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CAN EXCHANGE RATE PREDICTABILITY BE ACHIEVED WITHOUT MONETARY CONVERGENCE?

Evidence from the EMS

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The evidence presented here suggests that the European Monetary System has indeed coincided with more predictable exchange rates (nominal and real) between France, Germany and Italy. But if increased monetary policy coordination is the main explanation, then it is surprising that the conditional variance of real interest differentials between these countries does not appear to have fallen (unless the disturbances are mostly real, in which case fixed rates are suboptimal.) High onshore–offshore interest differentials for franc and lira assets, and the very slow convergence of intra-EMS inflation rates, suggest that capital controls have played a large role.

1. Introduction

When the European Monetary System was founded in March 1979, many skeptics argued that countries with widely divergent and highly variable inflation rates could not possibly hope to stabilize their bilateral nominal exchange rates. In one sense, the skeptics were right: the EMS experienced seven central parity realignments during its first five years. Nevertheless, recent studies by the International Monetary Fund and the European Community conclude that there has been less month-to-month volatility in intra-EMS exchange rates — both nominal and real — since the formation of the EMS.[1]

One purpose of the present paper is to investigate whether the EMS has coincided with reduced variability in *unanticipated* nominal and real exchange rate movements. Exchange rate expectations are measured using Euromarket forward rates, as well as random walk and vector autoregressive forecasting models. This study also extends earlier analyses by considering volatility at horizons greater than one month. (The efficient estimation of multi-month conditional forecast error variances turns out to require esti-

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[1]See Ungerer, Evans and Nyberg (1983), and the European Community Directorate-General for Economic and Financial Affairs (1984).

mation of a moving average process.) In order to examine the frequent assertion that the EMS provides a blueprint for reducing exchange rate variability among the world's largest currencies, the analysis focuses on France, Germany and Italy.[2]

Our findings based on conditional variances (variances of forecast errors) are broadly similar to the unconditional variance results reported in the IMF and EC studies. The variances of the French franc and Italian lira exchange rates against the Deutsche Mark (DM) were significantly lower during the first five years of the EMS than during the preceding five years; this result holds for nominal exchange rates at horizons from one month through one year. Real exchange rate volatility was also significantly lower at short horizons, though the evidence is not decisive at twelve-month horizons. In contrast, the DM exchange rates of the dollar, yen and pound were all more variable (though not always statistically significantly) during the EMS period than during the pre-EMS period. Bilateral cross-exchange rates for the dollar, yen and pound have also generally become more volatile since March 1979. Nor do results based on multilateral trade-weighted exchange rates reverse the conclusion that EMS members have experienced more forecastable exchange rates, at least at short-term horizons.

Whereas there is substantial evidence that the EMS period has been characterized by reduced intra-European exchange rate volatility, it is not certain how much this success should be attributed to increased monetary policy coordination. For one thing, only recently has there been any perceptible convergence between French, German and Italian inflation rates. Also, the evidence presented here suggests that the EMS has not led to a reduction in the *conditional variance* of real interest rate differentials between France and Germany. (The results are mixed for differentials between Italy and Germany.) This is troublesome because in the presence of financial disturbances, an optimal intervention policy directed at reducing unanticipated movements in real exchange rates should also end up reducing unanticipated movements in real interest rate differentials. It is true that the above result does not hold for real disturbances, but then standard intervention models do not prescribe fixing the exchange rate (à la EMS) in response to real disturbances. The relationship between the conditional variances of real interest differentials and real exchange rates also breaks down in the presence of capital controls. A comparison of onshore and offshore interest rates for the lira and the franc suggests that capital controls were operative both before and after the formation of the EMS. The presence of these controls may explain why Italy and especially France were able to reduce fluctuations in their exchange rates against the DM without reducing fluctuations in real interest differentials. The important role of capital

[2]In addition to France, Germany and Italy, the other participants in the EMS are Ireland, Netherlands, Belgium, Luxembourg and Denmark.

controls may limit the relevance of the EMS coordination experience for coordination between large countries with open capital markets, such as Germany and the United States.

Section 2 of the paper contains a brief description of the EMS, and evidence on the convergence of inflation rates. Section 3 examines nominal exchange rate variability, and section 4 looks at real exchange rate variability. Section 5 investigates the relative importance of capital controls and monetary policy coordination in stabilizing intra-EMS exchange rates. Efficient estimation of multi-step conditional forecast error variances is discussed in a statistical appendix.

