the United States.

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14. The test statistic for this test is $\tau[\ln(1-\hat{\rho}(\tau)) - \ln(1-\hat{\lambda}(\tau))]$ for the case where the cointegration space is restricted to one vector. τ represents the number of observations for the subsample; $\hat{\rho}(\tau)$ is the eigenvalue associated with the subsample (conditional on the full sample cointegration vector); $\hat{\lambda}(\tau)$ is the largest eigenvalue for the full sample. This test statistic is distributed chi-squared with $(p-r)r$ degrees of freedom, where p represents the dimensions of the cointegration vector (for models estimated with an unrestricted constant $p = 4$ and a restricted constant $p = 5$). See Hansen and Juselius (1995) for details and references.

15. The critical values used are from Table 1 of Andrews (1993). These critical values remove the bias against the null hypothesis that exists when testing parameter instability in a rolling regression using standard critical values.

4. EMPIRICAL RESULTS

With the parameter estimates of the demand-for-money function established, restrictions to the matrix of long-run multipliers will now allow us to perform a dynamic analysis of the permanent innovations in the system. In this section we present the impulse-response functions for the monetary shocks and examine the speed at which the models reach equilibrium.[16] The following experiment is conducted: an unanticipated contemporaneous increase in the money stock that generates a proportionate permanent 1 per cent change in the stock of money and the price level, with no long-run impact on interest rates or output.

Figure 1 illustrates the impulse-response functions and their one-standard deviation confidence bands for the six countries.[17] The first row is the response of money. We interpret these response functions as a central-bank action unanticipated by agents that generates a permanent increase in M1. The initial response of the money stock differs across countries but is less than its long-run equilibrium, except in Germany. This suggests the presence of a possible "multiplier effect" that varies across countries. One reason for this variation could be the substitution out of M1 into interest-bearing assets. Note that the countries estimated to have the lowest degree of M1 substitutability also have the largest relative long-run increase in the money stock: Japan and the United States, with an interest rate semi-elasticity of -0.03. In contrast, the United Kingdom, the country estimated to have the highest degree of M1 substitutability with an interest rate semi-elasticity of -0.57, has the lowest long-run increase in M1 relative to its contemporaneous change.

For all countries, the interest rate responses follow a similar pattern. Consistent with the liquidity effect, the impact response of the interest rate is negative following the expansionary monetary shock. This effect is significantly different from zero for Canada, Germany, the United Kingdom and the United States. For Germany, the volatile interest rate response may appear counterintuitive, given the relatively stable interest rates this country has had.[18] However, this result merely suggests that monetary shocks may not have been a major source of interest rate fluctuations in Germany. Note also that the magnitudes of the initial interest rate responses. In countries with relatively low interest rate semi-elasticities, the initial response of both the interest rate and the money stock

16. The VAR models were estimated with six quarterly lags across all countries.

17. A Monte Carlo simulation was used to estimate the standard deviation of the response functions.

18. Over the sample, the variance of the money-market interest rate in Germany is 6.22 compared with 11.85 for the United States.

tends to be small (Japan, France). In contrast, there appears to be a relatively large initial response for both the interest rate and the money stock in countries with relatively high interest rate semi-elasticities (Canada, the United Kingdom).

After the initial fall in the interest rate, for all countries other than the United States, the response functions increase above zero and then converge to steady state. This is consistent with the view that, over a shorter horizon, a liquidity effect dominates an inflation effect on the nominal interest rate while, over a longer horizon, the reverse is true. For Germany, the inflation effect is quite strong but short-lived. This may be attributed to the relatively fast equilibrium adjustment of the price level.

As expected following an expansionary monetary policy shock, the output response is positive across countries, although briefly negative for Canada and Germany. Excluding France, the increase in output peaks seven to ten quarters after the monetary shock. This positive response is significant for France, Germany, the United Kingdom and the United States. For the case of Germany, the variability of the output response is far greater than the other countries. Von Hagen (1995) attributes part of the variability of output in Germany to the fact that the central bank attempted to offset the negative supply shocks in 1973-74 and 1980-81 even though the Bundesbank’s objective was specified in terms of inflation rather than the price level.

