Impact of EMS Membership on Its Nominal Exchange Rate Volatility: An Approach with Univariate and Multivariate GARCH Models*

Kabsoo Hong

The Bank of Korea

This paper analyzes the exchange rate volatility after EMS was formed to stabilize bilateral nominal exchange rates among member countries. The analysis compares the exchange rate volatilities before and after the advent of the EMS, comparing members' currency volatilities with non-EMS currency volatilities.

Multivariate GARCH models as well as univariate GARCH models indicate that since March, 1979 the existence of the EMS has coincided with a marked reduction in the volatilities of exchange rates within the EMS. However, in the EMS versus non-EMS cases, or between-non-EMS currency cases, some countries show at least constant volatilities, even after March, 1979. This somewhat weakens claims that the reduction in volatility of inter-EMS currency rates has been a result of the creation of the system alone.

I. Introduction

A number of studies considered the evidence that the European Monetary System (EMS) has reduced exchange rate volatility. Ungerer *et al.* (1983) noted that "the exchange rate variability of EMS currencies has diminished since the introduction of the system..."[1] and updated the conclusion with a later paper, Ungerer *et al.* (1986). The European Commission (1982), Ungerer (1983), Dennis and Nellis (1984), Bank of England (1984), and Rogoff (1985) also studied the variability of EMS currencies.

In the notable study by Ungerer *et al.* (1983, 1986), various

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[1]Ungerer *et al.* (1983, pp. 8-9).

approaches to this question were used with various choices of exchange rates (bilateral, effective, nominal, and real), data frequency (daily, weekly, and monthly), and the level and change in exchange rates.

However, all of these studies which have tested for a downward shift in exchange-rate volatility for members of the EMS post-March, 1979 have generally relied on the unconditional distribution, independently and identically drawn from a normal distribution. It is by now an accepted fact that exchange-rate distributions tend to be leptokurtic (fat-tailed, highly-peaked) and that the variance shifts through time with new information available at time t-1, as noted by Taylor and Artis (1988). They applied non-parametric tests for volatility shifts which do not require actual estimation of the distribution parameters as well as tests for a shift in the conditional variance with a random walk with Autoregressive Conditional Heteroskedasticity (ARCH) disturbances. They found a significant reduction in the conditional variance of exchange rate for the EMS currencies against the Deutsche mark and signs of a significant rise in the conditional variance against the US dollar (see Taylor and Artis 1988, p. 12). However, they didn't demonstrate how to derive the likelihood ratio test which played a key role in their tests for shift in volatility. To derive a likelihood ratio test is not easy, given different observations and/or different distributions in each period i.e., pre-EMS and post-EMS. Also, after discussing the leptokurtic distribution in the exchange-rate change in one section, they ignored this distribution in their ARCH model and estimated the parameters under the normality assumption.

This paper will stress the importance of this stylized leptokurtic characteristic with the student t-distribution and also the possibility of a time-dependent conditional heteroskedasticity with multivariate GARCH (generalized ARCH) models as well as univariate models. I will test intra-EMS volatility against the Italian lira instead of the Deutsche mark (D-mark) to eliminate any possible impact of the role of the D-mark as a reserve currency or leading currency in the EMS. The US dollar will be used as a base currency to test the volatility change for non-EMS currencies, and the Pound Sterling will also be used to see whether there will be any difference in measuring volatility with a choice of a base currency.

In section II, I use unit root tests to check the stationarity in the weekly exchange-rate series. In section III, the univariate GARCH models are used to explain how the time-dependent conditional heteroscedasticity is built after diagnostic tests with Ljung-Box $Q(k)$ and $Q^2(k)$ statistics to check serial correlations. In the same section, the test results of the EMS currency volatility after March 13, 1979 are also analyzed, and in section IV, the multivariate GARCH models are estimated. Conclusions are given in section V.

II. Tests for a Unit Root in Weekly Exchange Rate Series

Autoregressive time series with a unit root have been the subject of much recent attention in the econometrics literature. In part, this is because the unit root hypothesis is of considerable interest, not only with data from financial and commodity markets where it has a long history, but also with macroeconomic time series. Initially, the research was confined to cases where the sequence of innovations driving the model is independent with common variance. Frequently, it was assumed that the innovations were i.i.d. $(0, \sigma^2)$ or, further, that they were i.i.d. $N(0, \sigma^2)$. However, independence and homoskedasticity are rather strong assumptions to make about the errors in most empirical econometric work. There is now a substantial body of research that exchange-rate series exhibit time-dependent heteroskedasticity.[2]

I have used the unit root test methods of Phillips (1987) and Phillips and Perron (1986, 1988) which are robust to a wide variety of serial correlation and time-dependent heteroskedasticity. These tests involve computing the OLS regressions:

$$s_t = \tilde{\mu} + \tilde{\beta}(t - T/2) + \tilde{\alpha} s_{t-1} + \tilde{u}_t \tag{1}$$

$$s_t = \overset{*}{\mu} + \overset{*}{\alpha} s_{t-1} + \overset{*}{u}_t \tag{2}$$

and

$$s_t = \hat{\alpha} s_{t-1} + \hat{u}_t \tag{3}$$

where s_t is the log of spot exchange rates, T denotes the sample size, and the innovation sequences $\tilde{u}_t$, $\overset{*}{u}_t$ and $\hat{u}_t$ are allowed to follow a wide variety of stochastic behavior including conditional heteroskedasticity. The testing strategy recommended by Phillips and Perron is to start equation (1) and to test the null hypothesis H_0^1: $\tilde{\mu} = 0$, $\tilde{\beta} = 0$, $\tilde{\alpha} = 1$ and H_0^2: $\tilde{\beta} = 0$, $\tilde{\alpha} = 1$ by means of the statistics $Z(\Phi_2)$ and $Z(\Phi_3)$ respectively. If H_0^1 and H_0^2 can be rejected, then one should next test H_0^3: $\tilde{\alpha} = 1$ by means of the $Z(t_{\tilde{\alpha}})$ statistic. If H_0^1

[2]See Baillie and Bollerslev (1989), Bollerslev (1987), Milhøj (1987), McCurdy and Morgan (1983) and the references therein.

and H_0^2 can not be rejected (i.e. they show both random walk and random walk with drift), then the strategy is to proceed to exclude the time trend and to test H_0^4: $\overset{*}{\mu} = 0$ and $\overset{*}{\alpha} = 1$ by the use of the $Z(\Phi_1)$ test statistic for testing a unit root without drift.

