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International Transmission of Inflation among G-7 Countries: A Data-Determined VAR Analysis

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Abstract

We investigate the international transmission of inflation among G-7 countries using data-determined vector autoregression analysis, as advocated by Swanson and Granger (1997). Over the period 1973 to 2003, we find that unexpected changes in U.S. inflation have large effects on inflation in other countries, although they are not always the dominant international factor. Similarly, shocks to some other countries also have a statistically and economically significant influence on U.S. inflation. Moreover, our evidence indicates that U.S. inflation has become less vulnerable to foreign shocks since the early 1990s, mainly because of the diminished influence from Germany and France.

Keywords: directed acyclic graphs; forecast error variance decomposition; recursive estimation

JEL Classifications: G15; C32

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1. Introduction

Most monetary authorities around the world would agree that maintaining price stability should be their main objective.[1] For example, an increasing number of central banks, including those of the U.K. and Canada, have adopted explicit inflation targeting over the past 15 years (e.g., Johnson, 2002). Others, the U.S., Germany, and Japan, for example, have also acted aggressively to contain inflation since the late 1970s, although they do not have an explicit inflation target (e.g., Clarida et al., 1998). The collective efforts among central banks for fighting inflation have coincided with a noticeable decline of inflation rates in most industrial countries (e.g., Levin and Piger, 2003). This casual observation is consistent with the notion that the international transmission is a significant part of the dynamic of inflation under both the fixed- and flexible-exchange-rate regimes (e.g., Darby and Lothian, 1983, 1989).[2]

Indeed, as the world economy is growing more and more integrated in nature, the transmission of inflation across countries has become an increasingly important concern for the conduct of optimal monetary policies. This issue is especially relevant also because most countries are small open economies, which are vulnerable to external influence. Therefore, a good understanding of the international transmission of inflation provides valuable guidance for central banks in coordinating their monetary policies to maintain price stability. Surprisingly, there are few empirical studies on this critical issue after the publication of classic works by

[1] Woodford (2004) and others argue that such a policy is also (approximately) optimal in standard monetary models.

[2] Early authors (e.g., Friedman, 1953) argue that the current floating exchange rate provides complete insulation, and thus a country's inflation is determined solely by its own monetary policies. However, this conjecture is likely to be unrealistic because of central bank interventions in the foreign exchange markets and the lack of pass-through in import goods prices (e.g., Devereux and Engel, 2002). Also see Svensson (2000), Clarida et al. (2002), Benigno and Benigno (2003), and others for recent theoretical analysis of monetary policies in open economies.

Darby and Lothian (1983, 1989) and others; in this paper, we try to fill this gap by providing some preliminary results using more recent data and more sophisticated statistical techniques.

We analyze the CPI (consumer price index) inflation transmission among G-7 countries over the period 1973 to 2003, using a vector autoregression (VAR) specification. In particular, we apply the directed acyclic graphs (DAG) technique (Pearl, 2000; Sprites et al., 2000) to determine the contemporaneous causal flows, which are then used to conduct a data-determined structural decomposition of the VAR shocks. The advantage of this approach is that, as advocated by Swanson and Granger (1997), it allows for the properties exhibited in the data and thus is less arbitrary than the recursive causal structure embedded in the commonly used Cholesky decomposition. This difference is found to be important in this paper.

Our main results can be summarized as follows. First, in *contemporaneous* time, U.S. inflation is substantially affected by unexpected changes in inflation (UCII, thereafter) originating from Canada, Germany, and Italy, which jointly account for over 12 percent of total variations in the U.S. In contrast, foreign UCII explain no more than 3 percent for the other G-7 countries, except the U.K. (8 percent) and France (19 percent). Also, U.S. UCII contemporaneously affect only French inflation. These results indicate that it is inappropriate to assume that U.S. UCII are the most important factor in the Cholesky decomposition.

Second, variance decompositions show that foreign influence on U.S. inflation increases moderately with forecast horizons, from 12 percent in contemporaneous time to 19 percent at the 24-month horizon. This result suggests that foreign UCII are transmitted into U.S. inflation very quickly but that their effects are mainly transitory. For the other G-7 countries, however, foreign influence becomes much more important as forecast horizons increase: It accounts for 48 to 74 percent of price variations at the 24-month horizon. Thus U.S. inflation is actually the least vulnerable to external shocks in the long run.

Third, despite their small contemporaneous effects, U.S. UCII explain a large portion of the long-run (24-month horizon) price variations of the other G-7 countries, with an average of 30 percent. We also document significant transmission among the other G-7 countries. In particular, the U.S. exerts less influence on Japanese and German inflation than Canada and the U.K., respectively, do. Therefore, there is a broad linkage of inflation among G-7 countries.

Lastly, the VAR system is found to be potentially unstable because of the Lucas (1976) Critique: Inflation is affected by monetary policies, which have changed over time. In particular, our recursively estimated forecast error variance decompositions show that U.S. inflation became less vulnerable to external shocks in the recent period mainly because of the diminishing influence from Germany and France. The effect of U.S. UCII on the other G-7 countries, however, appears to be relatively stable over time.

Crowder (1996) documents a strong convergence of inflation among G-7 countries. Cheung and Yuen (2002) investigate the interaction between the U.S. and two small open economies (Hong Kong and Singapore). Our paper is most closely related to Eun and Jeong (1999), who use the Cholesky decomposition to analyze inflation transmission among G-7 countries. However, as mentioned above, the limitations of the Cholesky decomposition make their results untenable.

The remainder of the paper is organized as follows. Section 2 discusses the empirical framework and section 3 describes the data and presents empirical results. We offer some concluding remarks in section 4.

2. Empirical Framework

2.1. Error Correction Models and Innovation Accounting

We assume that CPIs (consumer price indices) of G-7 countries follow an integrated process of order one, and we will discuss the unit root test in the next section. X_t denotes a vector of nonstationary CPIs, which can be modeled in an error correction model (ECM):

$$(1) \qquad \Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + e_t \, (t = 1, \ldots, T) \, .$$

Equation (1) resembles a VAR model in first differences, except for the presence of a (lagged) level, X_{t-1}. The parameters in matrix Π contain information about the long-run cointegration relationship among seven price indices.

If CPIs are cointegrated among G-7 countries, as shown in the next section, we can estimate an ECM with appropriate lags. Because the individual coefficients, especially those related to the short-run dynamics, Γ, do not have a straightforward interpretation, we use the Sims (1980) innovation accounting method to illustrate the short-run dynamic structure.[3] To illustrate this, we rewrite ΔX_t of equation (1) as an infinite moving average process:

$$(2) \qquad \Delta X_t = \sum_{i=0}^{\infty} A_i \varepsilon_{t-i} \, , \quad t = 1,2, \ldots, T.$$

The error from the forecast of ΔX_t at the *n*-step-ahead horizon, conditional on information available at *t*–1, Ω_{t-1}, is

[3] To impose cointegration constraints, we actually invert the estimated ECM to derive the level VAR representation and then use innovation accounting based on the equivalent level VAR to summarize the short-run dynamic interactions among G-7 countries.