2. Convergence of inflation rates

The EMS is something of a hybrid between a crawling peg system and a fixed but adjustable rate system. There are bilateral central rates for each pair of currencies, around which fluctuations of up to ± 2.25 percent are permitted (6 percent for the lira). Theoretically, when the margins are reached, the participating central banks are obliged to intervene in unlimited amounts. However, the central banks sometimes change the central rate instead. The bilateral intervention limits are supplemented by a 'divergence indicator', which may be viewed as measuring the deviation of a country's weighted average EMS-currency exchange rate from a weighted average of its bilateral central rates. When a currency reaches its 'threshold of divergence', set at 75 percent of the maximum possible divergence spread, there is a 'presumption' that the authorities in the deviating country will intervene and undertake changes in economic policy. Because divergence indicator movements do not strictly obligate intervention, and because the bilateral limits are often reached first, the divergence indicator does not play as large a prophylactic role as originally planned.[3]

The fluctuation margins around the bilateral central rates allow the system some scope to tolerate countries with different desired inflation rates. As a bilateral market rate approaches its intervention limit, the authorities can choose to move the bilateral central rate rather than intervene (though considerable consultation and negotiation are involved). By making a small enough adjustment in the central rate, the authorities can ensure that the lower tail of the new band overlaps with the upper tail of the old band. Thus, the realignment does not necessarily precipitate a movement in the market rate. Given the 12 percent band for lira bilateral rates, and 4.5 percent bands for the other rates, there can be significant intervals between realignments even in the face of persistent inflation rate differentials.

In practice, the EMS realignments are not typically as smooth as the

[3]Vaubel (1980) discusses some of the limitations of the divergence indicator.

process just described. Opposing the desire to defuse speculative capital movements is a wish to use fixed nominal exchange rates as a means of forcing the convergence of domestic inflation rates. The hope is that countries will undertake policy changes which would obviate the need for exchange rate adjustment. Such thinking underlies some episodes of substantial speculative pressures, such as March 1983, during which the authorities hoped to avoid or at least postpone the need for a realignment.[4]

The formation of the EMS did not produce a rapid convergence of inflation rates. Table 1 presents pre- and post-EMS inflation rates for seven countries. Comparing five-year averages, there is no evidence whatsoever of any convergence between France's, Germany's and Italy's inflation rates. (*GDP* deflators yield a similar picture to the *CPI* rates.) Indeed, any converging which took place was between the inflation rates of Germany, Japan and the United Kingdom. Comparing the final twelve months prior to the EMS, February 1978–February 1979, with the recent twelve-month period March 1983 through March 1984, we can detect some convergence in EMS inflation rates. Again, however, there is a much more discernable convergence between Germany's inflation rate and those of the U.S., U.K., and Japan. Even if French, German and Italian inflation rates do ultimately converge at a low level, one should be cautious in attributing this success to the existence of the EMS.

3. The variance of unanticipated nominal exchange rate movements, pre- and post-EMS

In most modern macromodels, unanticipated disturbances have far greater effects than perfectly anticipated shocks. Thus a natural measure of exchange rate volatility is conditional variance; that is, the variance of unanticipated movements in the exchange rate. Throughout most of this section, nominal exchange rate expectations will be measured using Euromarket forward rates. The implicit assumption is that expectations are rational and there is no exchange rate risk premium. The results are shown not to be sensitive to allowing for the type of (small) time-varying risk premium which has been detected in some exchange market 'efficiency' studies.

Table 2 contains estimates of the variance of the forecast error $e_t^{t+k} \equiv \log S_{t+k} - \log F_t^{t+k}$, where S_{t+k} is the exchange rate in month $t+k$, and F_t^{t+k} is the k-month ahead forward rate in period t. [Under the null hypothesis of risk neutrality, the mean prediction error is known and equal to zero; thus the root-mean-squared-error (*RMSE*) is equal to the standard deviation.] The data is monthly; the two subperiods considered are the five years preceding

[4]Collins (1983) employs the term structure of interest rates to analyze investors' expectations concerning the timing and the magnitude of the March 1983 realignment.