Individual unit root tests of the null hypothesis on (2) and (3) of the form H_0^5: $\overset{*}{\alpha} = 1$ and H_0^6: $\hat{\alpha} = 1$ are tested by the statistics $Z(t_{\overset{*}{\alpha}})$ and $Z(t_{\hat{\alpha}})$ (see Phillips and Perron (1988) for the precise formula for each test statistic).

In our analysis, I took weekly spot exchange rate data from the New York Foreign Exchange Market between January 3, 1973 and September 28, 1988. The series were constructed by taking observations every Wednesday, and in the even of the market being closed, an observation on the next business day (i.e. Thursday; if the market was closed on that Thursday also, then Friday, and so on) was used. The data provided by the Federal Reserve System, are bid prices taken at noon, constituting a total sample of 827 observations for the EMS currencies and 8 major countries[3] against the US dollar.

Six different unit root test statistics were estimated for all currencies. Calculating the test statistics requires that consistent estimates of the variances of the sum of the disturbances $\tilde{u}_t$, $\overset{*}{u}_t$ and $\hat{u}_t$ in (1) to (3) and a truncation lag, l, corresponding to the maximum order of non-zero autocorrelation in the disturbances be chosen; see Phillips and Perron (1986) and Newey and West (1987) for details. Hence, the statistics were computed for $l = 0, 2, 4, 6$ and 10, but were found to be remarkably similar for different values of l. The results with lag 10 are reported in Table 1.

Both simple unit root tests of the t-statistic type, $Z(t_{\overset{*}{\alpha}})$ and $Z(t_{\hat{\alpha}})$, confirm the unit root with drift. At the same time, the $Z(\Phi_1)$ statistics accepts the random walk without drift, and the inclusion of a time trend and use of the $Z(\Phi_2)$ statistics show the same results. However, the $Z(t_{\hat{\alpha}})$ statistics reject the random walk without a drift at the usual 95% level for the Swiss franc.

The overall indication is that there is strong evidence for the presence of unit root with a drift for all currency series, and hence, all the series appear to be stationary in their first differences.

[3]The EMS currencies in this paper include German D-mark, French franc, Italian lira, Belgian franc, Netherlands guilder, Irish pound, and Danish krone. The other major currencies include the US dollar with weighted value, Canadian dollar, Pound sterling, Austrian schilling, Swiss franc, Japanese yen, Swedish krona, and Norwegian krone.

TABLE 1
PHILLIPS-PERRON UNIT ROOT TESTS ON EXCHANGE RATE SERIES

$$S_t = \tilde{\mu} + \tilde{\beta}(t - T/2) + \tilde{\alpha} S_{t-1} + \tilde{u}$$

$$S_t = \overset{*}{\mu} + \overset{*}{\alpha} S_{t-1} + \overset{*}{u}_t, \quad S_t = \hat{\alpha} S_{t-1} + \hat{u}_t$$

$$Z(\Phi_2): \tilde{\mu} = 0, \ \tilde{\beta} = 1, \ \tilde{\alpha} = 1$$

$$Z(\Phi_3): \tilde{\beta} = 0 \text{ and } \tilde{\alpha} = 1, \quad Z(t_{\hat{\alpha}}): \hat{\alpha} = 1$$

$$Z(t_{\tilde{\alpha}}): \tilde{\alpha} = 1, \quad Z(\Phi_1): \overset{*}{\mu} = 0 \text{ and } \overset{*}{\alpha} = 1, \ Z(t_{\overset{*}{\alpha}}): \overset{*}{\alpha} = 1$$

Lag = 10 in Newey and West (1987)

Currencies (against U$)	$Z(\Phi_3)$	$Z(\Phi_2)$	$Z(t_{\tilde{\alpha}})$	$Z(\Phi_1)$	$Z(t_{\overset{*}{\alpha}})$	$Z(t_{\hat{\alpha}})$
Canadian $	0.8985	1.1039	−0.4762†	1.6543	−1.3388	−1.2617
Pound Sterling	1.1994	0.9803	−1.3970	1.3142	−1.4455	−0.7816
Irish Pound	0.9044	1.0334	−1.0969	1.4514	−1.2731	−1.1898
Italian lira	0.9932	1.9399	−0.8229†	2.8737	−1.3907	−1.7145
French franc	1.0638	0.7877	−1.4448	0.5014	−0.8800	−0.5495
Belgian franc	1.1551	0.7876	−1.4757	0.7641	−1.2138	−0.0503
Danish krone	1.0127	0.6799	−1.4258	0.5445	−1.0358	−0.2049
Deutsche mark	1.8170	1.5909	−1.8384	2.3862	−1.9050	0.9654
Dutch guilder	1.4877	1.2487	−1.6841	1.8510	−1.7106	0.7888
Swedish krona	1.0141	0.8579	−1.4159	0.5575	−0.7591	−0.7956
Aust. schilling	1.8367	1.6774	−1.8514	2.5031	−1.9090	0.9662
Swiss franc	2.6288	2.5395	−2.1850	3.5574	−2.1850	1.3835†
Japanese yen	1.3073	1.7751	−1.3598	1.3454	−0.1924	−1.6233
Norweg. krone	3.1458	2.1024	−2.3415	0.5861	−1.0746	−0.1906
Weighted-US dollar	0.8523	0.5976	−1.2980	0.8362	−1.2577	−0.3342

Note: 1. Under the null hypothesis, the 95% and 99% critical values of $Z(t_{\tilde{\alpha}})$, $Z(\Phi_3)$ and $Z(\Phi_2)$ are −0.94 and −0.33, 4.68 and 6.09, and 6.25 and 8.27, respectively, and values for $Z(t_{\tilde{\alpha}})$ are −3.96 and −3.44 at 1% and 5%. Also, at the 95% and 99% level the critical values of $Z(t_{\overset{*}{\alpha}})$, $Z(t_{\overset{*}{\alpha}})$ and $Z(\Phi_1)$ are 1.28 and 2.00, −0.07 and 0.6 and 4.59 and 6.43. For $Z(t_{\hat{\alpha}})$ and $Z(t_{\tilde{\alpha}})$, values are −2.58 and −1.95, − 3.43 and −2.86 at 1% and 5%, respectively (see Phillips and Perron 1986.).
2. †: indicates significant at the 0.95 percentile.