(3) $$\xi_{t,n} = \sum_{l=0}^{n} A_l \varepsilon_{t+n-l} \ .$$

Therefore, the variance-covariance matrix of the total forecasting error is

(4) $$Cov(\xi_{t,n}) = \sum_{l=0}^{n} A_l \Sigma A_l^{'} \ ,$$

where Σ is the variance-covariance matrix of the error term, e_t, as in equation (1). The remaining basic problem is how to orthogonalize the ECM residuals. The early research usually adopts the Cholesky factorization to achieve a just-identified system in contemporaneous time. This assumption leads to the following variance decomposition for the forecasting error:

(5) $$\theta_{ij}^{c}(n) = \frac{\sum_{l=0}^{n} (e_i^{'} A_l \Sigma P e_j)^2}{\sum_{l=0}^{n} (e_i^{'} A_l \Sigma A_l^{'} e_i)}, \quad i, j = 1, 2, ..., 7 \ ,$$

where P is the Cholesky factor of the residual variance-covariance matrix Σ, and e_i is a selection vector, with the *i*th cohort equal to 1 and all the other cohorts equal to 0. Therefore, $\theta_{ij}^{c}(n)$ measures the contribution of the *j*th-orthogonalized innovation to the variance of the total *n*-step-ahead forecasting error for variable ΔX_{it}.

In the Cholesky decomposition, we assume that there exists a recursive contemporaneous causal structure. This assumption, however, is restrictive and often unrealistic (e.g., Swanson and Granger, 1997). More fundamentally, economic theories rarely provide guidance for contemporaneous causal orderings, and VAR practitioners usually need to rely on various stories to determine them arbitrarily. Hence, it would be (more or less) ironic that the VAR method that originated as a way of getting away from incredible identifying restrictions on large scale macroeconomic models has to rely heavily on hardly more-credible stories to identify contemporaneous causal orderings (Demiralp and Hoover, 2003, p. 747). However, as advocated by Swanson and Granger (1997), the DAG can be used to uncover contemporaneous causal orderings in a data-determined and, thus, less ad hoc manner, which we discuss next.

2.2. Directed Acyclic Graphs Analysis

The DAG technique, which represents the recent advance in causality analysis, has received an increasing amount of attention in the empirical literature. In this subsection, we briefly describe how we conduct the DAG analysis using the variance-covariance matrix of the ECM residuals in equation (1). Also see Hoover (2003), Granger (2003), Demiralp and Hoover (2003), and Bessler and Yang (2003), among others, for detailed discussion on the DAG analysis.

A directed graph is essentially an assignment of the contemporaneous causal flow (or lack thereof) among a set of variables or vertices based on observed correlations and partial correlations. The edge relation characterizing each pair of variables represents the causal relation (or lack thereof) between them. In the context of the DAG used in this study, there are five possible edge relationships: (1) No edge (X   Y) indicates (conditional) independence between two variables. (2) Undirected edge (X — Y) signifies a covariance that is given no causal interpretation. (3) Directed edge ($Y \rightarrow X$) suggests that a variation in Y, with all other variables held constant, produces a (linear) variation in X that is not mediated by any other variable in the system. (4) Directed edge ($X \rightarrow Y$) has an analogous interpretation as (3). (5) Bidirectional edges ($X \leftrightarrow Y$) denote the bidirectional causal interpretation between the X and Y.

The basic idea of DAG (Pearl, 2000; Spirtes et al., 2000) builds on the insight of a non-time sequence asymmetry in causal relations, whereas the well-known Granger causality exploits the time sequence asymmetry that a cause precedes its associated effect (and thus an effect does not precede its cause). To illustrate, consider a causally sufficient set of three variables X, Y, and Z. A causal fork that X causes Y and Z can be illustrated as Y ← X → Z. Here the unconditional association between Y and Z is nonzero (as both Y and Z have a common cause in X), but the conditional association between Y and Z, given knowledge of the common cause X, is zero: C*ommon causes screen-off associations between their joint effects.* Now consider the so-called inverted causal fork, that X and Z cause Y, as X→ Y ← Z.[4] Here the unconditional association between X and Z is zero, but the conditional association between X and Z, given the common effect Y, is not zero: *Common effects do not screen-off association between their joint causes*. See Demiralp and Hoover (2003) for a lucid discussion on this point.

Assuming that the information set, Ω_{t-1}, is causally sufficient, Spirtes et al. (2000) provide a directed graph algorithm (i.e., PC algorithm) for removing edges between variables and directing causal flows of information between variables. The PC algorithm begins with an undirected graph, in which shocks to each variable are connected with shocks to all the other variables. It then proceeds in two stages: elimination and orientation. In the elimination stage, the algorithm removes edges from the undirected graph, based on unconditional correlations between pairs of variables: Edges are removed if they connect variables that have zero correlation. The remaining edges are then checked for whether the first-order partial correlation (correlation between two variables conditional on a third variable) is equal to zero. If it is zero, we remove the edges connecting the two variables. The remaining edges are then checked against zero second-order conditional correlation and so on. For N variables, the algorithm continues to check up to (N –

[4] As pointed out by a referee, so-called inverted causal forks are frequently known as "colliders." The discussion here is valid only for "*unshielded colliders*," in which X and Y are not directly connected. It does not apply, however, to "*shielded colliders*," in which X and Y are directly connected.

2)th-order conditional correlation.

In applications, Fisher's z statistic is used to test whether conditional correlations are significantly different from zero. To test whether conditional correlations are significantly different from zero, we use Fisher's z statistic, $z(\rho[i,j|k]n) = 1/2(n - |k| - 3)^{1/2} \times \ln\{(|1 + [i,j|k]|) \times (|1 - [i,j|k]|)^{-1}\}$. In this statistic, n is the number of observations used to estimate the correlations; $\rho(i,j|k)$ is the population correlation between variables i and j conditional on variables k (i.e., removing the influence of variables k from variables i and j); and $|k|$ is the number of variables in k. If variables i, j, and k are normally distributed and $r(i,j|k)$ is the sample conditional correlation of i and j given k, $z(\rho[i,j|k]n) - z(r[i,j|k]n)$ has a standard normal distribution.

Once the elimination stage is completed, the algorithm proceeds to the orientation stage. The notion of *sepset* is then used to assign the direction of contemporaneous causal flow between variables remaining connected after we check for all possible conditional correlations.[5] The sepset of a pair of variables whose edge has been removed is the conditioning variable(s) on the removed edge between two variables. For vanishing zero-order conditioning (unconditional correlation), the sepset is an empty set. Edges remaining connected are directed by considering triples X — Y — Z, in which the pair X and Y and the pair Y and Z are adjacent but X and Z are not. Edges are directed between triples X — Y — Z as X $\rightarrow$ Y $\leftarrow$ Z if Y is not in the sepset of X and Z. If (1) X $\rightarrow$Y, (2) Y and Z are adjacent, (3) X and Z are not adjacent, and (4) there is no arrowhead at Y, then Y — Z should be positioned as Y $\rightarrow$ Z. If there is a directed path from X to Y and an edge between X and Y, then X — Y should be positioned as X $\rightarrow$Y.

[5] See Yang and Bessler (2004) for more illustrations on the notion and use of the sepset.