Table 1

A cross-country comparison of wholesale price index and consumer price index inflation rates before and after March 1979.[a]

	Feb. 74–Feb. 79 (mean 12-month rate)		March 79–March 84 (mean 12-month rate)		Feb. 78–Feb. 79		March 83–March 84	
	WPI	*CPI*	*WPI*	*CPI*	*WPI*	*CPI*	*WPI*	*CPI*
Germany	3.2	4.2	5.3	4.6	3.1	2.6	5.6	3.2
France	5.0	10.3	10.8	11.6	11.2	10.1	14.5	8.6
Italy	15.0	16.4	15.0	17.0	11.2	13.1	11.4	11.7
U.S.	8.4	7.9	6.7	8.0	10.3	9.9	2.9	4.7
Japan	2.2	7.8	4.6	4.3	−0.9	2.5	−1.3	2.5
U.K.	16.2	15.5	9.2	10.4	8.7	9.6	6.4	5.2
Switzerland	0.1	3.0	3.7	4.6	0.6	2.1	4.0	3.4

[a]See the data appendix for a description of the indices used. Note that the *CPI*/*WPI* differential for the early five-year subperiod is sensitive to the starting date. In some countries, the 1973 oil shock passed more quickly into *WPI*'s than into *CPI*'s.

Table 2

The variance of forward rate prediction errors, $\log S_{t+k} - \log F_t^{t+k}$.[a]

Forecast horizon (months)	Exchange rates versus DM				Multilateral trade-weighted[b] exchange rates			
	Feb. 1974–Feb. 1979		March 1979–March 1984		Feb. 1974–Feb. 1979		March 1979–March 1984	
	One month	Twelve months	One month	Twelve months	One month	Twelve months	One month	Twelve months
Germany					4.3	41 (15)	4.3	45 (17)
France	4.5[c]	72[c] (26)	1.3[c]	21[c] (8)	3.0	57 (23)	3.9	28 (8)
Italy	10.3[c]	86[c] (21)	2.2[c]	24[c] (8)	6.6[c]	107[c] (35)	2.5[c]	18[c] (6)
U.S.	8.9[d]	89 (29)	12.6[d]	161 (55)	5.5[c]	71 (21)	9.1[c]	90 (27)
U.K.	8.2	54 (14)	10.8	79 (21)	5.2[d]	43[d] (14)	7.8[d]	94[d] (29)
Japan	7.1[c]	84 (21)	14.1[c]	133 (45)	5.2[c]	69 (27)	10.8[c]	135 (58)
Switzerland	5.2	88[c] (26)	4.2	31[c] (13)	7.7	133 (46)	6.3	89 (27)

[a]The units are approximately percent-squared/forecast horizon. Twelve-month-horizon maximum likelihood estimates were obtained using the Multiple Time Series Arima package at the Federal Reserve Board. Standard deviations of the twelve-month variance estimates are in parentheses.

[b]See the data appendix for a description of the trade weights.

[c]Variance in pre-EMS subperiod is significantly different from variance in post-EMS subperiod at 5 percent level. A one-tailed F-test is used for one-month horizons; the test described in the statistical appendix is used for twelve-month horizon comparisons.

[d]Variance for pre-EMS subperiod is significantly different at 10 percent level.

the EMS and the first five years of the EMS.[5] Examining the one-month DM prediction errors, we find that the conditional variances of the French franc/DM and lira/DM exchange rates were both approximately a quarter as large during the EMS period as during the pre-EMS period. Using a one-tailed F-test, the hypothesis of equality of variances can be rejected at the 95 percent level for both rates. The Swiss franc/DM rate also has lower one-month conditional variance in the EMS period, though the difference is not statistically significant. The variances of the dollar/DM, pound/DM and yen/DM exchange rates are all higher under the EMS; the dollar/DM rate variance is significantly higher at the 90 percent level of confidence, and the yen/DM variance is higher at the 95 percent significance level.