III. Models with Time-Dependent Conditional Heteroskedasticity: GARCH (1, 1)

For time series analysis, the autoregressive heteroskedastic process (ARCH) type of model has proven useful in several different economic applications. Among many others, see Engle (1982), Engle and Kraft (1982), Coulson and Robins (1985), Engle, Lilien and Robins (1987), and Weiss (1984). However in this paper, the

GARCH (Generalized Autoregressive Conditional Heteroskedasticity) is considered for empirical study of the EMS currency volatility, since it allows for a much flexible lag structure.[4]

A. Implication of GARCH Model

The first set of data consists of weekly exchange rates of EMS currencies against the US dollar and EMS currencies against the Dentsche mark from January 3, 1973 until September 28, 1988 for a total of 827 observations. The log of spot rates, s_t, are converted to continuously compounded rates of return,[5]

$$y_t = 1000 \times (s_t - s_{t-1}).$$

The dependent variable y_t denotes the change in the logarithm of the exchange rates between time t and $t - 1$ and is shown to be stationary in its first differences from the results of section II. The full model is then:

$$\begin{aligned} y_t &= b_0 + u_t \\ u_t \mid \Omega_{t-1} &\sim D(0, h_t) \\ h_t &= \omega_0 + \alpha_1 u_{t-1}^2 + \beta_1 h_{t-1} \end{aligned}$$

where Ω_{t-1} is the set of all relevant and available information at time $t - 1$, and where $D(0, h_t)$ represents some distribution with mean 0 and variance h_t. The assumed process is a regression model with innovations that have either conditional normal or student t densities with time-dependent variance. The conditional-variance equation is assumed to follow a generalized ARCH (or GARCH) model.

Before estimating the coefficients of GARCH models, the serial correlations are checked for implications of the GARCH model. First, most currencies were found to have moving average terms with significant levels. For example, the D-mark against the US dollar shows the value of $Q(10) = 22.5$ in the Ljung-Box (1978) portmanteau test statistic[6] for up to tenth-order serial correlation

[4]See Baillie and Bollerslev (1987) and Bollerslev (1986, 1987) for its applications with conditional t-distributed errors.

[5]For convenience of calculation, I multiplied 1000 by s_t, which doesn't change the statistical results.

[6]This is a test of the joint hypothesis that all autocorrelation coefficients are zero and as such as chi-square with k degrees of freedom.

in $(y_t - \hat{b}_0)$, which is very significant at any reasonable level in the corresponding asymptotic χ^2_{10} distribution. After adding those moving average disturbance terms, the value of $Q(10)$ is reduced to 8.5, which is not significant at any reasonable level (see Table 3, column 1). This $Q(10)$ reduction is the same for other currencies, with some exceptions, for example, the D-mark against the Italian lira and the D-mark against the Netherlands guilder need no moving average disturbance terms at all. After considering these moving average disturbance terms, we have

$$y_t = b_0 + \theta(L)u_t \tag{4}$$

$$\theta(L)u_t = \varepsilon_t + \theta_1 \varepsilon_{t-1} + \theta_2 \varepsilon_{t-2} \tag{5}$$

$$\varepsilon_t \mid \Omega_{t-1} \sim D(0, h_t) \tag{6}$$

$$h_t = \omega_0 + \alpha_1 \varepsilon^2_{t-1} + \beta_1 h_{t-1} \tag{7}$$

On the other hand, $(y_t - \hat{b}_0)^2$ is clearly not uncorrelated over time to all currencies, as reflected by the significant Ljung–Box test statistic for absence of serial correlation in the square, $Q^2(10)$, which is distributed asymptotically as a χ^2_{10} distribution (see McLeod and Li 1983). For example, when we don't use GARCH model, the D-mark against the US dollar shows $Q^2(10) = 21.2$, a very significant indication for the presence of serial correlation (see Table 3). The null hypothesis of no ARCH effects can be decisively tested with this $Q^2(k)$. Some series could have the squared residuals which appear to be autocorrelated even though the residuals do not (for our example, the Swiss franc against the US dollar; see Table 2). This absence of serial dependence in the conditional first moments, along with dependence in the conditional second moments, is one of the implications of the ARCH or GARCH (p, q) model (see Bollerslev 1986).

With the GARCH model we estimated the parameters by the Berndt *et al.* (1974) algorithm. The maximum likelihood estimates of the parameters are presented in Table 2 with asymptotic standard

$$Q(\hat{r}) = n(n+2)\sum_{k=1}^{M} \hat{r}_a^2(k)/(n-k) \sim \chi^2(M-p-q)$$

$$\text{where } \hat{r}_a(k) = \sum_{t=k+1}^{n} \hat{a}_t\hat{a}_{t-k} \Big/ \sum_{t=1}^{n} \hat{a}_t^2 \ (k = 1, 2, \ldots)$$

and $(1 - \phi_1 L - \ldots - \phi_p L^p)\omega_t = (1 - \phi_1 L - \ldots - \phi_q L^q)\, a_t$, where $\{a_t\} \sim$ i.i.d. $N(0, \sigma^2)$ with a discrete time series $\omega_1, \ldots, \omega_n$. In the case of $Q(10)$, critical values of those yield 18.307 and 15.987 at 5% and 10% level, respectively.

TABLE 2
ESTIMATION OF GARCH MODELS WITH DEUTSCHE MARK AND US DOLLAR AS BASE CURRENCY

$$y_t = b_0 + u_t;\ u_t = \varepsilon_t + \theta_1 \varepsilon_{t-1} + \theta_2 \varepsilon_{t-2};\ \varepsilon_t \mid \Omega_{t-1} \sim D(0, h_t, \nu);$$

$$h_t = \omega_0 + \omega_1 D_t + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1}$$

Parameters & Diagnostic Statistics	Inter-EMS			Non-EMS	
	DM-LIRA	DM-NGL	DM-FFR	US\$-CN\$	US\$-SFR
Log L	−2758.883	−2017.937	−2684.132	−2452.528	−3482.335
b_0	0.607**	0.047	0.816*	−0.209	0.641
	(0.238)	(0.062)	(0.382)	(0.169)	(0.608)
θ_1	—	—	0.219**	0.068*	0.069*
			(0.048)	(0.037)	(0.035)
θ_2	—	—	0.147**	0.117**	0.106**
			(0.035)	(0.036)	(0.026)
ω_0	19.397**	0.222**	66.990**	2.263**	156.463**
	(2.182)	(0.005)	(2.799)	(0.902)	(21.779)
ω_1	−14.580**	−0.209**	−40.981**	3.385**	140.164**
	(1.722)	(0.058)	(2.739)	(1.430)	(39.797)
α_1	0.266**	0.048**	0.274**	0.169**	0.311**
	(0.033)	(0.005)	(0.047)	(0.045)	(0.080)
β_1	0.631**	0.947**	—	0.694**	—
	(0.033)	(0.004)		(0.079)	
ν^{-1}	normal	normal	normal	0.201**	0.212**
				(0.007)	(0.056)
m^3	1.996	0.597	4.754	−0.868	0.240
m_4	13.109	8.941	51.352	8.788	5.583
$Q(10)$	12.404	10.296	17.116	12.631	6.209
$Q^2(10)$	18.750	4.571	0.382	8.582	141.254
$3(\hat{\nu}-2)/(\hat{\nu}-4)$	N.A.	N.A.	N.A.	9.0	11.36