The PC algorithm discussed above is commonly used and implemented in the program TETRAD III (Scheines et al., 1996), which we also use for the empirical analysis in this paper. To be robust, we also test the contemporaneous causal pattern identified by the DAG using the likelihood ratio test by Sims (1986). This is equivalent to testing certain combinations of zero restrictions on a_{ij}, which result in an overidentified *A* matrix. The likelihood ratio test on the parameter restrictions relating observed shocks (e_t) to orthogonal shocks (v_t) can be derived from the equation $Ae_t = v_t$. Specifically, the test statistic is given as $T[\log(\det(\Omega)) - \log(\det(\Sigma))]$, where Ω is the variance-covariance matrix derived from the *A*-matrix restrictions, Σ is the variance-covariance matrix derived from the observed nonorthogonal shocks, *T* is the number of observations used to estimate the model, *log* is the logarithmic transformation and *det* is the determinant operator. The test statistic has a chi-squared distribution with $(\frac{n(n\text{-}1)}{2} - m)$ degrees of freedom, where *n* is the number of series in the VAR and *m* is the number of overidentifying restrictions.

3. Empirical Results

3.1. Unit Root and Cointegration Tests

The CPI data are obtained from International Financial Statistics (IFS) for G-7 countries. We focus on the post-Bretton Woods period July 1973 to June 2003 because, as shown by Crowder (1996), inflation of the other G-7 countries moved closely with U.S. inflation before the collapse of the Bretton Woods system. Consistent with Eun and Jeong (1999), we cannot reject at the 5 percent significance level the null of a unit root for any country except Japan.

To estimate equation (1), we first select the optimal number of lags by minimizing the Akaike information criterion (AIC), with the maximum number set to 12. The AIC suggests an optimal number of four lags, or k = 4, for the level VAR and we thus use three lags, or k = 3, for the ECM in equation (1). We then conduct the Johansen (1991) trace test for the cointegration among seven CPIs and report the results in Table 1. We fail to reject at the 5 percent significance level that there are four cointegrating vectors, either with a constant included in the cointegrating space or with a linear trend. To be robust, we also have conducted recursive estimation of cointegration rank and found that the cointegration rank of 4 is rather stable over time. The estimated correlation matrix of the ECM shocks is

$$(6) \qquad V = \begin{bmatrix} 1 & & & & & & \\ .047 & 1 & & & & & \\ .151 & .060 & 1 & & & & \\ .278 & .198 & .128 & 1 & & & \\ .169 & .166 & .073 & .139 & 1 & & \\ .004 & .280 & .091 & .276 & -.027 & 1 & \\ .248 & -.151 & -.153 & .180 & .010 & -.115 & 1 \end{bmatrix}$$

In equation (6), we report only the lower triangular entries in the following order: ΔX_1, ΔX_2, ΔX_3, ΔX_4, ΔX_5, ΔX_6, ΔX_7, where the subscripts 1 through 7 denote the U.S., Japan, Germany, France, Italy, the U.K., and Canada, respectively. This matrix is the major input for the DAG analysis, which we discuss next.

3.2. Contemporaneous Causal Flows

Figure 1 plots the final directed graph of the residuals obtained from our seven-country ECM model at the 5 percent significance level. The figure is obtained using the PC algorithm, with the assumption of causal sufficiency, as programmed in Tetrad III (Scheines et al., 1996).

The 5 percent significance level is chosen based on the sample size and simulation evidence in Scheines et al. (1996); however, we obtain very similar results using the 1 percent level.

Interestingly, and perhaps somewhat surprisingly, U.S. inflation is among the most influenced by other countries in contemporaneous time, with edges running to it from Canada, Germany, and Italy. Similarly, U.S. inflation contemporaneously affects only French inflation. Our results thus cast doubt on the assumption of U.S. inflation as the most important factor in the Cholesky decomposition, as in Eun and Jeong (1999), for example.

Figure 1 also reveals strong contemporaneous links among G-7 countries. In addition to four edges from and to the U.S., there are three edges from Canada, Japan, and the U.K. to France. Also, there are two edges from Japan to Italy and the U.K., one edge from Germany to Canada, and one edge from Canada to Japan. We find no undirected edges and thus no additional contemporaneous causal flows between any pair of these countries. Overall, our results indicate that inflation is transmitted contemporaneously among G-7 countries.

We derive Figure 1 using the assumption of causal sufficiency, e.g., that we use a sufficiently rich set of theoretically relevant variables; however, failure to include a relevant variable may lead one to put an edge between two variables when in fact both are effects of an omitted third variable. The causal sufficiency assumption is unlikely to be completely satisfied by our parsimonious VAR specification in equation (1). For example, Stock and Watson (1999) show that economic fundamentals suggested by the Philips curve help forecast inflation. Also, inflation in G-7 countries may be affected by UCII in other countries. But adding these additional variables requires more identification assumptions and thus may obscure the dynamic interaction of inflation across countries, which is the main focus of this paper. More importantly, omitting these variables is likely to have a small effect because inflation is explained mostly by its own lags. Similarly, given that G-7 countries account for a dominant portion (e.g., 67 percent in 2001) of world output, their inflation is likely to be affected mainly by their own economic fundamentals. To summarize, while the omitted variable problem is a potential issue, it is unlikely to influence our main results in any qualitatively manner, as we discuss next.

To be robust, we investigate potential effects of omitted variables, using the FCI algorithm without the causal sufficiency assumption. The identified causal pattern remains the same for all the countries, except that there are bidirectional edges between Germany and Canada and between France and Canada. The latter result indicates that some latent variables may account for the causal flow pattern between these two pairs, results of which should thus be interpreted with caution. However, given that the other eight pairs in Figure 1 are unaffected, the omitted variable problem is unlikely to affect our main results, particularly the interaction between the U.S. and the other G-7 countries, in any qualitative manner. With these caveats in mind, we report the empirical results based on the DAG (in Figure 1) below.

3.3. Forecast Error Variance Decompositions

To illustrate the economic significance and the dynamic pattern of the international transmission of inflation, we present in Table 2 the forecast error variance decomposition—the percentage of price variations in each country at time $t+k$ that are due to UCII in all the countries at time t. The decomposition is based on the contemporaneous causal pattern derived from the DAG (Figure 1), which is not rejected by the Sims' (1986) likelihood ratio test at the 5 percent significance level. Table 2 reports the decomposition at the 1-month (contemporaneous time), the 3- and 6-month (short), and the 12- and 24-month (long) horizons.

Consistent with Figure 1, U.S. inflation is among the most influenced by other countries in contemporaneous time. For example, foreign inflation accounts for over 12 percent of U.S. price variations at the 1-month horizon, compared with 2 percent for Japan, 0 percent for Germany, 3 percent for Italy, 8 percent for the U.K., and 2 percent for Canada. France is the only country that is more vulnerable than the U.S. to foreign shocks in contemporaneous time.

The picture is quite different at the longer horizon, however. In particular, unlike in all the other countries, the effect of foreign shocks on U.S. inflation increases only moderately over forecast horizons. As a result, U.S. inflation is the least influenced by other countries at the 24-month horizon: Foreign UCII account for 19 percent of total U.S. price variations, compared with 48 to 74 percent for the other G-7 countries. The U.S. also plays a very important role in the international transmission of inflation at the longer horizon, although it contemporaneously affects few other countries. For example, at the 24-month horizon, U.S. UCII on average account for 30 percent of price variations in the other countries.