In addition to the response functions, we generated variance decomposition functions to assess the real effects of monetary policy. In general, the variability of output attributable to the monetary shock was quite small at all horizons. For example, for Canada, monetary shocks explained, at most, 10 per cent of output fluctuations (over a 28-quarter horizon). In contrast, as predicted by real-business-cycle models, productivity shocks explained approximately two-thirds of output variability.

Finally, consider the response of the price level to the monetary shock. For each country, the initial response of the price level is positive but small. For Germany, consistent with the dominant inflation effect that is observed in the interest rate response function, the price level adjusts more quickly compared with the other countries. Also for Germany, a slight overshooting of the price level occurs; this is consistent with the overshooting of money. What is most clear across countries, however, is the lagged equilibrium adjustment of the price level in response to a monetary shock.

The responses are generally consistent with the view that the central banks influence the money stock independent of other factors in the economy; that is, money has an active role in the transmission of monetary policy. Specifically, consider the equilibrium adjustment of the models, which is quite slow.[19] This suggests that, in response to an unanticipated increase in money, in the aggregate it takes several quarters for agents to rebalance their portfolios. Since the price level is slow to adjust because of, for example, menu costs, the subsequent temporary increase in spending leads to an increase in real economic activity. As the price level adjusts, monetary equilibrium is restored, and interest rates and output return to their pre-shock levels.

19. The equilibrium adjustment is quickest in the case of Germany. This could be explained by the fact that Germany is the only country that has targeted a monetary aggregate over most of the sample period.

5. SUMMARY AND CONCLUSIONS

This paper presents new evidence on the transmission of monetary policy shocks across industrialized countries. Monetary shocks are identified as those that have a proportionate effect on the stock of money and the price level but with no long-run impact on output or the interest rate. The response functions generally suggest that the monetary shock identified can be interpreted as a monetary policy shock. An expansionary shock generates an increase in the stock of money, a short-run fall in the interest rate, a temporary rise in output, and an impact increase in the price level. Unlike previous literature on short-run dynamics of monetary policy shocks, we do not rely on over-identifying restrictions. In other words, the results presented in this paper are not dependent on somewhat arbitrary instrumental variables.

In general, the results of this study support the view that the narrow stock of money has an active role in the transmission of monetary policy. Such results are consistent with a broad class of theories, such as the buffer-stock model and more recent general-equilibrium models with both real and nominal rigidities. The results also suggest that monetary policy has a limited influence on real variables, such as real output. For the purposes of nominal targeting, however, the results are quite promising. In particular, across countries, the estimated demand-for-money function is relatively stable (aside from the price elasticity in France and the United States) and monetary shocks have a significant effect on the nominal equilibrium path in the economy, namely the path of money and prices. This suggests that one avenue for controlling the evolution of the price level may be through monetary targeting.

With the goal of generating a richer set of dynamics, several extensions of the basic modeling strategy could be investigated in future research. Explicit open-economy factors could be examined, most notably the inclusion of an exchange rate. The relationship between monetary policy and the term structure of interest rates would also be an interesting avenue for future research.

REFERENCES

Andrews, D. 1993. "Tests for Parameter Instability and Structural Change with Unknown Change Point." *Econometrica*, 61, 821-856.

Armour, J., W. Engert, and B. S. C. Fung. 1996. "Overnight Rate Innovations as a Measure of Monetary Policy Shocks." Working Paper 96-4, Bank of Canada.

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

Christiano, L. and M. Eichenbaum. 1992. "Identification and the Liquidity Effect of a Monetary Policy Shock." In *Political Economy, Growth, and Business Cycles*, edited by A. Cukierman, L. Hercowitz and L. Leiderman (MIT Press, Cambridge, MA), 335-370.

Christiano, L., M. Eichenbaum and C. Evans. 1994. "The Effects of Monetary Policy Shocks: Evidence from the Flow of Funds." Working Paper WP-94-2, Federal Reserve Bank of Chicago.

Cooley, T. and S. Leroy. 1985. "Atheoretical Macroeconometrics: A Critique." *Journal of Monetary Economics*, 16, 283-308.

Faust, J. and E. Leeper. 1994. "When Do Long-Run Identifying Restrictions Give Reliable Results?" International Finance Discussion Paper No. 62, Board of Governors of the Federal Reserve System.