The results reported in table 2 do not appear sensitive to the assumption of risk neutrality. One can allow for a time-varying risk premium in the one-month forward rates by regressing the time t prediction error, e_t^{t+1}, against e_{t-1}^{t} and e_{t-2}^{t-1}. Risk premia of this form have been detected in other studies.[6] The residual of the autoregression may be treated as the 'unanticipated' component of the forward rate forecast error. Relaxing the assumption of risk neutrality in this fashion yields one-month conditional variance estimates extremely similar to those reported in table 2. Another potential problem with table 2 is that variance estimates can be quite sensitive to outliers. This is especially worrisome if the distribution has fat tails and converges only slowly to normality. Indeed, a joint chi-square test of the excess skewness and excess kurtosis of the forward rate prediction errors indicates significant deviation from normality in over half the cases.[7] However, when mean absolute deviations rather than variances are used as a measure of variability, the comparisons across subperiods are qualitatively unaffected. The other variability comparisons presented below are also generally robust to using mean absolute deviations rather than variances. Yet another issue which should be addressed is recent evidence that forward rate prediction errors are conditionally heteroskedastic.[8] The variance estimates in table 2 are still meaningful in the presence of conditional heteroskedasticity, as they are still consistent estimates of the unconditional prediction error variances. In other words, if a regime were left in place for a long period of time and if the

[5]It is not really possible to draw a strict demarcation between the pre- and post-EMS regimes. The formation of the EMS was not entirely unanticipated (indeed, it was scheduled to go into effect in January 1979, but was delayed until mid-March). Anticipations of the new regime might well have affected exchange rate behavior in the old regime. Flood and Garber (1983) study the effects of anticipated 'process-switching' on exchange rates.

[6]Hansen and Holdrick (1980) present evidence that lagged forward rate prediction errors help predict future forward rate prediction errors. It has proven difficult to estimate a structural model of this risk premium; see Rogoff (1984).

[7]The test for normality is based on Jarque and Bera (1980) who demonstrate that the statistic $T/6\,\{(\mu_3^2/\mu_2^3)+[(\mu_4/\mu_2^2)-3]^2/4\}$ is distributed chi-square with two degrees of freedom, where μ_i is the ith sample moment about the mean and T is the sample size.

[8]See, for example, Cumby and Obstfeld (1984).

process generating the prediction errors variances were stationary, then the procedures followed here would yield a consistent estimate of the mean prediction error variance. The mean variance is an appropriate criterion for comparing regimes if the social loss function is quadratic in exchange rate prediction errors and is constant over time.

Because the EMS is subject to periodic realignments, one might conjecture that the EMS has not made exchange rates any easier to predict at longer horizons. This does not seem to be the case, however, at least for nominal exchange rates. The results for three-, six- (both not reported) and twelve-month forward rate prediction errors are qualitatively extremely similar to the one-month horizon results.[9] Note that for these longer horizons, the sample variance is a consistent but not an efficient estimator. When monthly data is employed, the overlapping multi-month horizon forecast errors follow a moving average process; see, for example, Hansen and Hodrick (1980). The estimates in table 2 are thus based on the maximum likelihood procedure discussed in the statistical appendix. (Qualitatively similar results obtain when sample variances or sample mean absolute values are employed.) The appendix also discusses the asymptotic test used to compare the pre- and post-EMS forecast error variances. (One cannot form an F-test based on the sample variances, because the overlapping forecast errors are serially correlated.) Even at twelve-month horizons, the franc/DM and lira/DM rates are significantly less variable in the post-EMS period.[10]

Though the focus of this study is on whether and how the EMS has succeeded in stabilizing intra-EMS rates, table 2 also reports results for multilateral trade-weighted exchange rates. The weights are based on each country's share of total trade among the G-11 countries; see the data appendix. (Clearly, one cannot derive a single ideal measure of the trade-weighted exchange rate without reference to a particular theoretical model. The measure employed here is a popular one, but it may be of interest to apply our methodology to alternative measures.) One reason for considering a trade-weighted exchange rate is that it is possible to construct examples where stabilizing the bilateral rate between two countries destabilizes the trade-weighted exchange rate of one or both countries. Canzoneri (1982) constructs an example based on real disturbances. Marston (1984) demonstrates that destabilization is possible even with purely financial disturbances. He shows that when intra-European bilateral rates are fixed, portfolio shifts

[9]The multi-horizon prediction errors are constructed so that there is no overlap between the last prediction error of the first subperiod, and the first prediction error of the second subperiod. Twelve-month forward rate prediction errors are based, for example, on forward rate data for February 1974 through February 1978, and for March 1979 through March 1983.