It is interesting to note that U.S. UCII are not always the dominant international factor for Japanese and German inflation. Similarly, Japanese and German UCII explain a negligible amount of variations of U.S. inflation. As mentioned in the introduction, Clarida et al. (1998) show that monetary authorities in the U.S., Japan, and Germany acted aggressively to contain inflation during most of the period under study in this paper. Therefore, our results seem to suggest that these central banks have tried to neutralize the inflationary pressure from each other. Consistent with this interpretation, we show in the next subsection that U.S. inflation actually reacted negatively to Japanese and German UCII.

Like many other G-7 countries, Japan is barely influenced by other countries at the 1-month (3 percent) and 3-month (12 percent) horizons. However, it becomes highly vulnerable to foreign influence at the longer horizon. For example, foreign UCII explain over 50 percent of its price variations at the 24-month horizon. Canada—a major exporter of raw materials to Japan—accounts for about 16 percent of Japanese price variations at the 24-month horizon. Similarly, the U.S.—a major trade partner—accounts for an additional 12 percent. Italy, Germany, and the U.K. also explain a substantial portion of the variations in Japanese prices. In contrast, Japanese UCII apparently explain few price variations in the other countries.

Among G-7 countries, German inflation is the least affected by foreign UCII in contemporaneous time. Germany remains barely influenced by other countries up to the 12-month horizon (18 percent). This result might be consistent with the well-known German dominance hypothesis that the German monetary authority has been traditionally very independent in setting its monetary goals (e.g., Uctum, 1999). However, at the 24-month horizon, the U.S. (14 percent), the U.K. (15 percent), and Italy (9 percent) can exert substantial influence on German inflation. It is interesting to note that the U.K. and Italy are the only two countries that dropped out of the EMS (European Monetary System) because of difficulties in keeping up with German monetary policy. Thus, the result based on the whole sample here may largely reflect the interaction between Germany and the other European countries since 1990, which we will discuss further below.

French inflation is the most vulnerable to the influence of foreign inflation at the contemporaneous horizon as well as the 3-month horizon. U.S. UCII are the main international source of its price variations at the horizon of 6 months or longer.

Italy is also barely influenced by other countries at the contemporaneous horizon (3 percent) and at short horizons (11 to 19 percent). However, Italian inflation becomes the most vulnerable to the influence of foreign inflation at the 24-month horizon—only 26 percent of the price variations are explained by its own shocks. The U.S. and France explain about 45 percent and 24 percent, respectively, of the price variations in Italy. Interestingly, Italian UCII also account for a substantial portion of price variations in most other countries. For example, it is 11 percent for the U.S., 10 percent for Japan, 9 percent for Germany, 14 percent for the U.K., and 5 percent for Canada at the 24-month horizon. This result might reflect the fact that Italy has the worst record on price stability among G-7 countries.

The U.K. is also highly independent at the contemporaneous and short horizons, but it is much less so at the longer horizons. The U.S. and Italy, respectively, account for 33 percent and 14 percent of price variations in the U.K. at the 24-month horizon, while the other countries exert little influence.

Lastly, Canada is similar to Italy in that Canada is among the most influenced by the other countries at the 24-month horizon and among the least influenced at the contemporaneous and short horizons. At the 24-month horizon, Canada is vulnerable to shocks originating from the U.S. (36 percent), the U.K. (23 percent), and Germany (12 percent).

3.4. Impulse Responses

We also use the ordering of shocks derived from the DAG analysis to generate the impulse responses, which illustrate how domestic inflation reacts to UCII in foreign countries. To conserve space, we plot only the responses of U.S. inflation to shocks (defined as a 1 percent unexpected price increase) from all the countries (Figure 2) and the responses of all the other countries to shocks from the U.S. (Figure 3). Following Christiano, Eichenbaum, and Evans

(1996), among others, one-standard-deviation bands (dashed lines) are also plotted to show whether the point estimates of the impulse response are statistically significant.[6]

Figure 2 shows that, at most horizons (including the 24-month horizon), U.S. inflation responds *significantly* to shocks from Japan, Italy, and the U.K., in addition to its own shocks. Therefore, inflation does transmit from the other countries to the U.S. economy. We also observe an interesting asymmetry in U.S. responses to foreign shocks. In particular, U.S. inflation reacts *negatively* to UCII in Canada, Japan, and Germany at long horizons, although it has a positive response to shocks from the U.K., France, and Italy. The response is also statistically significant at the 24-month horizon for all countries except France and Canada. Given that the monetary policies in the U.K., France, and Italy are greatly influenced by the Bundesbank (Clarida et al., 1998), the U.S. Fed might not want to react directly to the monetary shocks in these countries. In contrast, the central banks of Canada, Japan, and Germany focus mainly on their domestic objectives. Therefore, the negative responses to their UCII indicate that the U.S. Fed tried to neutralize the inflationary pressure originating from these countries.[7] As mentioned above, these

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[6] Recently, Jorda (2004) developed a method to conduct impulse response analysis based on local projections. However, as we focus on the data-determined identification of contemporaneous causal relationships and its impact on impulse response analysis, we leave application of this potentially important technique for future research.

[7] Eun and Jeong (1999) find that, at the 24-month horizon, U.S. inflation reacts negatively to the innovations from all the other G-7 countries except France. They explain that negative responses might happen in the U.S. if an inflationary foreign shock is accompanied by an "overshooting" depreciation of the foreign currency against the U.S. dollar, actually lowering the dollar prices of foreign imports. However, their explanation does not explain the positive response to shocks in France, whose economy shares many similar features to those of the U.K. and Italy. The difference between Eun and Jeong (1999) and our paper reflects the different identification assumptions used in the two papers.

interpretations are consistent with the result that UCII in Canada, Japan, and German have a negligible effect on U.S. inflation at the 24-month horizon (Table 2).

Figure 3 shows that all the other G-7 countries respond positively and significantly to shocks from the U.S. Interestingly, we observe that the response decreases around the 4-month horizon for Canada, Japan, and Germany, indicating that these countries also attempt to neutralize inflationary pressure from the U.S. In contrast, we do not observe such a pattern for France, Italy, and the U.K.

There are three main channels, namely, the monetary channel, the income channel, and the price channel, through which inflation is transmitted across countries (e.g., Darby and Lothian, 1983). While a formal investigation of the relative importance of these three channels is important for understanding the international transmission of inflation, it is beyond the scope of our paper and we leave it for future research. Here, we draw some casual observations based on the empirical analysis above. First, the price channel generates a direct effect, through which inflation might transmit more quickly than through the other two channels. The contemporaneous transmission of inflation across G-7 countries thus provides evidence for an important role of the price channel. Second, foreign influence increases substantially over time, indicating that the monetary and income channels are also important. Third, as shown in Figures 2 and 3, central banks tend to neutralize external influence, revealing a special role of the monetary channel. In summary, consistent with Darby and Lothian (1983, 1989), inflation transmission is a slow process and all three channels, particularly the monetary channel, appear to be important.

Finally, we have conducted a number of robustness tests on the results reported above, which are available upon request. In particular, we experimented with different lags (three to six lags) in the ECM and found qualitatively the same results. We also estimated a model assuming stationary Japanese CPI and nonstationary CPIs for all the other countries. Again, the basic inference remains qualitatively the same. Therefore, the results presented above are quite robust.

