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Pacific-Basin Finance Journal 7 (1999) 331–349

PACIFIC-BASIN FINANCE JOURNAL

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International price level linkages: Evidence from the post-Bretton Woods era

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Abstract

In this paper, we investigate the dynamic interdependence structure of national price levels of the Group of Seven (G-7) countries during the post-Bretton Woods era, i.e., 1973–1996. We find that a significant proportion of each country's domestic inflation rate variance is attributable to foreign inflation shocks, especially in the long run. Also, all foreign countries are found to import U.S. inflation during the sample period as they used to under the Bretton Woods system. The empirical findings imply that flexible exchange rates do not insulate the domestic price levels from foreign inflation shocks, which invalidates a key argument for the flexible exchange rate regime. © 1999 Elsevier Science B.V. All rights reserved.

JEL classification: F3

Keywords: Domestic price level; Foreign inflation shocks; Exchange rate innovations; Flexible exchange rate regime; G-7 countries

1. Introduction

For the last few decades both international financial and commodity markets have been steadily liberalized at both the global and regional levels. At the global level, GATT, which was recently succeeded by the World Trade Organization

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PII: S0927-538X(99)00011-6

(WTO), has played a key role in dismantling the import quotas and tariff worldwide, promoting duty-free trade, and extending the rules of free trade to cover financial services and intellectual properties. At the regional level, such formal arrangements as ASEAN, NAFTA, and EU have been instituted to promote economic integration. As a result of these developments, the world commodity and financial markets are likely to have become substantially integrated, forging close linkages among national markets.

Reflecting this trend toward a greater integration, a strand of papers examined the interdependence structure of national financial markets, e.g., (Eun and Shim, 1989; Hamao et al., 1990; King and Wadhnani, 1990). These studies generally documented a highly integrated, dynamic world financial market where the U.S. leads other national markets. In contrast, the interdependence structure of national commodity markets remains largely unexplored. It is pointed out that there exists a related strand of literature, such as (Adler and Lehmann, 1983) and (Abuaf and Jorion, 1990), mainly investigating the validity of the purchasing power parity for currency pairs. In the current paper, however, we are not concerned with the *bilateral* PPP issue but with the *multilateral* pattern of interactions among national price levels during the period of flexible exchange rates.

When the exchange rates were fixed, a price shock in a country is likely to be transmitted to other countries unless hampered by serious trade barriers. As discussed by Bordo (1993) and others, U.S. inflation in the late 1960s and early 1970s was indeed transmitted to other countries via fixed exchange rates. This situation contributed to the eventual collapse of Bretton Woods system in 1973. In the words of Bordo (p. 177): ''The only alternative to importing U.S. inflation was to float—a route taken by all countries in 1973.'' Presumably, under the flexible exchange rate system, U.S. inflation may be offset by the depreciating dollar, without being transmitted to other countries. Whether this situation really prevailed since 1973 is a matter of considerable interest.

In the current paper, we propose to fill the existing gap in the literature by investigating the following questions:

1. How much of a country's domestic inflation dynamics can be attributable to foreign-originating inflation shocks and exchange rate shocks?
2. How does a country's inflation rate adjust over time to foreign inflation shocks?
3. Do fluctuating exchange rates insulate a country's domestic price level from foreign inflation shocks as argued by the advocates of the flexible exchange rate regime?

In addressing the above questions, we interpret the monthly inflation rate data for the G-7 countries using the vector autoregression (VAR) and other time-series techniques. Our investigation is focused on establishing stylized facts about the *multilateral, dynamic* interactions among national price levels. Our sample period covers July 1973 through June 1996 during which the exchange rates were generally allowed to fluctuate.

The key empirical findings of the paper can be summarized as follows. First, the price levels of G-7 countries are found to be cointegrated, suggesting that these price levels are highly interlinked in the long-run. The U.S. is found to be the least susceptible to foreign shocks and Japan the most. Specifically, our estimation of the 7-variable VAR system indicates that at 24-month horizon, the proportion of domestic inflation rate variance attributable to foreign inflation shocks ranges from 17.2% for the U.S. to 73.0% for Japan; the proportion is 46.8% for Germany, 47.0% for France, 60.9% for Italy, 41.8% for the U.K., and 69.0% for Canada. No country's price level is exogenous to the VAR system in the long run.

Second, not surprisingly, the U.S. is found to be the most influential market. A U.S. price shock is 'positively' transmitted to each foreign country over time, showing that foreign countries are indeed *importing* the U.S. inflation. Likewise, the French and U.K. price shocks are positively transmitted to the U.S. However, shocks to the Canadian, German, Italian, and Japanese prices are actually transmitted negatively, eliciting deflationary U.S. responses. This unexpected result may arise when an inflationary foreign shock is accompanied by an *overshooting* depreciation of the foreign currency. Our impulse response analysis shows that the domestic price level reacts to a foreign price shock over an extended period time, attesting to a rather sticky nature of commodity prices.

Third, our estimation of the expanded 13-variable VAR system comprising both national inflation rates and exchange rates shows that exchange rate innovations collectively account for a significant fraction of each country's inflation variance, ranging from 18.9% for Canada to 50.1% for France at 24-month horizon. In addition, the relative importance of exchange rate innovations as opposed to foreign inflation innovations as sources of a country's domestic inflation rate variance tends to rise as the horizon gets longer. In other words, foreign inflation innovations tend to *pass-through* to the domestic price levels faster than exchange rate innovations, suggesting an intriguing possibility, i.e., exchange rate innovations may be perceived as more reversible than permanent, whereas the opposite may hold for foreign inflation innovations.

Fourth, our empirical findings have a few major policy implications. In view of the significant foreign influences, a country in isolation has a rather limited capability of controlling its domestic inflation dynamics in the long run. International coordination of monetary and exchange rate policies will be necessary to effectively control national inflation rates. In addition, highly 'interlinked' national price levels documented in this study do not support a key argument in favor of the flexible exchange rate regime advanced by Milton Friedman (1953): Exchange rate changes, if allowed, would absorb foreign-originating inflation shocks, insulating the domestic price level. Our findings suggest that the flexible exchange regime needs to be advocated on other grounds.

The rest of the paper is organized as follows. In Section 2, we describe the data and perform some preliminary statistical analyses including the unit root and cointegration tests. Section 3 presents the main empirical findings of the VAR analysis. Lastly, Section 4 offers concluding remarks.