[10]For the three-month conditional forecast errors, both a test based on the maximum likelihood estimates and an F-test based on non-overlapping data were constructed. These tests similarly reject the hypothesis of equality of the pre-EMS and post-EMS conditional forecast error variances, for both the French franc/DM and lira/DM rates.

between the U.S. and Germany can have a greater effect on the trade-weighted exchange rates of third-party European countries than such shocks would have under a pure float. Since it is quite possible that private portfolio shifts between the U.S. and Germany are an important phenomenon, Marston's result raises the question of whether the EMS might have destabilized the trade-weighted lira and the trade-weighted franc. The evidence in table 2 suggests that this theoretically troublesome possibility did not strongly manifest itself, at least for nominal rates. (As we shall see below, the evidence is much less clear for real rates at longer forecast horizons.) Whereas the one-month forecast error variances rose significantly for the yen and the dollar (at the 95 percent level), and for the pound (at the 90 percent level); the conditional variances for the DM and French franc did not rise significantly, and the lira variance fell significantly (at the 95 percent level). The conditional variance of the multilateral trade-weighted lira at three and twelve months is also significantly lower during the post-EMS period. If bilateral trade weights had been used instead of multilateral trade weights, the reduction in the relative variance of the EMS country exchange rates would probably have been more pronounced, since most EMS trade is intra-European.

4. The variance of unanticipated real exchange rate movements, pre- and post-EMS

Some would argue that the EMS has served to destabilize real exchange rates, by delaying nominal exchange rate adjustments even where they are necessary to offset inflation rate differentials. Here we shall attempt to investigate this hypothesis. It is, unfortunately, even more difficult to measure expectations about the real exchange rate than the nominal exchange rate; one cannot directly observe inflationary expectations. (The problem is less severe at short horizons, since short-term exchange rate volatility is typically an order of magnitude greater than short-term price level volatility.)

Fig. 1 plots the logarithm of six bilateral real exchange rates against the DM; prices are measured using consumer price indices.[11] Note that between March 1979 and March 1984, the lira appreciated substantially in real terms against the DM; the change in the logarithm of the real exchange rate was 0.24. By contrast, the French franc appreciated against the DM by only 2 percent in real terms over the first five years of the EMS. Herein lies the difficulty of trying to use a relatively small data set to measure long-horizon conditional forecast error variances. Elaborate time series techniques cannot obscure the basic fact that it is very difficult to say whether private agents anticipated these trend movements.

[11]The *CPI* indices in figs. 1 and 2 have been seasonally adjusted, though the estimated seasonal adjustment factors are very small.