Fisher, L., P. Fackler and D. Orden. 1995. "Long-Run Identifying Restrictions for an Error-Correction Model of New Zealand Money, Prices and Output." *Journal of International Money and Finance*, 14, 127-147.

Fung, B. S. C. and R. Gupta. 1994. "Searching for the Liquidity Effect in Canada." Working Paper 94-12, Bank of Canada.

Gerlach, S. and F. Smets. 1994. "The Monetary Transmission Mechanism: Evidence from the G-7 countries." Unpublished, Bank for International Settlements, October.

Gonzalo, J. 1994. "Five Alternative Methods of Estimating Long Run Equilibrium Relationships." *Journal of Econometrics*, 60, 203-233.

Gonzalo, J. and S. Ng. 1996. "A Systematic Framework for Analyzing the Dynamic Effects of Permanent and Transitory Shocks." Report 0396, University of Montreal, Centre of Research and Economic Development, March.

Gordon, D. and E. Leeper. 1994. "The Dynamic Impacts of Monetary Policy: An Exercise in Tentative Identification." *Journal of Political Economy*, 102, 1228-1247.

Grilli, V. and N. Roubini. 1995. "Liquidity and Exchange Rates: Puzzling Evidence from the G-7 Countries." Unpublished, Yale University, March.

Grilli, V. and N. Roubini. 1996. "Liquidity Models in Open Economies: Theory and Empirical Evidence." *European Economic Review*, 4, 847-859.

Hansen, H. and K. Juselius. 1995. *CATS in RATS: Cointegration Analysis of Time Series*. 1-87.

Hoffman, D., R. Rasche and M. Tieslau. 1995. "The Stability of Long-Run Money Demand in Five Industrial Countries." *Journal of Monetary Economics*, 35, 317-339.

Johansen, S. 1991. "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models." *Econometrica*, 59, 1551-1580.

Johansen, S. and K. Juselius. 1990. "Maximum Likelihood Estimation and Inference on Cointegration—With Applications to the Demand for Money." *Oxford Bulletin of Economics and Statistics*, 52, 169-210.

Kasumovich, M. 1996. "Interpreting Money-Supply Shocks and Interest Rate Shocks as Monetary-Policy Shocks." Working Paper 96-8, Bank of Canada.

Kim. S. 1995. "Does Monetary Policy Matter in the G-6 Countries? Using Common Identifying Assumptions about Monetary Policy across Countries." Unpublished, Yale University, November.

Kim, S. and N. Roubini. 1995. "Liquidity and Exchange Rates: A Structural VAR Approach." Unpublished, Yale University, February.

King, R., C. Plosser, J. Stock and M. Watson. 1991. "Stochastic Trends and Economic Fluctuations." *American Economic Review,* 81, 819-840.

Lastrapes, W. and G. Selgin. 1995. "The Liquidity Effect: Identifying Short-Run Interest Rate Dynamics Using Long-Run Restrictions." *Journal of Macroeconomics*, 17, 387-404.

Leeper, E. and D. Gordon. 1992. "In Search of the Liquidity Effect." *Journal of Monetary Economics*, 29, 341-369.

Osterwald-Lenum. M. 1992. "A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics." *Oxford Bulletin of Economic and Statistics*, 54, 461-472.

Sims, C. 1980. "Macroeconomics and Reality." *Econometrica*, 48, 1-48.

Sims, C. 1986. "Are Forecasting Models Usable for Policy Analysis?" *Federal Reserve Bank of Minneapolis Quarterly Review*, 10, 2-16.

Sims, C. 1992. "Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy." *European Economic Review*, 36, 975-1000.

Stock, J. 1990. "A Class of Tests for Integration and Cointegration." Unpublished, Kennedy School of Government, May.