2. Preliminary analyses

2.1. Summary statistics

The empirical analysis covers the post-Bretton Woods period, a 23-year period between July 1973 and June 1996. We obtained the consumer price indexes and exchange rates from *International Financial Statistics* for G-7 countries. The descriptive statistics for the monthly inflation rates of each country are presented in Table 1. Panel A provides the mean, standard deviation, skewness, and kurtosis. The panel shows that countries have rather disparate inflation rates. Germany and Japan have the lowest mean inflation rates at less than 4% per annum, whereas Italy has the highest, 10.08% per annum, followed by the U.K. with 8.16% per annum. The inflation rates of the other three countries, i.e., U.S. (5.32%), Canada

Table 1
Summary statistics for monthly inflation rates (G-7 countries). Sample period (July 1973–June 1996)

Panel A: Descriptive statistics[a]

	Mean	S.D.	Skewness	Kurtosis	J.-Bera	$Q(6)$
US	0.0046	0.0035	0.7082	3.6323	27.57	426.26
Japan	0.0033	0.0073	1.4263	6.6623	246.92	35.97
Germany	0.0027	0.0029	0.6887	3.7234	27.74	90.96
France	0.0054	0.0040	0.4903	2.7526	11.72	798.68
Italy	0.0084	0.0059	1.1575	4.3811	83.26	525.01
UK	0.0068	0.0075	1.8133	7.9691	433.64	207.12
Canada	0.0048	0.0041	0.7654	4.9765	71.62	1583.00

Panel B: Correlation matrix of monthly inflation rates

	US	Japan	Germany	France	Italy	UK	Canada
US	–						
Japan	0.35	–					
Germany	0.40	0.27	–				
France	0.61	0.47	0.40	–			
Italy	0.51	0.40	0.38	0.67	–		
UK	0.39	0.46	0.41	0.56	0.37	–	
Canada	0.54	0.16	0.28	0.54	0.46	0.25	–

[a] The J.-Bera statistic is the Jarque and Bera (1980) test statistic for normality. The hypothesis of normality is rejected for all countries at any conventional significance level. The $Q(k)$ is the Ljung–Box statistic. We reject the null hypothesis that all auto-correlations up to lag k are jointly zero if the test statistic, $Q(k)$, is above the critical value.

(5.76%), and France (6.48%), fall in the medium range. The standard deviation of monthly inflation rate ranges from 0.29% for Germany to 0.75% for the U.K. Among G-7 countries, Germany is thus found to have the lowest mean inflation rate as well as the lowest volatility. The sample moments also indicate that the distributions of monthly inflation rates are all skewed and leptokurtic when compared to the normal distribution. Also, the Jarque and Bera (1980) test of normality confirms that the monthly inflation rate of each G-7 country is not normally distributed. Furthermore, the hypothesis that all correlation coefficients for lags up to sixth order are jointly zero is rejected for all countries based on the Ljung–Box *Q*-statistic. Thus, the monthly inflation rate series for all countries are serially correlated over time during the post-Bretton Wood period.

Panel B of Table 1 provides the correlation matrix of monthly inflation rates of G-7 countries. Some pairwise correlations are surprisingly high. For example, the correlation is 0.67 for France/Italy, 0.61 for France/U.S., 0.56 for France/U.K., and 0.54 for Canada/U.S. as well as for Canada/France. It is noted that France and the U.S. tend to have relatively high correlations with other countries, whereas Germany and Japan low correlations.

2.2. Unit root tests

We first test the hypothesis that the price level of each country contains a unit root so that it is nonstationary and thus follows a random walk process. To test for the presence of unit roots, we use the augmented Dickey–Fuller (ADF) test and the Philips–Perron (PP) test. The regression equations for the ADF and PP tests are as follows,

$$\text{ADF test:}\ \Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \sum_{k=1}^{k} \gamma_k \Delta y_{t-k} + \varepsilon_t, \tag{1}$$

$$\text{PP test:}\ y_t = \beta_0 + \beta_1 y_{t-1} + e_t, \tag{2}$$

where y_t represents the price level of each country, ε_t is a white noise error term, and Δ is the first difference operator. The null hypothesis of unit root (i.e., nonstationarity) is $\alpha_1 = 0$ for ADF test and $\beta_1 = 1$ for PP test. The two tests differ from each other regarding the assumptions made concerning the distribution of error terms. [1] However, the same critical values, represented as the τ statistic in Fuller (1976), are used to test the null hypothesis of unit root against the stationarity of price levels.

The results of unit root tests are reported in Table 2. Specifically, the table tabulates test statistics (τ-statistics) of the estimates of α_1 and β_1. For the ADF

[1] While the ADF test assumes that the disturbance terms are uncorrelated and have constant variance, the PP test allows for heteroskedasticity and serial correlations in error terms.

Table 2
Preliminary test: Unit root test

Price level	Dickey–Fuller test		Philips–Perron test	
	$k=3$	$k=4$	$k=3$	$k=5$
US	1.00	−0.99	−1.00	−0.93
Japan	−7.00*	−7.12*	−11.54*	−9.16*
Germany	−0.93	−0.71	−0.90	−0.87
France	−2.43	−2.38	−2.87*	−2.47
Italy	−1.19	−0.89	−1.89	−1.58
UK	−1.08	−1.12	−0.93	−0.88
Canada	−2.02	−1.80	−2.11	−1.86

Dickey–Fuller test: $\Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \sum_{k=1}^{k} \gamma_k \Delta y_{t-k} + \varepsilon_t$.
Philips–Perron test: $y_t = \beta_0 + \beta_1 y_{t-1} + e_t$.
y_t represents the price level of each country. Δ is the first difference operator. Tabulated is $\tau(b)$, the t-statistic for the parameter estimate α_1 and β_1 (the ratio of the OLS estimate to its standard error). Special critical values are tabulated in (Fuller, 1976). If the number of observations is more than 250, the critical value is -2.87 (-3.44) at 5% (1%) level of significance. For the time series in question, if $\tau(b)$ is below the critical value, then we reject the null hypothesis of nonstationarity in favor of the alternative hypothesis of stationarity.
* Indicates significance at 1% level.

test, we choose two lag lengths, i.e., $k=3$ and $k=4$. For PP test, we need to determine truncation lag q for the process of calculating test statistics which account for the serial correlation in e_t. We choose two truncation lags, i.e., $q=3$ and $q=5$. The critical values are -2.87 (-3.44) at 5% (1%) significance level with more than 250 observations.