3.5. Recursive Forecast Error Variance Decompositions

The Lucas Critique suggests that the dynamic of inflation is potentially unstable because it depends on monetary policies, which have changed over time. To illustrate this point, we first present recursively estimated forecast error variance decompositions: Substantial changes across time are an indication of structural breaks. However, the recursive results should be interpreted with caution if there is a structural break: Parameter estimates after the break might be imprecise or even invalid because they use data from two different regimes. To formally address this issue, we also use Bai and Perron's (1998, 2003) structural break tests to explicitly date potential breaks and then conduct the subperiod analysis in the next subsection.

Figure 4 plots the recursive variance decompositions of U.S. inflation at the 24-month horizon. In particular, we use the sample from July 1973 to December 1989 to conduct the first decomposition and the sample from July 1973 to January 1990 for the second decomposition, and so on. It shows that the influence of Germany and France on the U.S. has dramatically diminished since the early 1990s, while the influence of Italy has risen moderately. The influence of the other countries, however, has remained relatively weak except for a spike in the influence of the U.K. around 1992. Overall, the portion of variations in U.S. inflation explained by foreign shocks decreased from 40 percent in the latter 1980s to 20 percent in the early 2000s.

Several historic episodes might explain these changes in the international transmission of inflation. The U.S. Fed has become more aggressive in containing inflation in the Volcker-

Greenspan period than in the pre-Volcker period (Clarida et al., 2000). Our results thus might indicate that the Fed successfully insulated the inflationary pressure from other countries, especially Germany and France, which were the main sources of international influence on U.S. inflation in the late 1980s (Figure 4). Also, our results might reflect the vanished German dominance after its unification, as documented by Uctum (1999).[8] Over a long period, the Bundesbank had been effectively running monetary policies for France and Italy; however, it lost dominance after the Bundesbank was forced to raise the short-term interest rate to curb the rising inflation caused by the unification. Moreover, the monetary tightening eventually led to the EMS collapse—another potential structural break—because interest rates in the other European countries had become much higher than domestic macroeconomic conditions warranted.

Figure 5 plots the recursive variance decompositions of the other countries to U.S. shocks at the 24-month horizon. The responses of Japan, Italy, and France are quite stable over time, except that Japan shows some increases in the mid-1990s. U.S. shocks appear to have an increasingly important effect on the U.K. and Canada—the only two G-7 countries that explicitly adopted inflation targeting (in the early 1990s). Their effects on German inflation, however, seem to have become weaker. Nevertheless, the changes are much smaller than those in Figure 4.

3.6. Structural Breaks

In this subsection, we use Bai and Perron's (1998, 2003) methodology to formally analyze structural breaks in our data. We apply the test separately for each equation in the VAR system because the test is designed for the univariate regression. We do not include all 29 right-

[8] We find a similar change for France possibly because the Bank of France has followed the moves of the Bundesbank most closely among the other European countries (e.g., Clarida et al., 1998).

hand-side variables for each equation in the test because Bai and Perron (2003) provide critical values for testing with up to only 10 variables subject to structural changes. Since the dynamic effects appear to be mainly captured by first lags, we consider partial structural change models, including a constant and seven first-lagged terms. We allow for a maximum of five structural breaks and use the 5 percent significance level in making inferences. In all cases, we reject the null hypothesis of no breaks (either against a known or against an unknown number of breaks).

The sequential F-tests indicate three breaks in five CPIs and two breaks in the other two CPIs. The first break was identified as occurring around the period October 1981 to August 1982 for most countries except the U.K. (July 1979) and Germany (August 1988). This result confirms that the shift of U.S. monetary policies in the Volcker period has a significant effect on inflation. Dating is less consistent across countries for the second and third breaks, although a few of these breaks occurred around the German unification and the EMS collapse.

To investigate effects of the first break on our results, we exclude the above break points or periods and analyze two subsamples: July 1973 to September 1980 and September 1982 to June 2003. The final directed graphs based on the two subsamples are somewhat different from the one based on the full sample, in terms of both edge inclusion/exclusion and edge direction. Nevertheless, the associated variance decompositions are qualitatively similar to those obtained from the recursive estimation (Figures 4 and 5). In particular, U.S. inflation is less affected by external shocks after the break mainly because of the diminishing impact from France. For brevity, these results are not reported here but are available upon request.

The German unification (in July 1990) might also have led to structural changes in the international transmission of inflation but evidence is somewhat weaker: Only Canada, France, and the U.K. had breaks around this period. We also consider two subsample periods to investigate the effect of the German unification: July 1973 to March 1990 and February 1991 to June 2003. Results obtained from the latter subsample are similar to those obtained from the post-Volcker subsample, as discussed above. However, the PC algorithm does not direct the remaining edges for the former subsample, even if we exclude the observations prior to the first break. Germany also had a break around the 1992 EMS collapse. However, it is difficult to distinguish these effects from those of the German unification because these two breaks are only two years apart.

To summarize, consistent with the recursive estimation, we find significant breaks in the international transmission of inflation. Moreover, our subsample analysis confirms that U.S. inflation has become less vulnerable to external influence in the recent period.

4. Conclusion

We uncover a broad linkage of inflation among G-7 countries during the post-Bretton Woods period 1973 to 2003. Inflation is transmitted not only from the large countries, such as the U.S., to other smaller open economies, but also the other way around. This result indicates that it is difficult for a central bank to act alone in combating inflation. Therefore, the improved price stability in the past two decades among most industrial countries could reflect the collective efforts of their central banks for containing inflation.

We also document significant breaks in the international transmission of inflation. The first break occurred around the early 1980s, coinciding with a shift of the U.S. Fed to more restrictive monetary policies on inflation. The second break happened around the 1990 German unification and possibly also the 1992 EMS collapse. Consequently, U.S. inflation has become less vulnerable to foreign shocks, especially those originating from Germany and France. These results indicate that, as argued by many authors (e.g., Woodford, 2004), a long-run commitment to price stability is the key to winning the battle against inflation. Overall, our study suggests that, to effectively thwart global inflation, monetary authorities in G-7 countries might want to pursue more concerted policy coordination, including a joint policy goal of a long-term commitment to price stability.

It is important to note that some other structural changes might have potentially important effects on global inflation. In particular, an oil price hike always raises the concern that it might lead to higher inflation rates. In hindsight, we could say that there have been three oil price shock periods: 1973-74, 1979-80, and the recent one we are experiencing now. Also, it is arguable that China, being a provider of less expensive goods to G-7 countries, contributes to the import countries' lower inflation. Moreover, as China has become a major economic force, it might have to be included in the G-7 forum. These important issues are beyond the scope of this paper and we leave them for future research.

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Table 1. Johansen Trace Test

	Without a Linear Trend		With a Linear Trend	
H_0:	T	C(5%)	T	C(5%)
r = 0	323.79	132.00	298.46	123.04
r = 1	185.65	102.14	161.30	94.15
r = 2	100.10	76.07	89.29	68.52
r = 3	59.57	53.12	52.30	47.21
r = 4	31.16	34.91	25.374	29.68
r = 5	12.82	19.96	8.49	15.41
r = 6	3.20	9.24	0.71	3.76

Notes: Trace test statistics (T) are compared with the critical values (C). The lags in the underlying VARs are determined by the Akaike information criterion.