Note: 1. Asymptotic standard errors are in parentheses under corresponding parameter estimates.
2. * indicates significance at the 5% level and ** at the 1% level.
3. DM stands for the Deutsche mark, NGL stands for the Netherlands guilder, CN\$ stands for the Chanadian dollar, SFR for the Swiss franc, and FFR for the French franc.
4. ν denotes the degree of freedom with student t density.

TABLE 2
(CONTINUED)
ESTIMATION OF GARCH MODELS WITH EMS CURRENCIES AGAINST THE US DOLLAR

Parameters & Diagnostics Statistics	Non-EMS				
	US$-UK£	US$-DM	US$-DKR	US$-NGL	US$-BFR
Log L	−3297.485	−3316.977	−3309.170	−3287.396	−3287.438
b_0	−0.671	0.543	−18.033**	0.422	0.216
	(0.567)	(0.515)	(0.513)	(0.452)	(0.536)
θ_1	0.009	0.080*	0.081*	0.078*	0.084**
	(0.037)	(0.036)	(0.038)	(0.037)	(0.034)
θ_2	0.078*	0.152**	0.119**	0.150**	0.140**
	(0.037)	(0.035)	(0.035)	(0.035)	(0.035)
ω_0	15.239*	7.066**	12.020**	6.124**	5.773*
	(3.481)	(0.168)	(4.883)	(2.819)	(2.551)
ω_1	14.626**	12.572*	18.715*	10.754*	12.079*
	(5.437)	(5.742)	(8.348)	(4.996)	(5.546)
α_1	0.109**	0.168**	0.185**	0.156**	0.153**
	(0.023)	(0.039)	(0.045)	(0.034)	(0.036)
β_1	0.768**	0.782**	0.726**	0.793**	0.798**
	(0.047)	(0.044)	(0.064)	(0.042)	(0.040)
ν^{-1}	normal	0.159**	0.169**	0.123**	0.152**
		(0.032)	(0.031)	(0.020)	(0.013)
m_3	−0.292	0.314	−0.067	0.349	0.321
m_4	6.561	6.093	4.190	6.257	5.947
$Q(10)$	12.666	6.129	4.190	4.903	5.861
$Q^2(10)$	3.714	3.484	4.453	4.903	3.290
$3(\hat{\nu}-2)/(\hat{\nu}-4)$	N.A.	5.650	6.115	4.464	5.319

Note: DKR stands for the Danish krone and BFR stands for the Belgian franc.

errors in parenthesis. The summary of the relevant test statistics are shown in Table 3; for example, the Ljung-Box test statistic for the standardized residuals, $\hat{\varepsilon}_t \hat{h}_t^{-1/2}$ and the standardized squared residuals, $\hat{\varepsilon}_t^2 \hat{h}_t^{-1}$, from the estimated GARCH (1, 1) model takes the values $Q(10) = 6.13$ and $Q^2(10) = 3.48$, respectively, for the D-mark against the US dollar, which doesn't indicate any further serial dependence. On the other hand, the hypothesis of the constant conditional variance fails with $LR_{\alpha=\beta=0}$ test statistics (see Table 4), which is highly significant at any level in the corresponding

TABLE 3
SUMMARY STATISTICS WITH THE IMPLICATION OF GARCH (1, 1) MODEL

	$y_t - \hat{b}_0$			GARCH (1, 1) − t		
	$Q(10)$	$Q^2(10)$	k	$Q(10)$	$Q^2(10)$	k
EMS Currency						
US\$-DM	8.52	21.21	11.72	6.13	3.48	6.09
US\$-DKR	8.45	34.52	10.77	4.19	4.46	5.21
US\$-NGL	9.93	33.43	13.01	6.26	4.90	5.12
US\$-BFR	7.64	29.45	12.39	5.86	3.29	5.95
Non-EMS Currency						
US\$-CN\$	18.84	27.98	17.28	12.63	8.58	8.79
US\$-SFR	3.03	145.52	6.20	6.21	141.25	5.58
UK£-US\$	10.94	38.43	6.07	12.67	3.71	6.56

Note: $Q(10)$ and $Q^2(10)$ denote the Ljung-Box (1978) portmanteau tests for up to tenth-order serial correlation in the levels and the squares which are standardized, respectively. They have χ^2 distributions with a degree of freedom of 10, which has values of 15.987 and 18.307 with $p = 0.10$ and 0.05, respectively. k is the usual measure of kurtosis given by the fourth sample moment divided by the square of the 2nd moment.

TABLE 4
LIKELIHOOD RATIO TESTS

	US\$-DM	US\$-NGL	US\$-DKR	US\$-BFR	US\$CN\$	US\$-SFR	US\$-UK£
$LR_{1/\nu=0}$	46.334	39.112	28.578	41.994	77.694	56.966	—
$LR_{\alpha=\beta=0}$	88.410	68.994	93.428	97.472	81.936	105.728	104.758

Note: For our reference, $\chi_1^2 = 6.638$ (with $P = 0.01$) and $\chi_2^2 = 9.210$ (with $P = 0.01$).

asymptotic χ_2^2 distribution.[7]

As can be seen from Table 2, the estimated values for $\alpha + \beta$ are close to 1[8] for some currencies, indicating the probable exist-

[7] I didn't show all test results in the Table for other exchange rates, but obviously they have the same results (see each Table).

[8] The GARCH (1, 1) process is wide-sense stationary iff $\alpha + \beta < 1$. See Bollerslev (1986) for the proof. The time series $\{X_t, t \in Z\}$, with index set $Z = \{0, \pm 1, \pm 2, \ldots\}$ is said to be wide-sense stationary or covariance stationary if

(i) $E|X_t|^2 < \infty$ for all $t \in Z$,

(ii) $EX_t = m$ for all $t \in Z$,

ence of an integrated GARCH, or IGARCH process (see Bollerslev 1989, Engle and Bollerslev 1986). The autoregressive term i.e. the coefficients of h_{t-1}^2 are highly significant, which tells us that changes in volatility of exchange rates have a high degree of persistence.