We fail to reject the null hypothesis of unit root at 5% significance level with the exception of Japan since test statistics are not below the critical value (sufficiently negative). It means that price levels of the G-7 countries, with the exception of Japan, are nonstationary.[2] Our unit root test results imply, among other things, that a domestic price shock tends to be transitory in Japan but permanent in other G-7 countries.

2.3. Test for cointegration

We perform Johansen's cointegration test (Johansen, 1991) to see whether the price levels of G-7 countries are cointegrated even though price levels of individ-

[2] Although not reported here, unit root tests with trend as well as drift were also performed based on τ_τ (Dickey and Fuller, 1979) and ϕ_2 (Dickey and Fuller, 1981). Both test statistics fail to reject null hypothesis of unit root with the exception of Japan.

ual countries are $I(1)$, i.e., nonstationary. [3] The Johansen test is a multivariate generalization of augmented Dickey–Fuller test. Consider a VAR of order p:

$$y_t = a + \sum_{k=1}^{P} A_k y_{t-k} + e_t, \tag{3}$$

where y_t is a $n \times 1$ vector. By subtracting y_{t-1} from each side, we obtain

$$\Delta y_t = a + \sum_{k=1}^{p-1} \pi_k \Delta y_{t-k} + \pi y_{t-k} + e_t, \tag{4}$$

where

$$\pi = \sum_{k=1}^{p} A_k - I; \;\; \pi_k = \sum_{j=1}^{k} A_j - I.$$

The Johansen test of cointegration is focused on the rank of estimated matrix π and its characteristic roots. If rank $(\pi) = 0$, then there is no cointegration relationship among the variables and the usual VAR specification in first differences is appropriate. If rank $(\pi) = r$ $(r < n)$, where n represents the number of variables in the system, there exists r cointegrating relations (or vectors). It also implies πy_{t-k} is the error correction terms which modify the conventional VAR model in first difference form to reflect the equilibrium relationships in levels.

To determine the size of rank (r) for estimated matrix π, Johansen developed two likelihood ratio test statistics: Trace statistic (λ_{trace}) and maximum eigenvalue statistic $(\lambda_{\max})$. The critical values for more than five variables are tabulated in Osterwald-Lenum (1992).

The results of cointegration tests are reported in Table 3. The null hypothesis that the price levels are not cointegrated $(r = 0)$ against the alternative of one or more cointegrating vectors $(r > 0)$ is clearly rejected since the $\lambda_{\text{trace}}(0)$ statistic (280.46 with 3 lags and 259.65 with 6 lags) exceeds critical values (124.24 and 133.57 at 5% and 1% significance levels respectively). However, the λ_{trace} statistic suggests no more than four cointegrating vectors since we fail to reject H_0 of $r \leq 4$, i.e., $\lambda_{\text{trace}}(4)$ statistic (31.82 with 3 lags and 27.94 with 6 lags) is less than critical value of 35.65 at 1% significance level. The maximum eigenvalue test based on $\lambda_{\max}(r)$ provides similar results. The null hypothesis of no cointegration $(r = 0)$ is rejected in favor of cointegration since $\lambda_{\max}(0)$ statistic (102.27 with 3 lags and 84.85 with 6 lags) exceeds critical values (45.28 and 51.57 at 5% and 1% significance levels respectively). In addition, the $\lambda_{\max}$ statistic indicates no more than three cointegrating vectors since we fail to reject H_0 of $r = 3$, i.e., $\lambda_{\max}(3)$

[3] Engle and Granger (1987) show that if variables are cointegrated, such econometric specification as the vector autoregression (VAR) in the first differences is misspecified.

Table 3
Preliminary test: Johansen cointegration test

Trace test [$\lambda_{trace}(r)$]						Maximum eigenvalue test [$\lambda_{max}(r)$]					
H_0	H_A	$k=3$	$k=6$	Critical values		H_0	H_A	$k=3$	$k=6$	Critical values	
				5%	1%					5%	1%
$r \leq 6$	$r > 6$	0	0.02	3.76	6.65	$r=6$	$r=7$	0	0.02	3.76	6.65
$r \leq 5$	$r > 5$	7.95	7.75	15.41	20.04	$r=5$	$r=6$	7.95	7.73	14.07	18.63
$r \leq 4$	$r > 4$	31.82	27.94	29.68	35.65	$r=4$	$r=5$	23.87	20.19	20.97	25.52
$r \leq 3$	$r > 3$	59.63	57.23	47.21	54.46	$r=3$	$r=4$	27.81	29.29	27.07	32.24
$r \leq 2$	$r > 2$	101.13	102.14	68.52	76.07	$r=2$	$r=3$	41.50	44.91	33.46	38.77
$r \leq 1$	$r > 1$	178.19	174.80	94.15	103.18	$r=1$	$r=2$	77.06	72.66	39.37	45.10
$r=0$	$r > 0$	280.46	259.65	124.24	133.57	$r=0$	$r=1$	102.27	84.85	45.28	51.57

The rank of estimated matrix π in Eq. (4) is r, representing the number of cointegrating vectors. k refers to number of lags in the VAR. H_0 (H_A) refers to null (alternative) hypothesis. Critical values are tabulated in (Osterwald-Lenum, 1992). Two likelihood ratio test statistics are obtained as follows: $\lambda_{trace}(r) = -T\sum_{i=r+1}^{n}\ln\left(1-\hat{\lambda}_i\right)$; $\lambda_{max}(r,r+1) = -T\ln\left(1-\hat{\lambda}_{r+1}\right) = \lambda_{trace}(r) - \lambda_{trace}(r+1)$ where T is the number of observation and λ_i is the estimated values of the characteristic roots from the π matrix.

statistic (29.29 with 6 lags) is less than critical value of 32.24 at 1% significance level.

3. The vector autoregression (VAR) analysis

3.1. Econometric specifications

Engle and Granger (1987) show that if variables are cointegrated, then the usual VAR(p) representation in first difference form is misspecified. Instead, they suggest a vector error correction representation as follows,

$$\Delta y_t = a + \sum_{k=1}^{P} A_k \Delta y_{t-k} - d(\alpha' y_{t-1}) + e_t, \tag{5}$$

where $\alpha' y_{t-1}$ are called the error correction terms and d is $n \times r$ matrix of coefficients. Since the price levels of G-7 countries are found to be cointegrated in preliminary tests, we estimate the vector error correction representation in Eq. (5) where Δy_t is a $n \times 1$ vector of monthly inflation rates, a is a $n \times 1$ vector of constant, A_k's are $n \times n$ coefficient matrices, e_t is the $n \times 1$ vector of serially uncorrelated mean zero innovations, and p is the lag length. [4]

[4] In order to determine the number of cointegrating vectors for the estimation of the error correction model, we evaluated the results of formal testing, innovation accounting, and residuals. We choose the number of cointegrating vector to be one.