Table 2. Forecast Error Variance Decompositions

Month	US	Japan	Germany	France	Italy	UK	Canada
	Variance of US explained by price shocks to the seven countries						
1	87.86	0.06	1.97	0.00	2.28	0.00	7.15
3	86.25	0.24	4.45	0.29	1.63	0.73	6.08
6	82.40	1.53	4.90	0.54	5.50	1.25	4.34
12	83.38	2.43	1.95	0.57	10.04	1.94	1.66
24	80.99	3.44	1.95	0.47	10.54	5.41	0.93
	Variance of Japan explained by price shocks to the seven countries						
1	0.00	97.72	0.05	0.00	0.00	0.00	2.23
3	1.07	88.28	0.11	0.11	2.52	0.19	7.49
6	2.26	71.62	11.73	0.32	3.53	1.08	7.41
12	3.75	63.59	9.52	0.79	6.12	1.02	12.80
24	11.92	47.09	6.52	1.44	9.54	6.59	15.62
	Variance of Germany explained by price shocks to the Seven Countries						
1	0.00	0.00	100.00	0.00	0.00	0.00	0.00
3	0.13	0.03	97.94	0.09	0.60	0.01	0.83
6	0.15	0.04	93.17	0.06	2.04	2.33	2.69
12	2.12	0.03	81.98	0.43	3.80	11.07	3.07
24	14.09	0.28	59.76	1.81	9.07	14.73	3.85
	Variance of France explained by price shocks to the seven countries						
1	4.54	4.70	0.01	80.65	0.12	6.03	4.06
3	11.29	2.82	0.06	77.88	0.32	3.32	4.39
6	10.73	1.19	0.68	81.22	0.69	1.34	2.84
12	17.76	0.50	0.32	80.29	0.34	0.68	1.26
24	37.37	0.30	1.44	62.36	0.10	2.44	0.39
	Variance of Italy explained by price shocks to the seven countries						
1	0.00	2.71	0.00	0.00	97.23	0.00	0.06
3	1.19	3.81	0.13	0.63	89.26	0.75	2.90
6	3.50	1.71	0.62	3.08	81.82	2.20	4.25
12	15.84	0.86	0.45	12.09	62.07	1.33	4.60
24	44.61	0.51	0.25	23.63	26.35	2.04	2.81
	Variance of UK explained by price shocks to the seven countries						
1	0.00	7.68	0.00	0.00	0.00	92.14	0.18
3	1.99	6.29	0.21	0.59	0.52	91.33	0.12
6	5.14	4.03	0.67	0.80	5.07	85.83	0.09
12	14.19	2.79	0.81	1.58	12.07	72.23	0.11
24	32.64	1.20	4.50	2.11	14.48	50.78	0.06

Table 2. Forecast Error Variance Decompositions (Continued)

	Variance of Canada explained by price shocks to the seven countries						
1	0.00	0.00	2.33	0.00	0.00	0.00	97.67
3	1.20	0.71	1.08	0.03	0.25	1.53	96.24
6	1.47	0.42	0.84	0.50	0.99	6.24	91.91
12	9.93	0.29	4.58	1.49	2.88	14.00	71.82
24	36.02	1.18	11.87	2.22	4.51	23.29	28.60

Notes: The forecast error variance decomposition is conducted based on the directed graph on UCII, as shown in Figure 1. Month 1 is the contemporaneous month. Each panel shows how variance of inflation in a country is explained in percentage points by price shocks to the seven countries listed in the first row.