It is also interesting to note that the implied estimate of the conditional kurtosis,[9] $3(\hat{\nu} - 2)(\hat{\nu} - 4)^{-1}$ is in close accordance with the sample analogue for $\hat{\varepsilon}_t^4 \hat{h}_t^2$ (which is k in Table 3) for most currencies (see Table 2). This means that even in the weekly data, the t-distributed GARCH (1, 1) model works well.[10] This estimate of the conditional kurtosis differs significantly from the normal value of three, as seen by the $LR_{1/\nu=0}$ test for the GARCH (1, 1) model with conditional normal errors with χ_1^2 distribution (see Table 4). The estimated value of ν^{-1} are the inverses of the degrees of freedom parameter (see Table 2).

In conclusion, as expected, GARCH models worked very well for my purposes and this model is used to test the EMS currency volatility.

B. *Tests for EMS Currency Volatility*

A) *Test Method*

Because EMS implemented the Exchange Rate Mechanism (ERM)[11] in March 13, 1979, to test the volatility, I will designate the time period before March 13, 1979 as pre-EMS and after March 13, 1979 as post-EMS and see whether there is a difference in volatility in both periods.

To test volatility we simply could test the following null hypothesis:

H_0: Pre-EMS $\hat{\omega}_i$, $\hat{\alpha}_{1i}$, $\hat{\beta}_{1i}$ are same as those of post-EMS in our equation (8)

and (iii) $\gamma_x(r, s) = \gamma_x(r + t, s + t)$ for all $r, s, t \in Z$, where $\gamma_x(r, s) = \mathrm{Cov}(X_r, X_s)$. If $\alpha + \beta \geq 1$, then it blows up and we have an explosive ARMA model (See Bollerslev 1989 for discussions about IGARCH.).

[9]From Kendall and Stuart (1969), the fourth moment is equal to

$$E(\varepsilon_t^4 \mid \psi_{t-1}) = 3(\hat{\nu} - 2)(\hat{\nu} - 4)^{-1} h_{t-1}^2, \ \nu > 4.$$

[10]Baillie and Bollerslev (1987, 1989) have found that with weekly data the assumption of normality is generally appropriate.

[11]As of September 1988, Belgium, Luxembourg, Denmark, France, Germany, Ireland, Italy, and the Netherlands participate in the Exchange Rate Mechanism (ERM). Great Britain joined the ERM on October 1990, Spain on June 1989, and Portugal on April 1992. Greece are not in the ERM, but in the EMS. Hereafter, the term EMS is used to indicate these countries on their currencies in the ERM.

$$h_{ti} = \omega_i + \alpha_{1i} \varepsilon^2_{t-1} + \beta_{1i} h_{t-1} \quad (i = 1, 2). \tag{8}$$

However, if H_0 is rejected, does this imply increasing volatility, decreasing volatility, or neither? We can find no distinction between them. One possible way to structure our test would be to test $\hat{\omega}_i$, $\hat{\alpha}_{1i}$, and $\hat{\beta}_{1i}$, but, $\hat{\alpha}$'s and $\hat{\beta}$'s have really nothing to do with volatility levels. Therefore, $\hat{\omega}_i$'s will be used to test the change in volatility, i.e., the differences in $\hat{\omega}_i$'s.

$$y_t = b_0 + \theta(L)\varepsilon_t \tag{9}$$

$$\theta(L)u_t = \varepsilon_t + \theta_1 \varepsilon_{t-1} + \theta_2 \varepsilon_{t-2} \tag{10}$$

$$\varepsilon_t \mid \Omega_{t-1} \sim D(0, h_t) \tag{11}$$

$$h_t = \omega_0 + \omega_1 D_t + \alpha_1 \varepsilon^2_{t-1} + \beta_1 h_{t-1} \tag{8'}$$

$$\text{where} \quad D_t \begin{cases} = 1 \text{ if post-EMS} \\ = 0 \text{ if pre-EMS} \end{cases}$$

$$H_0: \omega_1 = 0$$

B) Test Results of EMS Currency Volatility

First, the nominal exchange rate volatilities were estimated against intra-EMS (D-mark against Lira, D-mark against French franc, and D-mark against the Netherlands guilder). The existence of the EMS since 1979 has coincided with a marked reduction in the volatility of exchange rates within the EMS. This was one major goal of the system, and to this end, the intervention arrangement and other elements of the exchange rate mechanism were established (see Table 2, columns 1 to 3). However, in terms of the nominal volatility against the US dollar, the EMS currency volatility increased during the EMS period. It is statistically significant at the 5% level for the cases which I have studied with the D-mark, Danish krone, the Netherlands guilder and Belgian franc against the US dollar (see Table 2). To compare the volatility level change between EMS currencies with that between non-EMS currencies, I have estimated the volatility of the Canadian dollar, Swiss franc, and Pound sterling against the US dollar, which showed an increase in nominal volatility during the EMS period in each case (see Table 2, last three columns).

In the previous case we checked the volatility level changes against two key currencies, the US dollar and the D-mark, which are both major reserve currencies and transaction currencies. In addition, we used the D-mark, because Germany is the leading country in the EMS. However, due to those factors, the measure of the exchange rate volatility might be distorted. To eliminate this problem we used the Pound sterling instead of the US dollar and the Italian lira instead of the D-mark as base currencies and tested the significance of the change in the level of the volatility again. The results are shown in Table 5.

As expected, in the case of intra-EMS currencies (the French franc and the Netherlands guilder against the Italian lira), there were significant decreases in the volatility after March, 1979 (see the first two columns in Table 5). But in the case of the EMS currencies, (Italian lira and Netherlands guilder) against the non-EMS currencies (Pound sterling), the Netherlands guilder, which showed an increase in volatility after March, 1979 when it is measured against the US dollar, showed at least constant volatility after March, 1979 (see Table 5, fourth column). Also, between non-EMS currency volatility, the Swiss franc, which showed an increase in volatility against the US dollar, accepts the null hypothesis that there was no change in the volatility even after March, 1979 (see Table 5 column 5).

As we suspected that during the sample period United Kingdom might try to stabilize her currency volatility against the other EMS currencies, as they are her neighbors, we tested the non-EMS currency volatility against the Japanese yen and Canadian dollar as base currencies. Table 6 shows that the Yen/guilder and Yen/Sfr had at least constant volatilities again, and we can confirm our results.