Although VAR estimation is based on the autoregressive representation, most interpretations of VARs are based on moving average representation. By inverting or successive substitution, Eq. (5) has a moving average representation as follows:

$$\Delta y_t = \sum_{k=0}^{\infty} B_k e_{t-k}, \tag{6}$$

where e_t are serially uncorrelated, but are contemporaneously correlated with variance Σ. Furthermore, this representation is transformed into an orthogonal form in which Δy_t is a linear combination of current and past orthogonalized innovations as follows: [5]

$$\Delta y_t = \sum_{k=0}^{\infty} C_k u_{t-k}, \tag{7}$$

where u_t are contemporaneously and serially uncorrelated. This orthogonalization enables us to perform variance decomposition and impulse response analysis, which provides insight into the main channels of interaction among the countries in the system.

The variance decomposition of T-step-ahead forecast shows the proportion of total variance of each variable that is attributable to each of the orthogonalized innovations. Specifically, the forecast error variance in the T-step-ahead forecast of Δy_i, $\sum_{j=1}^{n}\sum_{k=0}^{T} C_{k,ij}^2$, can be decomposed into its respective sources, $\sum_{k=0}^{T} C_{k,ij}^2$, which is caused by innovations in Δy_j. The proportion measures the importance of the most recent inflation shocks (innovations) of a country to the inflation rate of other countries in the system, including its own country. The impulse response of the country i in k periods to an inflation shock generated in country j at time t is the i,jth element of the matrix C_k's.

3.2. Estimation results

Table 4 reports the decomposition of forecast error variance of each country. Specifically, the table provides the decomposition of 1-, 3-, 6-, 12-, and 24-month ahead forecast error variances of inflation rates into fractions that are attributable to innovations in each of the seven countries.

Overall, Table 4 reveals a highly interlinked world commodity market, especially in the long run. At 24-month horizon, for instance, the proportion of domestic inflation rate variance that can be collectively attributable to foreign inflation innovations ranges from 17.2% for the U.S. to 73.0% for Japan. The

[5] The orthogonalized innovation u_t has an identity covariance matrix. In order to obtain the orthogonalized innovation, we assume the relationship such that $e = \upsilon u$, where $E(ee') = \Sigma$ and $\upsilon\upsilon' = \Sigma$. The matrix υ can be obtained in several different factorizations of a positive definite matrix Σ. Most literature uses the Choleski factorization (a lower triangular matrix). For detailed exposition, see (Eun and Shim, 1989; Cooley and LeRoy, 1985).

Table 4
Decomposition of inflation innovations. Sample period (July 1973–June 1996)

Country explained	Horizon (in months)	By innovations in							
		U.S.	Japan	Germany	France	Italy	U.K.	Canada	Foreign[a]
U.S.	1	100.0	0	0	0	0	0	0	0
	3	96.9	0.1	0.7	1.5	0	0.7	0.1	3.1
	6	88.7	0.3	1.8	6.5	0	2.3	0.4	11.3
	12	81.2	1.1	0.6	10.8	0.1	4.2	2.0	18.8
	24	82.8	1.7	1.1	8.9	0.3	2.2	2.9	17.2
Japan	1	3.0	97.0	0	0	0	0	0	3.0
	3	7.0	87.9	0.1	0	1.7	1.0	2.2	12.1
	6	19.3	60.5	11.0	0.3	4.4	1.5	3.0	39.5
	12	21.2	43.0	15.3	2.4	6.7	3.3	8.1	57.0
	24	32.6	27.0	16.0	4.2	4.3	5.9	10.0	73.0
Germany	1	14.4	0.3	85.3	0	0	0	0	14.7
	3	14.2	0.8	83.8	0.3	0.2	0.5	0.2	16.2
	6	13.8	0.9	84.3	0.3	0.2	0.2	0.2	15.7
	12	23.0	0.9	71.7	2.0	2.0	0.1	0.3	28.3
	24	41.2	0.6	53.2	2.4	2.3	0.1	0.1	46.8
France	1	15.4	0.2	1.4	83.1	0	0	0	16.9
	3	19.8	0.4	2.9	75.6	0.7	0.4	0.1	24.4
	6	17.1	0.8	4.9	72.2	3.0	1.9	0.1	27.8
	12	24.9	1.9	5.1	59.9	5.3	2.6	0.4	40.1
	24	33.3	2.7	2.4	53.0	5.2	2.4	0.9	47.0
Italy	1	2.4	1.3	0	0	96.2	0	0	3.8
	3	3.8	1.2	1.1	1.3	92.4	0	0.1	7.6
	6	5.5	2.0	1.9	3.8	86.6	0.1	0.1	13.4
	12	12.9	3.0	1.5	20.7	60.8	0.9	0.2	39.2
	24	27.1	0.8	1.6	30.2	39.1	1.1	0.1	60.9
U.K.	1	0.4	1.8	5.7	8.1	0	83.9	0	16.1
	3	4.2	0.9	4.8	7.3	1.4	80.1	1.3	19.9
	6	7.0	0.5	4.6	8.9	6.8	70.8	1.4	29.2
	12	11.2	0.7	2.1	11.8	7.6	63.8	2.9	36.2
	24	20.6	1.0	1.2	11.2	4.5	58.2	3.3	41.8
Canada	1	0.4	0	0.8	0.2	0.1	0.9	97.6	2.4
	3	3.4	0.1	2.6	1.9	0.1	0.3	91.6	8.4
	6	2.7	0.1	1.7	5.0	0.3	0.9	89.2	10.8
	12	7.7	1.7	2.3	15.5	0.7	4.9	67.2	32.8
	24	26.9	4.5	5.8	22.5	2.4	7.0	31.0	69.0

[a]This column provides the percentage of forecast error variance of each country explained collectively by the ‘foreign’ countries.

proportion is 46.8% for Germany, 47.0% for France, 60.9% for Italy, 41.8% for the U.K., and 69.0% for Canada. These results indicate that in each country, a surprisingly large fraction of domestic inflation variance is attributable to foreign-originating shocks. The U.S. appears to be the only exception. The U.S. domestic inflation rate is the least susceptible to foreign innovations probably due to the sheer size of the U.S. economy and the leading role it plays in the world economy.