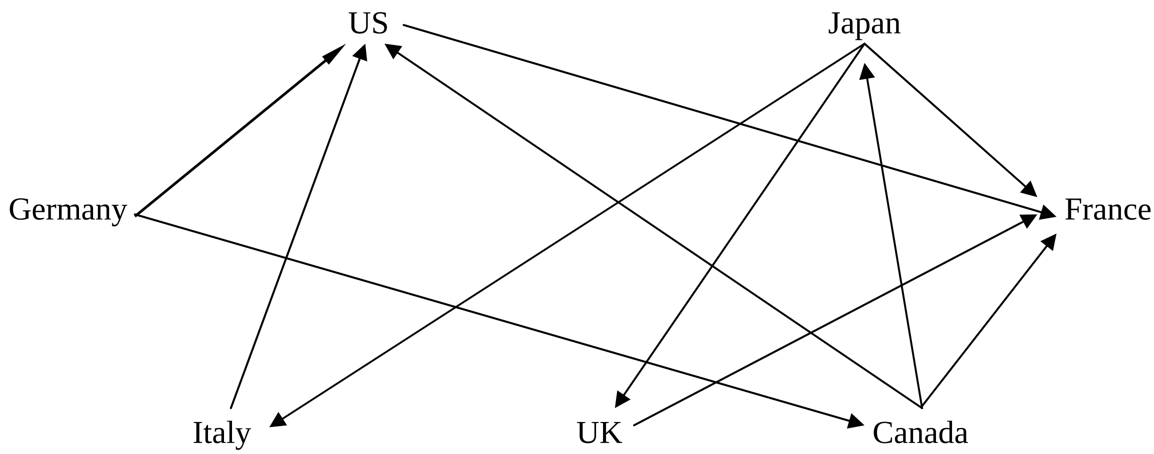

Figure 1. Directed Acyclic Graph

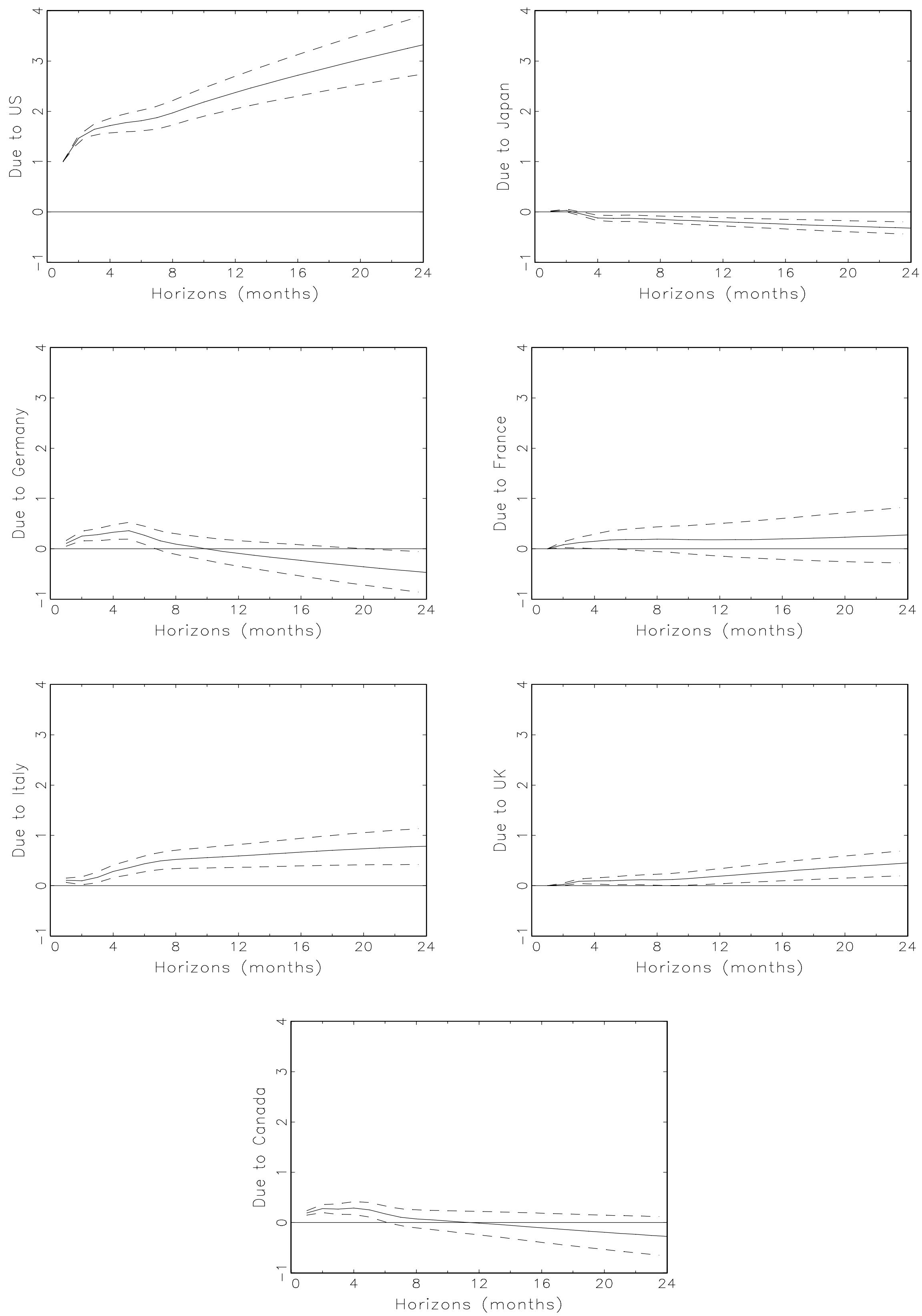

Figure 2. Impulse Responses of U.S. Inflation to UCII in all G-7 Countries

Notes: Solid lines are point estimates of the impulse responses, and dashed lines represent one-standard-deviation bands. The vertical axes measure the cumulative effects on U.S. inflation (percentage changes in prices) due to a 1 percent unexpected increase in prices in all seven countries.

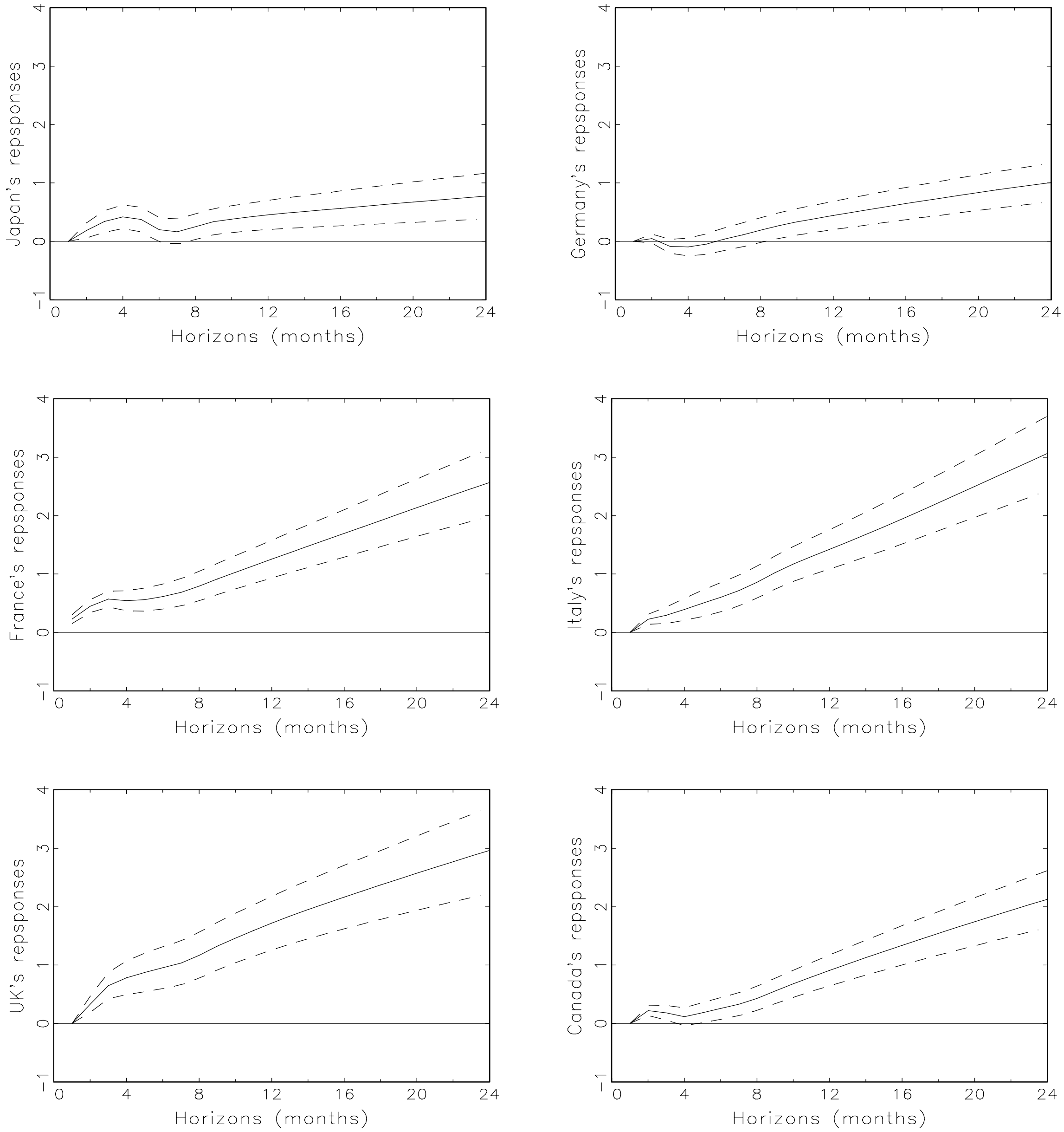

Figure 3. Impulse Responses of Inflation in other G-7 Countries to U.S. UCII

Notes: Solid lines are point estimates of the impulse responses, and dashed lines represent one-standard-deviation bands. The vertical axes measure the cumulative effects on inflation in the six non-U.S. G-7 countries (percentage changes in prices) due to a 1 percent unexpected increase in prices in the U.S.