The clear diminution of exchange-rate volatility in the case of intra-EMS is certainly consistent with the view that the system has been successful in contributing to exchange-rate stability among participating countries. However, as is shown in Table 2, in the exchange rate volatilities against the US dollar the volatility of the EMS currencies showed different patterns under their exchange-rate mechanisms from those of the non-EMS currencies. Hence, we can say that decreasing volatility of the intra-EMS does closely follow the increasing volatility against the US dollar. This was already noted by Cohen (1981), who said that "...effort to maintain the joint float could increase the volatility of fluctuations between participating and non-participating currencies..." (see p. 14). It appears that such effort may do so at the cost of increased instabil-

TABLE 5

ESTIMATION OF GARCH MODELS WITH THE ITALIAN LIRA AND THE POUND STERLING AS BASE CURRENCY

$y_t = b_0 + u_t$; $u_t = \varepsilon_t + \theta_1 \varepsilon_{t-1} + \theta_2 \varepsilon_{t-2}$; $\varepsilon_t \mid \Omega_{t-1} \sim D(0, h_t, v)$;
$h_t = \omega_0 + \omega_1 D_t + \alpha_1 \varepsilon^2_{t-1} + \beta_1 h_{t-1}$

Parameters & Diagnostics Statistics	EMS				Non-EMS	
	Inter-EMS		EMS vs Non-EMS			
	FFR-LIRA	NGL-LIRA	UK£-LIRA	UK£-NGL	UK£-SFR	UK£-CN$
Log L	−2773.098	−2775.547	−3173.470	−3162.507	−3331.852	−3258.569
b_0	0.161	0.934**	0.716*	−0.785*	−1.214**	−0.134
	(0.220)	(0.178)	(0.389)	(0.449)	(0.509)	(0.395)
θ_1	—	—	0.002	0.092**	0.044	0.0
			(0.039)	(0.039)	(0.038)	(0.037)
θ_2	—	—	0.006	0.019	0.033	0.067*
			(0.037)	(0.036)	(0.039)	(0.036)
ω_0	86.453**	84.280**	9.160**	9.818**	33.748**	8.985**
	(3.368)	(3.945)	(1.423)	(2.642)	(6.909)	(3.819)
ω_1	−64.247**	−63.423**	4.513**	−1.009	−0.681	21.505**
	(3.426)	(3.704)	(1.472)	(1.042)	(3.845)	(9.042)

TABLE 5
(CONTINUED)

Parameters & Diagnostics Statistics	EMS				Non-EMS	
	Inter-EMS		EMS vs Non-EMS			
	FFR-LIRA	NGL-LIRA	UK£-LIRA	UK£-NGL	UK£-SFR	UK£-CN$
α_1	0.392** (0.036)	0.532** (0.056)	0.145** (0.018)	0.081** (0.018)	0.201** (0.037)	0.152** (0.064)
β_1	—	—	0.781** (0.024)	0.850** (0.034)	0.646** (0.058)	0.755** (0.064)
ν^1	normal	normal	normal	normal	normal	0.185** (0.001)
m_3	−1.584	7.852	0.438	−0.592	−0.638	−0.936
m_4	26.800	85.781	7.414	6.129	4.903	11.186
$Q(10)$	11.792	8.696	13.275	9.754	7.028	28.704
$Q^2(10)$	3.554	8.499	15.476	2.312	2.462	1.348
$3(\hat{\nu}-2)/(\hat{\nu}-4)$	N.A.	N.A.	N.A.	N.A.	N.A.	7.271

Note: 1. Asymptotic standard errors are in parentheses under corresponding parameter estimates.
2. * indicates significance at 5% level and ** at 1% level.

ity of exchange rates against the US dollar.[12]

Even though there was a significant reduction in volatility after joining the EMS and although the study as a whole suggest fairly distinct patterns to the results, no strong conclusions as to cause or effect can be drawn. For example, it is impossible to say how far the reduced volatility among EMS currencies is due to the operations of the EMS itself. In addition, even if the coincident fall in volatility among EMS versus non-EMS currencies and the constant volatility among the non-EMS currencies are not a reliable indication of the way in which the EMS rates would have behaved in the absence of the system, it does nevertheless somewhat weaken the claim that the reduction in the volatility of intra-EMS rates is due to the creation of the system alone.

IV. The Multivariate Generalized ARCH Approach

In previous sections we estimated the univariate GARCH models, and they offered good statistical descriptions of exchange rate movements. However, they are not satisfying compared to a multivariate model because the multivariate approach gives some advantages for the following reasons:

(1) Nonzero covariances among exchange rate innovations require joint estimation of sets of regression if efficient estimation in parameters is to be achieved.
(2) Exchange rates are bilateral rates, and if new information comes to the foreign exchange market (e.g., the US money supply, the US budget deficit, the German trade surplus, etc.), it should affect all rates as market dealers change their demands of specific currency and it affects their portfolios.

The multivariate ARCH (q) model was originally introduced in Engle and Kraft (1982), and then used by Diebold and Nerlove (1986). Later it was generalized by Bollerslev, Engle and Wooldridge (1988). Baillie and Bollerslev (1990) modelled risk premia in forward exchange-rate markets with a multivariate GARCH approach.

[12]Also Marston (1980) says the volatility of the dollar exchange rate of that member country disturbs economic relationships between the two members of the union by changing cross-exchange rates between member currencies.

A. Estimation of Multivariate GARCH Models

In order to deal with the multivariate GARCH (1, 1) model, the following SUR system is estimated for the set of N currencies:

$$y_t = b_0 + \varepsilon_t \tag{12}$$

$$\varepsilon_t \mid \Omega_{t-1} \sim N(0, H_t) \tag{13}$$

$$\begin{aligned}\mathrm{Vech}(H_t) = C_0 + C_1 D_{1t} + A_1 \mathrm{Vech}(\varepsilon_{t-1}\varepsilon'_{t-1}) \\ + B_1 \mathrm{Vech}(H_{t-1})\end{aligned} \tag{14}$$

where y_t is vector of first-differentiated N currencies and b_0 and ε_t are $N \times 1$ vector of constants and innovation vector. The Vech($\cdot$) denotes the column-stacking operator of the lower portion of a symmetric matrix. C_0 and C_1 are $N(N+1)/2 \times 1$ vector, and A_1 and B_1 are $N(N+1)/2 \times N(N+1)/2$ matrices.