Also noteworthy from Table 4 is the fact that the degree of interdependence among national inflation rates tends to increase rather sharply as the horizon gets longer. Take Italy for instance. The fraction of Italian inflation variance attributable to foreign innovations is only 3.8% at 1-month horizon and 7.6% at 6-month horizon. But it rises to 39.2% at 12-month horizon and 60.9% at 24-month horizon. Other countries also exhibit the qualitatively similar patterns. One thus may say that in the short run, national inflation movements are largely due to domestic factors, but in the long run, they become highly interdependent internationally. This pattern may reflect the fact that commodity prices are relatively sticky in the short run, but become flexible in the long run.

As can be expected, the U.S. emerges as the most influential among individual countries. For example, the U.S. inflation innovations account for 32.6% of the Japanese and 41.2% of the German inflation variance at 24-month horizon. Meanwhile, few countries exert a significant influence on the U.S. inflation, with the possible exception of France whose innovations account for 8.9% of the U.S. inflation variance at 24-month horizon.

Despite its position as the world's second largest economy, Japan does not exert a significant influence on any country. In contrast, the Japanese domestic inflation variance is significantly attributable to innovations in other countries, e.g., the U.S. (32.6%), Germany (16.0%), Canada (10.0%), and the U.K. (5.9%). Considering this asymmetric pattern of interactions, one may say that Japan is rather a 'passive' market. Also, in view of the frequent reference to the closed nature of Japanese market, foreign influences on the Japanese domestic inflation are found to be surprisingly heavy.

A few other interesting patterns are noticeable as well. In accounting for the German inflation rate, no country, except the U.S., plays a significant role. In accounting for the Italian inflation rate, however, France plays a more influential role than the U.S. in the long run. Specifically, at 24-month horizon the French (U.S.) innovations account for 30.2% (27.1%) of the Italian inflation rate movements. The significant French influence on the Italian inflation may be, in part, due to the cultural and geographical proximity between the two countries. The U.K. inflation movements are also found to be significantly influenced by the U.S. (20.6% at 24-month horizon) and French (11.2%) innovations.

A highly interactive nature of international commodity markets documented in Table 4 implies that flexible exchange rates failed to insulate these countries from foreign inflation rate shocks, especially in the long run. The Friedman (1953) argument that the flexible exchange rate regime is desirable in part because a country's domestic price level can be insulated from foreign price shocks by exchange rate changes is not supported by our findings.

3.3. Impulse response analysis

To obtain additional insight into the structure of international price level linkages, we now examine the pattern of impulse responses of each foreign market

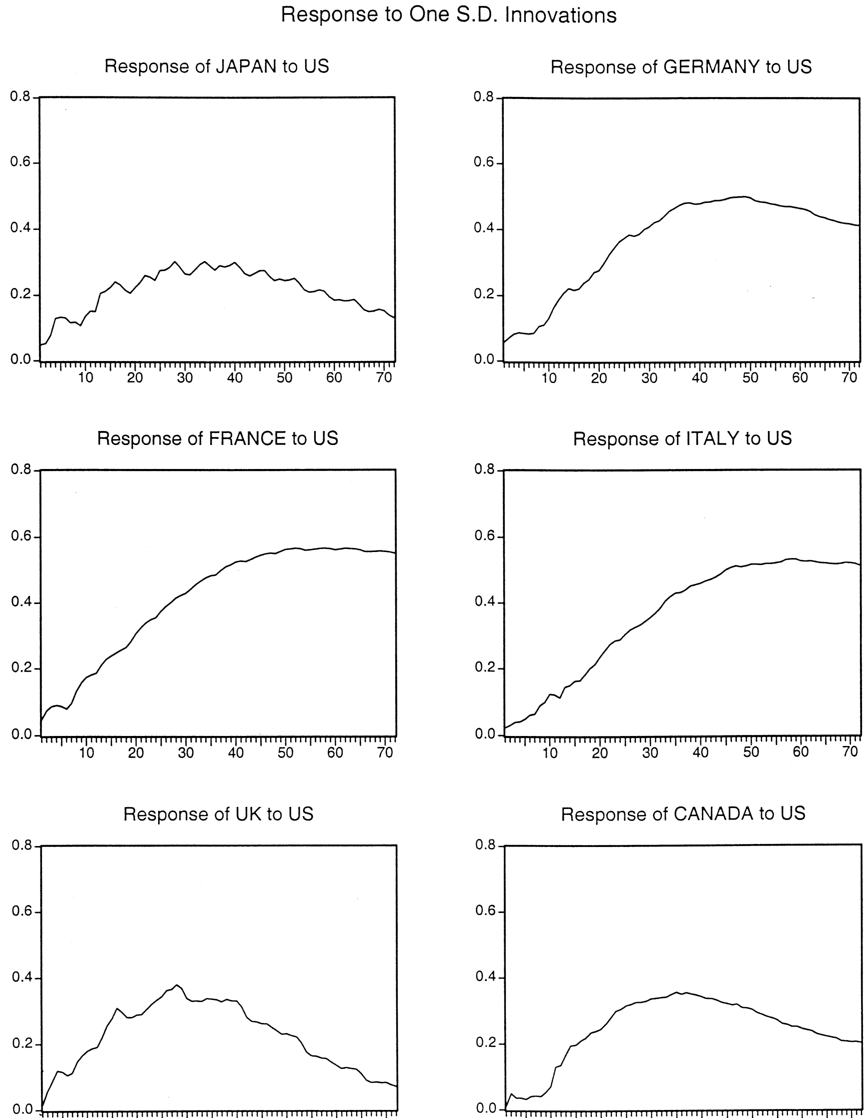

Fig. 1. Impulse responses of foreign prices to a U.S. shock (impulse responses are plotted vertically against the lapsed time measured in months).

to a shock in the U.S. market, which was shown to be the most influential, as well as that of U.S. responses to a shock introduced to each foreign market. We examine the impulse responses based on the simulated responses of the estimated

VAR system. For the ease of interpretation, we examine the impulse responses at the price level, rather than the inflation rate level.