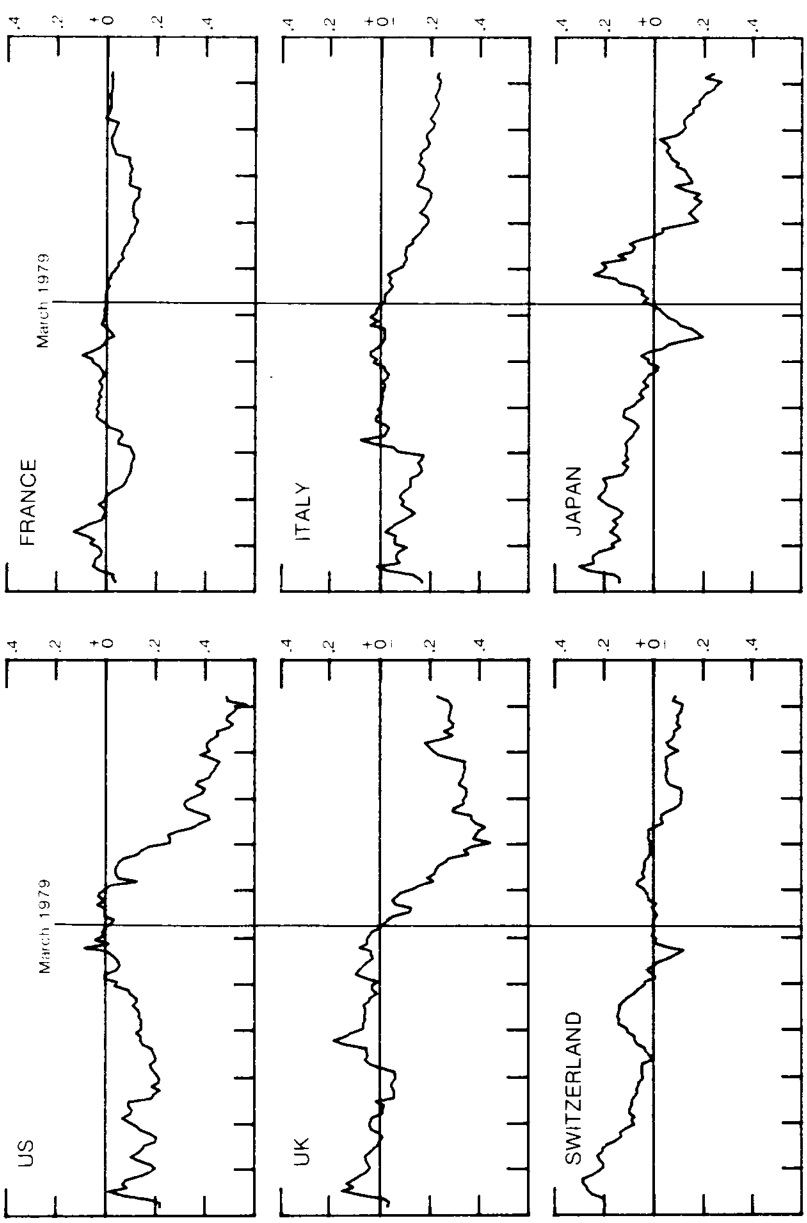

Fig. 1. Monthly movements in the log of real (*CPI*) exchange rates versus the DM, March 1973–March 1984. End-of-month foreign currency/DM rates, March 1979 = 0.

Having suitably qualified our results, we now turn to our two real exchange rate forecasting models. One is the random walk model, which predicts that the real exchange rate at any future date will be the same as today's real exchange rate. The out-of-sample forecasting results presented in Meese and Rogoff (1983) suggest that this naive forecasting model is difficult to significantly improve upon. The other forecasting model is a vector autoregression (*VAR*) which includes the logarithm of the real exchange rate, *CPI* inflation rate differentials, twelve-month Euromarket interest rate differentials, and the difference between the home and foreign cumulated trade balance (each measured in dollars and normalized by their respective mean absolute values). The *VAR*'s include contemporaneous values and two lags of each variable; monthly seasonal dummies are also included. To allow for a structural break at the point of formation of the EMS, separate *VAR*'s are estimated for the pre-EMS and post-EMS periods.[12] The estimated *VAR* forecast error variances are 'in-sample'; that is, the forecast errors are based on coefficients estimated over the entire subsample. Given Meese and Rogoff's results that most (nominal) exchange rate models fit poorly out-of-sample, one might argue that the in-sample *VAR* estimates provide a lower bound on the forecast errors, and the random walk model provides an upper bound. (The *VAR* forecasts might be improved though, by taking into account oil prices and/or structural breaks occurring at important elections.)

Table 3 lists the root-mean-squared-errors for the two forecasting models.[13] At one-month horizons, the *RMSE* for the *VAR* model are equal to or lower than those of the random walk model, but the differences are generally small. Using either measure, the results for one-month conditional real exchange rate forecasts are quite similar to those obtained for nominal exchange rates. Again, this is not surprising since price level movements are

[12] A likelihood ratio test (with a degrees of freedom correction) fails to reject the restriction that lag lengths are three instead of four in all twenty-six cases (across exchange rates and subsamples). The test rejects lag lengths of two in only four cases. (Three-month interbank interest rates from IFS were used in the lag length tests.) The short-horizon conditional variance of the real exchange rate generally changes little when a time trend is included. A likelihood ratio test rejects (at the 95 percent level) the hypothesis that all the coefficients in the two subsample *VAR*'s are equal for the French franc/DM, lira/DM, trade-weighted lira, and trade-weighted French franc. The hypothesis can only be rejected at the 80 percent level for the trade-weighted DM.

[13] The *VAR* forecasting begins two months into each subsample because two lags are required to generate a one-month ahead forecast. For the *VAR*, *RMSE* and standard deviations are the same at one-month horizons and very similar at longer horizons. *RMSE* and standard deviations are also very close at short-horizons for the random walk model. A second-order autoregressive model, with or without a time trend, yields in-sample one-month horizon *RMSE* which are slightly higher than those of the *VAR* model. Michel Galy informed me of unpublished work in which he has compared the pre- and post-EMS volatility of the French franc using shorter time periods and different price indices. He finds that the real franc may even have been more variable under the EMS.