The conditional log likelihood function for (12)-(14) for the single time period t can be expressed as

$$\begin{aligned}L_t(\theta) = -N/2 \log 2\pi - 1/2 \log | H_t(\theta) | \\ - 1/2\ \varepsilon_t(\theta)' H_t^{-1}(\theta)\varepsilon_t(\theta),\end{aligned}$$

where all the parameters have been combined into $\theta' = (b'_0, C'_0, C'_1, \mathrm{Vec}(A_1)', \mathrm{Vec}(B_1)')$. Thus, conditional on the initial values, the log likelihood function for the sample $1, 2, \ldots, T$ is given by

$$L(\theta) = \sum_{t=1}^{T} L_t(\theta).$$

As is obvious from univariate case, the log likelihood function $L(\theta)$ depends on the parameter in a nonlinear form, and the maximization of $L(\theta)$ requires iterative methods.

While the multivariate GARCH (1, 1) of manageable size is considered here, a natural simplification is to assume that each covariance depends only on its own past values. We restrict our attention to two currencies, the Italian lira and the Nethelands guilder, first against the D-mark and then against other cross currencies. Weekly data from the FRB tape are used, as in the univariate GARCH models.

The model considered here becomes the bivariate GARCH model

$$y_{it} = b_i + \varepsilon_{it} \tag{12)'$$

$$\varepsilon_t \mid \Omega_{t-1} \sim N(0, H_t) \tag{13)'$$

$$h_{ijt} = C_{0ij} + C_{1ij} D_{1t} + \alpha_{ij} \varepsilon_{it-1} \varepsilon_{jt-1} + \beta_{ij} h_{ijt-1} \quad (i, j = 1, 2) \tag{14'}$$

where subscript i refers to the ith elements of the corresponding vector and ij to the ijth element of the corresponding matrix.

B. Model Estimates

The maximum likelihood estimates of the model obtained by the Berndt *et al.* (1974) algorithm for the case of the Netherlands guilder (y_{1t}) and the Italian lira (y_{2t}) against the D-mark are:[13]

$$\begin{bmatrix} y_{1t} \\ y_{2t} \end{bmatrix} = \begin{bmatrix} -0.001 \\ (0.06) \\ 0.61^{**} \\ (0.22) \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix}$$

$$\begin{bmatrix} h_{11t} \\ h_{21t} \\ h_{22t} \end{bmatrix} = \begin{bmatrix} 0.31^{**} \\ (0.08) \\ 3.15^{**} \\ (1.17) \\ 16.49^{**} \\ (2.03) \end{bmatrix} + \begin{bmatrix} 0.05^{**} \ \varepsilon^2_{1t-1} \\ (0.01) \\ 0.104^{**} \ \varepsilon_{1t-1}\varepsilon_{2t-1} \\ (0.02) \\ 0.31^{**} \ \varepsilon^2_{2t-1} \\ (0.03) \end{bmatrix} + \begin{bmatrix} 0.93^{**} \ h_{11t-1} \\ (0.05) \\ 0.73^{**} \ h_{12t-1} \\ (0.05) \\ 0.61^{**} \ h_{22t-1} \\ (0.03) \end{bmatrix} + \begin{bmatrix} -0.29^{**} \\ (0.08) \\ -2.91^{**} \\ (1.13) \\ -11.59^{**} \\ (1.79) \end{bmatrix} D_{1t} \tag{15}$$

log likelihood function $= -3970.3668$.

In the case of the Netherlands guilder against the D-mark, in the conditional covariance equation of the h_{11t}, 0.31 is the intercept, $\alpha = 0.05$, and $\beta = 0.93$, with -0.29 as the intercept change after March 1979. The h_{21t} is for the conditional covariance of the Netherlands guilder against the Italian lira, and the significant value of the coefficient of D_{1t} (i.e., -2.91) shows the decreasing volatility after EMS as already verified in the univariate case (see the Table 5). The h_{22t} is for the conditional covariance of the Italian lira

[13]Hereafter, * indicates significance at the 5% level and ** at the 1% level.

against the D-mark, and the significant value of the coefficient D_{1t} (i.e., -11.59) also shows the decreasing volatility after joining the EMS.

The estimates for the model are appealing, and the estimate value for each coefficient is reasonable and highly significant, lending some support for the arguments that time series for exchange rates works well under the GARCH model and that the intra-EMS currencies show decreasing volatilities after participating in the EMS system. However, the likelihood ratios between the univariate GARCH (-2758.883 for DM-LIRA and -2017.937 for DM-Netherlands guilder) and the multivariate GARCH (-3970.3668) implies that the multivariate GARCH is more efficient than the univariate model. Compared with the univariate model, significant coefficients of α and β are achieved in the case of Guilder/Lira.

I estimated the Lira and the Netherlands guilder against the Japanese yen as one test of the change in volatility of the EMS against the non-EMS. In the univariate GARCH model, the decreasing volatility was shown at the 5% significant level in the case of the Yen-Lira and at least no change in volatility in the case of the Yen-Netherlands guilder (see Table 6). Even in the case of the EMS currency against the non-EMS currency, there was at least constant volatility after March 13, 1979 and demonstrated that the reduction in the volatility of EMS currency could result even in EMS vs. non-EMS cases. The following multivariate GARCH estimates assured such a claim. The volatility of the Yen-Lira (y_{1t}) is decreased after March 1979 with a 5% significant level and the volatility of the Yen-Guilder (y_{2t}) is shown to be at least constant. Here, again, the multivariate model becomes more efficient than univariate models when we consider their likelihood functions. Furthermore, the constant term in the case of Yen/Guilder (see Eq. (16)) shows significance at the 5% level, which was not significant in the univariate case.