Fig. 1 plots the time path of impulse responses of each foreign price level to a typical shock, i.e., one standard deviation shock, to the U.S. price level. As can be

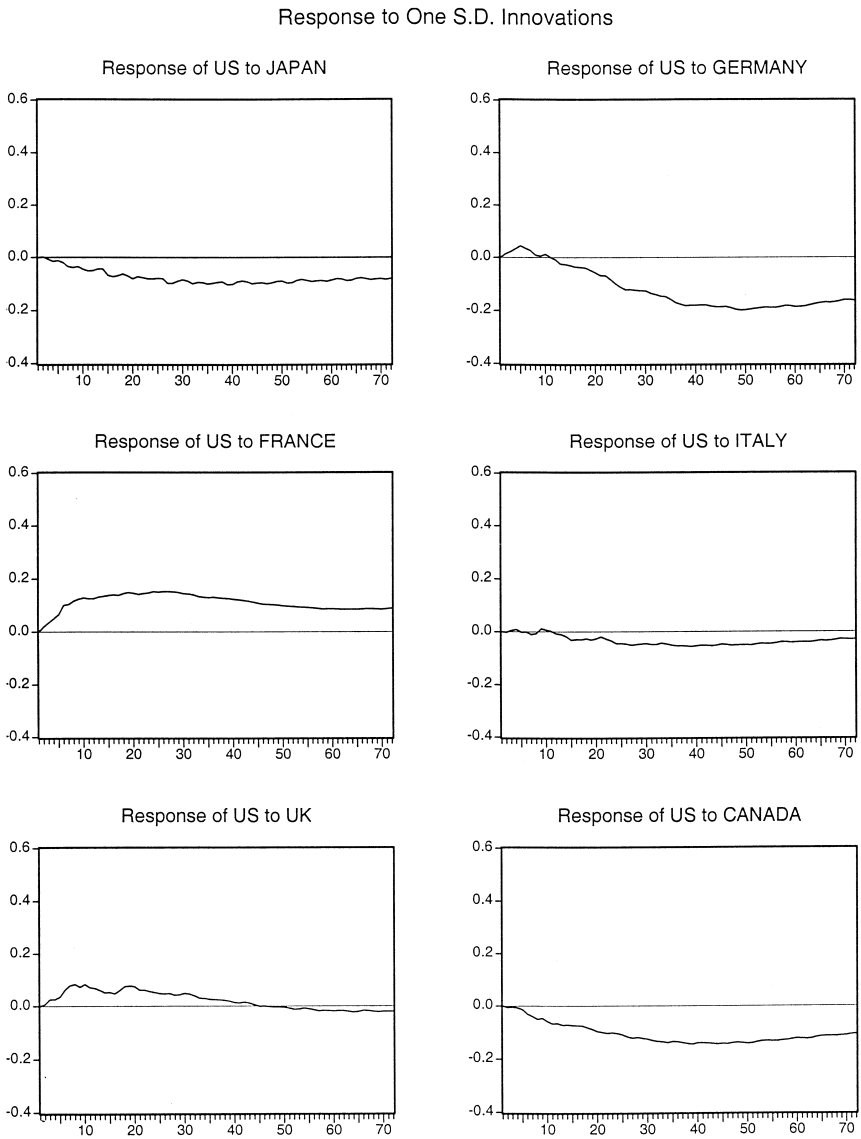

Fig. 2. Impulse responses of the U.S. price level to foreign shocks (impulse responses are plotted vertically against the lapsed time measured in months).

seen from the figure, a U.S. price shock is ‘positively’ transmitted to each market. In other words, foreign countries are clearly seen to *import* the U.S. inflation. The

Table 5
Decomposition of each country’s inflation innovations into their sources. Sample period (July 1973–June 1996)

Mos	Foreign inflation							Exchange rates							Domestic inflation
US															
	Jap.	Ger.	Fra.	Ita.	U.K.	Can.	Total	Jap.	Ger.	Fra.	Ita.	U.K.	Can.	Total	
1	0	0	0	0	0	0	0	0	0	0	0	0	0	0	100.0
3	0.3	0.4	0.9	0.9	0.3	0.2	3.0	0	0.2	0.4	0.4	0.3	4.1	5.4	91.6
6	1.5	1.4	0.8	1.5	0.5	0.6	6.3	1.5	0.5	1.4	1.3	1.8	10.4	16.9	76.8
12	3.1	0.6	0.3	3.0	1.3	1.7	10.0	7.6	0.3	3.4	1.3	3.3	10.0	25.9	64.1
24	3.9	3.0	1.3	3.1	0.5	4.3	16.1	7.2	0.5	6.0	1.3	4.2	9.6	28.8	55.1
Japan															
	U.S.	Ger.	Fra.	Ita.	U.K.	Can.	Total	Jap.	Ger.	Fra.	Ita.	U.K.	Can.	Total	
1	4.1	0	0	0	0	0	4.1	0	0	0	0	0	0	0	95.9
3	6.9	1.5	0.1	0.2	0.4	6.3	15.4	0.1	1.8	1.4	0.6	1.4	0.2	5.5	79.1
6	19.3	4.3	1.3	0.5	0.9	8.9	35.2	0.4	4.3	0.9	1.4	1.4	0.5	8.9	55.9
12	14.6	3.6	1.8	1.2	1.1	16.4	38.7	9.1	6.2	0.8	1.3	0.6	0.4	18.4	42.9
24	17.4	1.2	3.9	2.2	1.7	13.6	40.0	17.6	6.7	0.7	0.6	1.5	1.9	29.0	31.0
Germany															
	U.S.	Jap.	Fra.	Ita.	U.K.	Can.	Total	Jap.	Ger.	Fra.	Ita.	U.K.	Can.	Total	
1	18.3	1.9	0	0	0	0	20.2	0	0	0	0	0	0	0	79.8
3	13.1	2.6	1.7	0.1	0.2	0.2	17.9	0.9	0	0	2.0	0.1	4.0	7.0	75.1
6	6.8	1.1	6.4	0.4	1.3	0.5	16.5	2.9	2.3	0.1	4.8	0.1	8.3	18.5	65.0
12	6.0	0.6	6.0	0.2	1.6	0.3	14.7	7.8	3.7	3.9	3.7	0.7	12.8	32.6	52.7
24	12.6	0.7	3.5	0.2	2.7	1.3	21.0	9.0	4.4	5.7	3.1	0.3	15.5	38.0	41.0
France															
	U.S.	Jap.	Ger.	Ita.	U.K.	Can.	Total	Jap.	Ger.	Fra.	Ita.	U.K.	Can.	Total	
1	17.1	0.7	0.2	0	0	0	18.0	0	0	0	0	0	0	0	82.0
3	27.5	1.2	0.6	0	3.5	0.9	33.7	1.2	1.5	0.4	0.1	0	0.5	3.7	62.6
6	26.5	1.4	0.9	0.2	5.9	0.7	35.6	10.8	7.4	1.1	0.5	0.8	0.4	21.0	43.4
12	27.8	2.6	3.6	0.8	2.9	0.5	38.2	16.9	8.7	14.7	0.6	2.7	1.7	45.3	16.5
24	26.2	1.2	6.6	0.7	1.8	2.1	38.6	12.1	10.5	24.5	1.3	1.1	0.6	50.1	11.3
Italy															
	U.S.	Jap.	Ger.	Fra.	U.K.	Can.	Total	Jap.	Ger.	Fra.	Ita.	U.K.	Can.	Total	
1	4.8	2.7	1.0	6.4	0	0	14.9	0	0	0	0	0	0	0	85.1
3	7.3	2.2	1.0	8.5	0.9	1.4	21.3	0	0.3	0.4	0.2	0.1	0.3	1.3	77.4
6	4.9	0.9	1.8	18.5	0.7	1.7	28.5	0.7	0.3	0.3	0.9	0.2	0.5	2.9	68.6
12	5.6	0.8	0.8	17.6	0.7	3.7	29.2	4.7	1.5	2.1	1.1	0.8	1.5	11.7	59.1
24	12.1	2.3	0.4	13.3	0.3	1.8	30.2	11.6	1.6	9.6	0.3	0.2	1.3	24.6	45.2