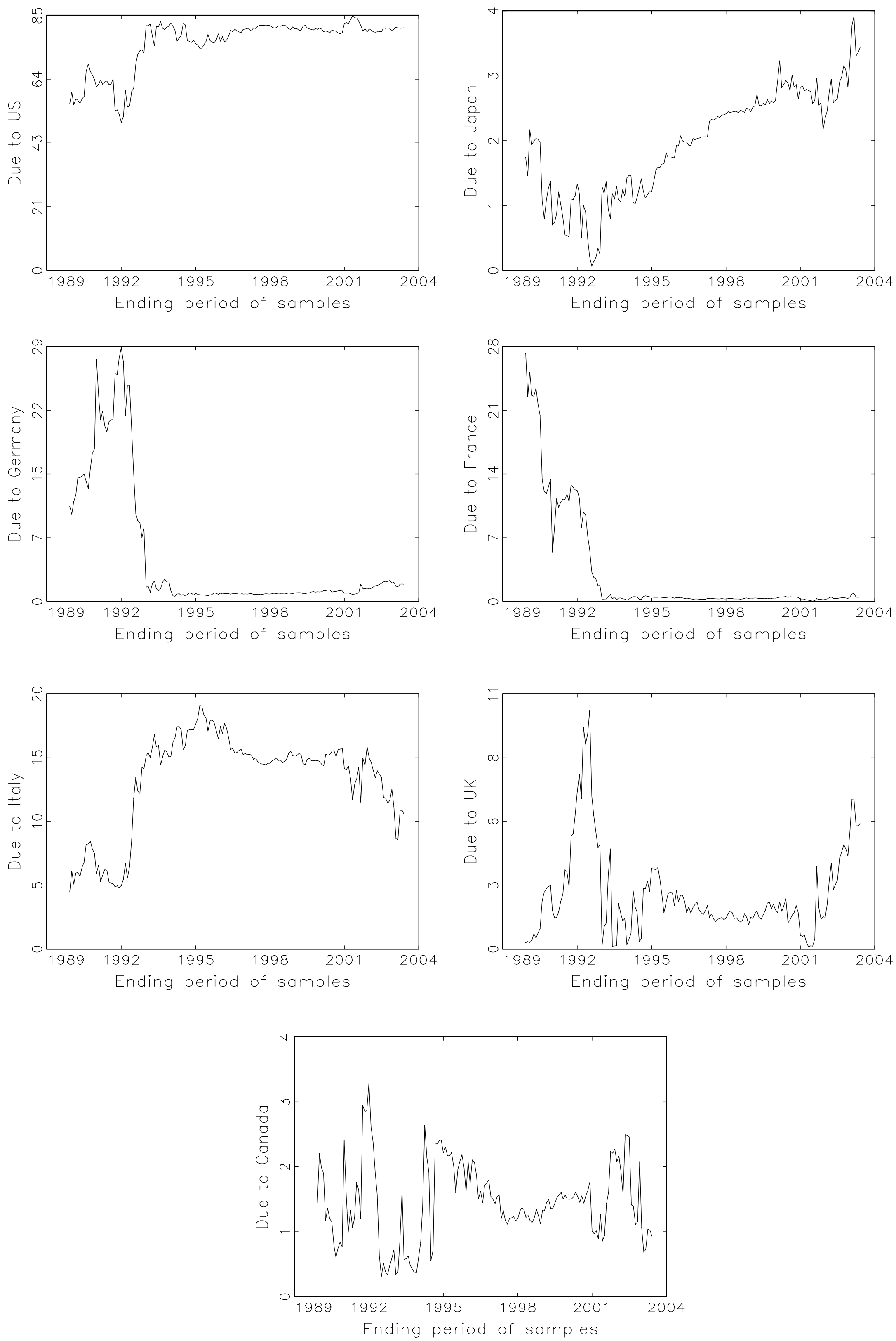
Due to US
Due to Japan
Due to Germany
Due to France
Due to Italy
Due to UK
Due to Canada
Ending period of samples
1989
1992
1995
1998
2001
2004

Figure 4. Recursively Estimated Forecast Error Variance Decompositions of U.S. Inflation due to UCII in all G-7 Countries at the 24-Month Horizon

Notes: The initial sample period is July 1973 to December 1989, and the variance decompositions are estimated recursively each month with an expanding sample. The final sample period is July 1973 to June 2003.

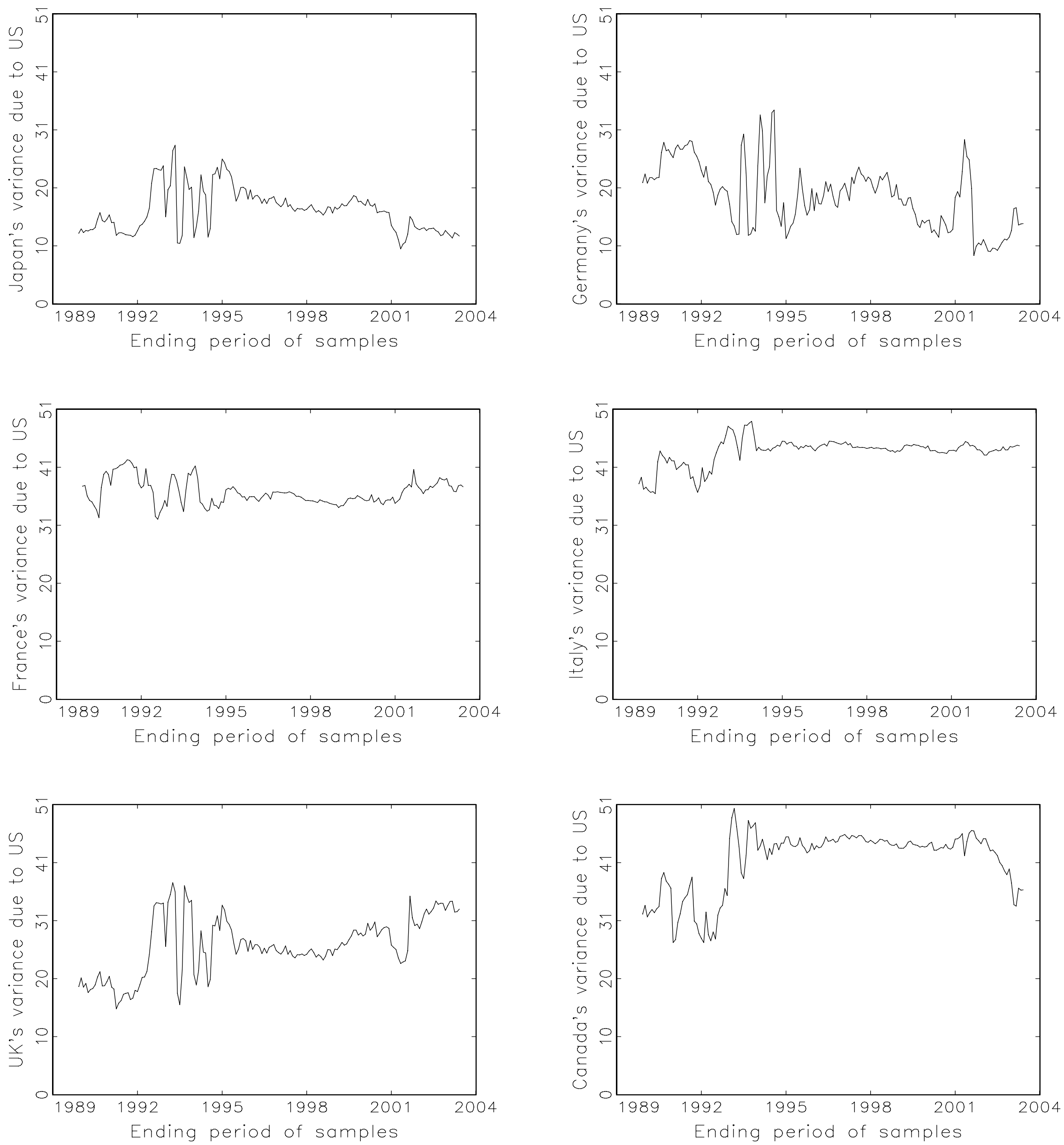

Figure 5. Recursively Estimated Forecast Error Variance Decompositions of inflation in all Other G-7 Countries due to UCII in the U.S. at the 24-month horizon

Notes: The initial sample period is July 1973 to December 1989, and the variance decompositions are estimated recursively each month with an expanding sample. The last sample period is July 1973 to June 2003.