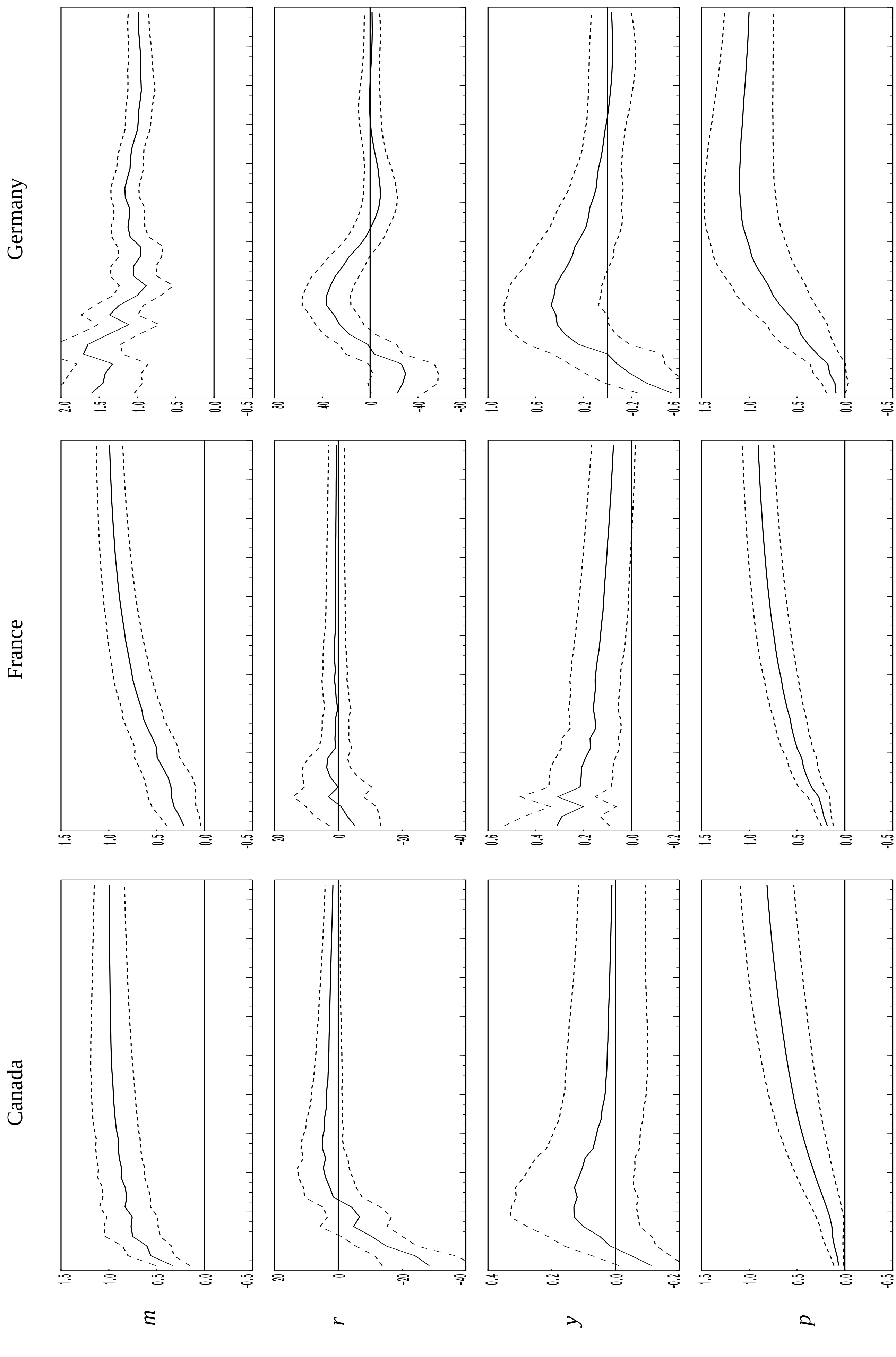
Strongin, S. 1995. "The Identification of Monetary Policy Disturbances: Explaining the Liquidity Puzzle." *Journal of Monetary Economics*, 35, 463-497.

Von Hagen, J. 1995. "Inflation and Monetary Targeting in Germany." In *Inflation Targets*, edited by L. Leiderman and L. Svensson.