Table 5 (continued)

Mos	Foreign inflation							Exchange rates							Domestic inflation
U.K.															
	U.S.	Jap.	Ger.	Fra.	Ita.	Can.	Total	Jap.	Ger.	Fra.	Ita.	U.K.	Can.	Total	
1	2.7	1.9	2.7	3.7	0.1	0	11.1	0	0	0	0	0	0	0	88.9
3	12.4	3.6	1.6	1.8	0.5	0.1	20.0	0.3	0.1	0.5	0.9	0	1.3	3.1	76.9
6	17.5	3.1	2.3	4.0	2.9	1.4	31.2	0.8	0.2	2.7	3.8	0	3.6	11.1	57.7
12	19.1	2.5	1.9	3.5	4.0	1.6	32.6	9.0	0.2	1.4	6.9	1.0	1.5	20.0	47.4
24	31.5	1.8	4.4	1.4	3.9	0.8	43.8	13.7	0.1	0.5	6.0	0.7	0.8	21.8	34.4
Canada															
	U.S.	Jap.	Ger.	Fra.	Ita.	U.K.	Total	Jap.	Ger.	Fra.	Ita.	U.K.	Can.	Total	
1	0.2	0.4	0.1	1.0	0.5	1.7	3.9	0	0	0	0	0	0	0	96.1
3	6.9	1.9	1.7	4.3	0.4	0.6	15.8	3.5	0.3	0.4	0.9	1.6	0	6.7	77.5
6	6.9	3.3	1.2	9.8	0.2	0.3	21.7	3.2	0.2	3.8	0.9	3.5	0	11.6	66.7
12	11.8	2.1	6.1	11.6	1.0	3.0	35.6	5.9	0.2	6.3	1.0	1.9	0.1	15.3	49.1
24	26.5	1.3	17.9	9.5	0.5	7.0	62.7	1.7	0.1	11.9	4.2	0.7	0.3	18.9	18.4

magnitude of impulse responses is low initially but increases over time before it levels off and/or finally starts to decline. For example, the U.K. (Canadian) impulse response to a U.S. shock steadily increases and peaks in 28 (35) months and starts to decline thereafter. In the case of Japan (Germany), the peak is reached roughly during 28–34 (45–50) months after the initial U.S. shock. In the case of France (Italy), the impulse response steadily grows stronger until it levels off in around 50 (55) months. Overall, Fig. 1 again attests to world commodity markets that are relatively sticky and insular in the short run but highly interlinked in the long run.

Fig. 2 plots the time paths of the U.S. price responses to a typical shock introduced to each foreign price level. As can be seen from the figure, the U.S. price level tends to respond positively to the French and U.K. shocks, but negatively to the Japanese, German, Italian and Canadian shocks. A foreign price shock thus is not always positively transmitted to the U.S. This result is rather unexpected in that an inflationary foreign shock can actually elicit a deflationary U.S. reaction. This might happen 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. [6] It is noted, however, that

[6] The recent 'overshooting' devaluation of certain Asian currencies against the U.S. dollar also appears to have resulted in negative transmission of foreign inflation, exerting a deflationary pressure on the U.S. economy.

the magnitude of U.S. responses, whether positive or negative, to foreign price shocks is rather modest over all horizons.

3.4. The effect of exchange rate innovations

In the preceding VAR analysis, we have not explicitly considered the effect of exchange rate changes. In this section, we expand our analysis by adding the U.S. dollar exchange rates of six foreign currencies. Our analysis in this section thus is based on the estimated VAR system comprising 13 variables—seven national price levels and six exchange rates.

Suppose the U.S. imports a variety of French products, and the French franc appreciates against the U.S. dollar. This situation may lead to a higher price level in the U.S. even if there is no change in the French domestic price level. Since the exchange rate may change irrespective of price level changes, it would be useful to explicitly consider the effect of exchange rate changes on the international price level dynamics. In estimating the 13-variable VAR system, we first stack the inflation rates and thereafter exchange rate changes. We thus examine the additional explanatory power of exchange rate innovations.

Table 5 provides the decomposition of each country's inflation innovation into their sources. A few things are noteworthy. First, exchange rate innovations collectively account for a significant proportion of each national inflation variance, ranging from 18.9% for Canada to 50.1% for France at 24-month horizon; the proportion is 28.8% for the U.S., 29.0% for Japan, 38.0% for Germany, 24.6% for Italy, and 21.8% for the U.K. In fact, exchange rate innovations collectively account for a greater fraction of the U.S. domestic inflation variance than foreign inflation innovations at every horizon. It is noted that foreign inflation and exchange rate innovations together account for 44.9% of the U.S. inflation variance at 24-month horizon, implying that the U.S. inflation is far from self-generating.

Second, as can be seen from Fig. 3, the importance of exchange rate innovations relative to foreign inflation innovations as sources of the domestic inflation variance tends to increase as the horizon gets longer. Take Italy for instance. The fraction of the Italian inflation variance attributable to exchange rate innovations increases rather sharply from 1.3% at 3-month horizon to 11.7% at 12-month horizon to 24.6% at 24-month horizon. In contrast, the fraction of the variance attributable to foreign inflation innovations is 21.3% at 3-month horizon, rising gradually to 29.2% at 12-month horizon to 30.2% at 24-month horizon. With the exception of Canada, all the other foreign countries, i.e., Japan, France, Germany and the U.K. exhibit a similar pattern. In the cases of France and Germany, exchange rate innovations eventually become more important than foreign inflation innovations in accounting for the domestic inflation dynamics.