$$\begin{bmatrix} y_{1t} \\ \\ y_{2t} \end{bmatrix} = \begin{bmatrix} 1.389^{**} \\ (0.403) \\ 0.604^{*} \\ (0.367) \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \\ \varepsilon_{2t} \end{bmatrix} \tag{16}$$

$$\begin{bmatrix} h_{11t} \\ \\ h_{12t} \\ \\ h_{22t} \end{bmatrix} = \begin{bmatrix} 30.63^{**} \\ (2.82) \\ 19.89^{**} \\ (2.65) \\ 21.41^{**} \\ (3.71) \end{bmatrix} + \begin{bmatrix} 0.205^{**}\ \varepsilon^2_{1t-1} \\ (0.02) \\ 0.196\quad \varepsilon_{1t-1}\varepsilon_{2t-1} \\ (0.02) \\ 0.186^{**}\ \varepsilon^2_{2t-1} \\ (0.02) \end{bmatrix} \tag{17}$$

$$+\begin{bmatrix} 0.67^{**} & h_{11t-1} \\ (0.02) & \\ 0.67^{**} & h_{12t-1} \\ (0.02) & \\ 0.701^{**} & h_{22t-1} \\ (0.025) & \end{bmatrix} + \begin{bmatrix} -5.65^{*} \\ (2.89) \\ 2.72 \\ (3.14) \\ -0.45 \\ (3.43) \end{bmatrix} D_{1t}$$

log likelihood function $= -5170.6051$

Lastly, the volatilities of the Canadian dollar and the Swiss franc against the Japanese yen were estimated to see whether among non-EMS currencies they have increased volatilities after March 1979. The following multivariate estimates show that, in the case of the Swiss franc against the Yen, there was a decreasing volatility after March 1979 at the 5% significant level. In the case of univariate estimates, it showed at least constant volatility (see Table 6). In Eqs. (18) and (19), y_{1t} denotes the first-differenced Canadian dollar/Yen, and y_{2t} denotes the first-differentiated Yen/Swiss franc.

$$\begin{bmatrix} y_{1t} \\ \\ y_{2t} \end{bmatrix} = \begin{bmatrix} -0.91^{*} \\ (0.47) \\ 0.61 \\ (0.41) \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \\ \varepsilon_{2t} \end{bmatrix} \tag{18}$$

$$\begin{bmatrix} h_{11t} \\ \\ h_{12t} \\ \\ h_{22t} \end{bmatrix} = \begin{bmatrix} 9.62^{**} \\ (1.89) \\ -8.05^{**} \\ (2.52) \\ 16.01^{**} \\ (3.97) \end{bmatrix} + \begin{bmatrix} 0.08^{**} & \varepsilon^2_{1t-1} \\ (0.01) & \\ 0.09^{**} & \varepsilon_{1t-1}\varepsilon_{2t-1} \\ (0.02) & \\ 0.16^{**} & \varepsilon^2_{2t-1} \\ (0.03) & \end{bmatrix} \tag{19}$$

$$+\begin{bmatrix} 0.86^{**} & h_{11t-1} \\ (0.01) & \\ 0.80^{**} & h_{12t-1} \\ (0.04) & \\ 0.77^{**} & h_{22t-1} \\ (0.03) & \end{bmatrix} + \begin{bmatrix} 5.30^{**} \\ (1.43) \\ 1.79 \\ (1.88) \\ -5.24^{*} \\ (2.84) \end{bmatrix} D_{1t}$$

log likelihood function $= -5844.3448$

As we have seen, the estimates of multivariate GARCH models are efficient relative to univariate GARCH estimates, and it is important to have simultaneous multivariate estimation for the reasons

TABLE 6
ESTIMATION OF GARCH MODELS WITH THE JAPANESE YEN AND THE CANADIAN DOLLAR AS BASE CURRENCY

$$y_t = b_0 + u_t;\ u_t = \varepsilon_t + \theta_1 \varepsilon_{t-1} + \theta_2 \varepsilon_{t-2};\ \varepsilon_t \mid \Omega_{t-1} \sim D(0, h_t, \nu);$$
$$h_t = \omega_0 + \omega_1 D_t + \alpha_1 \varepsilon^2_{t-1} + \beta_1 h_{t-1}$$

Parameters & Diagnostic Statistics	EMS vs Non-EMS			Non-EMS vs Non-EMS	
	YEN-LIRA	YEN-NGL	CN\$-NGL	CN\$-YEN	YEN-SFR
Log L	−3286.487	−3205.630	−3297.810	−3383.972	−3256.098
b_0	1.929**	0.211	−0.625	−0.931*	0.355
	(0.420)	(0.477)	(0.517)	(0.540)	(0.393)
θ_1	—	0.077*	0.090**	0.070*	—
		(0.036)	(0.037)	(0.034)	
θ_2	—	0.097**	0.137**	0.069*	—
		(0.035)	(0.036)	(0.038)	
ω_0	15.829**	9.516*	12.176**	1.826**	12.911**
	(4.261)	(4.548)	(5.528)	(0.458)	(5.138)
ω_1	−3.669*	3.170	11.721*	2.523**	−0.743
	(2.099)	(3.010)	(6.171)	(0.564)	(3.915)
α_1	0.101**	0.109**	0.148**	0.041**	0.195**
	(0.016)	(0.034)	(0.034)	(0.004)	(0.044)
β_1	0.824**	0.821**	0.761**	0.944**	0.755**
	(0.031)	(0.049)	(0.059)	(0.004)	(0.047)
ν^{-1}	normal	0.158**	0.110**	normal	0.158**
		(0.031)	(0.022)		(0.035)
m_3	1.434	0.574	−0.156	−0.965	0.243
m_4	12.459	6.101	4.607	11.057	5.110
$Q(10)$	8.312	6.345	8.860	9.945	9.892
$Q^2(10)$	6.840	4.507	8.553	2.123	6.303
$3(\hat{\nu}-2)/(\hat{\nu}-4)$	N.A.	5.576	4.178	N.A.	5.576

Note: 1. Asymptotic standard errors are in parentheses under corresponding parameter estimates.
2. *indicates significance at the 5% level and ** at the 1% level.

mentioned. However, the magnitude and sign of the coefficients which were used to test for the change in the volatility after joining in the EMS did not vary with the multivariate estimates.

V. Conclusion

The European Monetary System was established on March 13, 1979 and plans for the development of European Community set forth in the Delors Report (1989) envision a single European Currency managed by an independent European System Central Bank. The time is ripe to evaluate this scheme and consider its possible future contribution to European and world-wide monetary relations as well as to European integration.

I have empirically studied the after-EMS currency volatilities and demonstrated them with multivariate GARCH models as well as with univariate GARCH models. Although the intra-EMS showed stable volatility after March 1979, one can not say that these stable exchange rate volatilities result from the system itself, because we have found that even in some EMS vs. non-EMS cases, as well as among non-EMS country cases, there existed at least constant volatilities. Furthermore, decreasing volatility of intra-EMS closely follows the increasing volatility against the US dollar, and an effort to maintain the joint float increases the volatility of fluctuations between participating currencies and the US dollar. Proposals for policy coordination among the major industrial economies have been discussed in recent years. But if such proposals utilize the successful EMS-member coordination for stable exchange rates, they should be considered carefully, because our experience indicates it is not used without cost.

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