APPENDIX 1: FIGURES

FIGURE 1

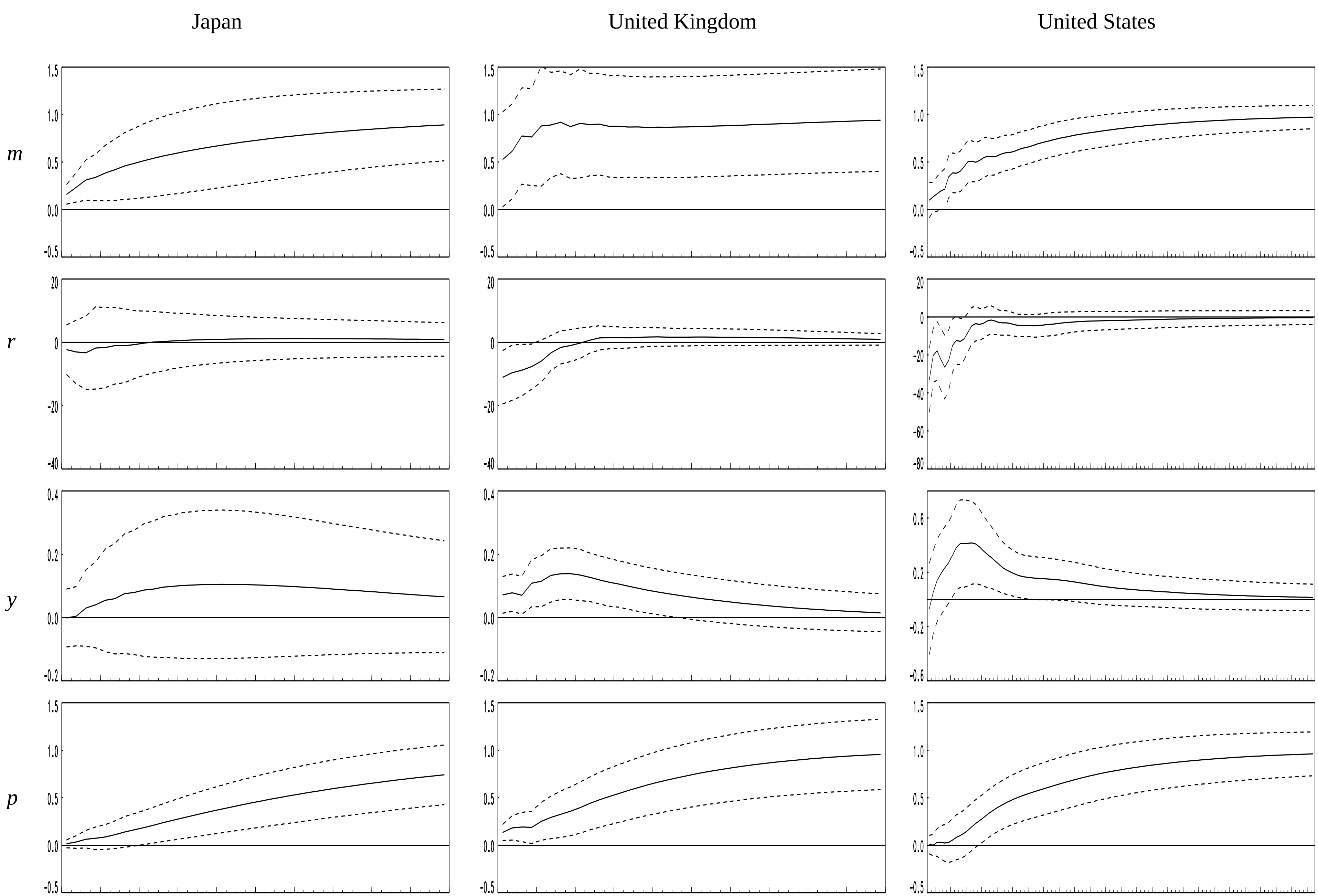
Dynamic responses to a monetary shock (40 quarters)



Notes: The interest rate response function is in terms of basis points. All other responses are in percentage terms. The (one-standard deviation) error bands are generated from a Monte Carlo simulation.

FIGURE 1 (cont.)

Dynamic responses to a monetary shock (40 quarters)



Notes: The interest rate response function is in terms of basis points. All other responses are in percentage terms. The (one standard deviation) error bands were generated from a Monte Carlo simulation.

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