This particular pattern implies an interesting phenomenon: *foreign inflation innovations pass-through to the domestic inflation faster than exchange rate*

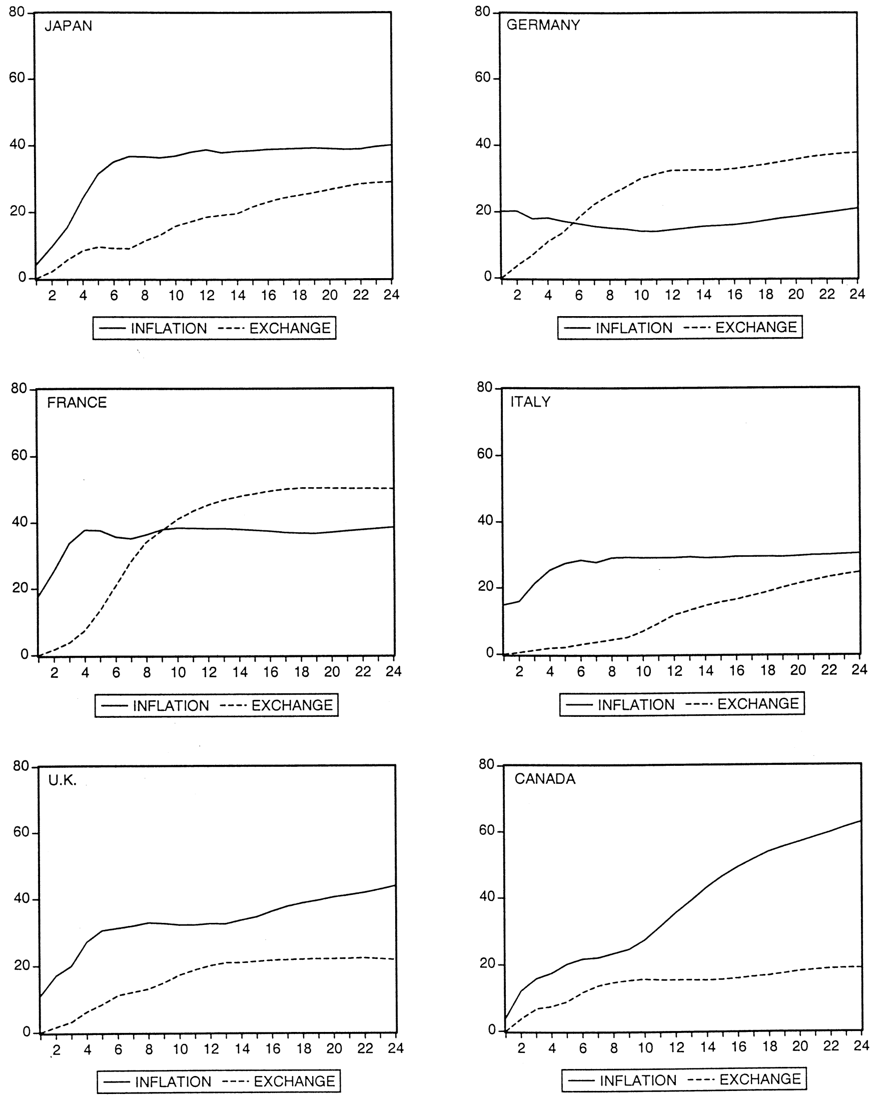

Fig. 3. Accounting for each country's inflation variance: foreign inflation vs. exchange rate innovations. [The vertical axis represents the percentages of inflation variance in each country which can be accounted for by foreign inflation (solid line) and exchange rate innovations (dotted line). The horizontal axis represents the forecasting horizon in months.]

innovations do. This phenomenon may arise if exchange rate innovations are viewed as more reversible than permanent, whereas foreign inflation innovations more permanent than reversible. Needless to say, it would be quite costly and impractical to continuously adjust the domestic prices of foreign imports as the exchange rate fluctuates up and down. When foreign prices change, however, it will be much less likely to be reversible and consequently, the domestic prices of those foreign imports would adjust more promptly.

Third, among all external factors (i.e., six foreign inflation and six exchange rates) influencing the U.S. inflation, the exchange rates of Canada and Japan, the two largest trading partners of the U.S., turned out to be the most important, followed by the French exchange rate and the Canadian inflation rate. In the case of Japanese inflation, the country's own exchange rate (17.6%), the U.S. inflation rate (17.4%) and the Canadian inflation rate (13.6%) emerge as the most important sources of innovations at 24-month horizon. Similarly, in the case of French inflation rate, the U.S. inflation rate (26.2%) and France's own exchange rate (24.5%) are found to be by far the most important sources of innovations at 24-month horizon. It is also noteworthy that, unlike the Japanese inflation rate exerting no major influences on any foreign inflation rates, the Japanese exchange rate innovations account for 9.0% of the German inflation variance, 12.1% of the French variance, 11.6% of the Italian variance, and 13.7% of the U.K. variance. Japan thus exerts its influences on foreign inflations not through its inflation innovations but through its exchange rate innovations. Further interpretation of the table is left to readers.

4. Concluding remarks

In this paper, we have investigated the interdependence structure of national price levels during the post-Bretton Woods era, i.e., 1973–1996, when the exchange rates were generally allowed to fluctuate. Unlike the previous literature, we have examined the linkages among national price levels in a framework that allows for dynamic, multilateral interactions.

National price levels were found to be cointegrated, suggesting that there exists an underlying equilibrium relationship binding these prices. In addition, the degree of interdependence among national price levels increases as the horizon gets longer. These findings imply that a country has a rather limited ability to control its domestic price level in the long run. This situation can be viewed as providing an incentive for countries to coordinate their monetary and exchange rate policies, curtailing national policy autonomy.

A highly interlinked nature of national price levels documented for the post-Bretton Woods era implies that exchange rate fluctuations failed to insulate the domestic price level from foreign price shocks. What's more, exchange rate innovations were found to actually add to the domestic inflation variance. Taken together, these findings substantially weaken the argument for the flexible exchange rate regime.

Acknowledgements

A summer research grant from the Howard University School of Business is gratefully acknowledged by Jin-Gil Jeong.

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