Table 3
The root-mean-squared-error of unanticipated movements in the logarithm of the real exchange rate, using vector autoregression and random walk forecasts.[a]

	Feb. 1974–Feb.1979				March 1979–March 1984			
	VAR		Random walk		*VAR*		Random walk	
Forecast horizon (months)	One month	Twelve months	One month	Twelve months	One month	Twelve months	One month	Twelve months
Bilateral rates versus DM								
France	2.0	5.1	2.1[b]	8.9	1.3	3.0	1.4[b]	5.9
Italy	2.8	3.9	3.0[b]	7.9	1.3	3.7	1.4[b]	6.5
United States	2.7	5.6	3.1	8.0	3.2	11.7	3.4	15.2
United Kingdom	2.4	5.6	2.8	8.7	2.9	8.7	3.3	14.1
Japan	2.6	7.5	2.8[b]	8.8	3.7	9.3	3.7[b]	16.3
Switzerland	2.0	7.3	2.2[b]	10.0	1.6	3.7	1.7[b]	6.1
Multilateral trade-weighted exchange rates								
Germany	1.9	3.8	2.2	5.1	2.0	5.9	2.0	10.2
France	1.7	3.8	1.7	7.0	1.9	4.5	1.9	8.0
Italy	2.2	3.9	2.3[b]	7.1	1.2	2.1	1.4[b]	5.3
United States	2.1	5.8	2.3[c]	8.7	2.7	10.1	2.8[c]	11.6
United Kingdom	1.6	4.7	2.2[b]	6.8	2.5	6.8	2.9[b]	10.7
Japan	2.2	10.2	2.3[b]	10.2	3.1	5.6	3.2[b]	11.8
Switzerland	2.5	8.1	2.7[c]	10.0	2.1	4.9	2.2[c]	8.1

[a]Real exchange rates are measured using consumer price indices (see figs. 1 and 2). *RMSE* are approximately in percentage terms. See the data appendix for the multilateral trade weights. The vector autoregressions employ exchange rates, relative interest rates, relative inflation rates and cumulated trade balances.

[b]Under the assumption that real exchanges rates follow a random walk, a one-tailed F-test yields that the pre-EMS one-month conditional variance is significantly different from the post-EMS conditional variance at the 5 percent significance level.

[c]Under the assumption that real exchanges rates follow a random walk, a one-tailed F-test yields that the pre-EMS one-month conditional variance is significantly different from the post-EMS conditional variance at the 10 percent significance level.

relatively predictable at short horizons. Using a one-tailed F-test on the random walk variances (assuming unbiasedness), one can reject the hypothesis of equality of pre- and post-EMS variances for both the franc/DM and the lira/DM rates. The Swiss franc/DM rate variance also fell significantly. The one-month conditional variance of the real dollar/DM, pound/DM and yen/DM rates were all higher during the EMS subperiod, but only the yen rise is statistically significant.[14]

The trade-weighted real exchange rates are plotted in fig. 2, and their conditional forecast error *RMSE* are presented in table 3. Note that there has been no significant trend movement in the real trade-weighted lira. The one-month horizon random walk results indicate that the variance of the real trade-weighted lira was significantly lower during the EMS period and the real trade-weighted Swiss franc variance was also lower (though only at the 90 percent level). The real trade-weighted exchange rate variance of the DM and French franc did not change significantly, whereas the conditional variances of the trade-weighted real pound, yen and dollar (90 percent) all rose significantly in the second subperiod.

The point estimates for the twelve-month horizon *RMSE* present a somewhat similar picture for the DM bilateral rates. (No formal tests are presented.) Note though, that the fall in the volatility of the real lira/DM rate is much less decisive at twelve months, whether measured by the *VAR* or by the random walk model. The twelve-month horizon trade-weighted results are even more ambiguous. Using the random walk model, the real trade-weighted DM *RMSE* doubled during the post-EMS period; using the *VAR* model, the *RMSE* for the yen halved. Thus one cannot entirely dismiss the empirical relevance for the EMS of the Canzoneri–Marston point: Attempts to stabilize a bilateral rate may destabilize the trade-weighted rate.

One comparison that has not been made thus far is to ask what has happened to the variability of exchange rates between non-EMS currencies. Could it be that the post-EMS period is characterized by increased volatility between EMS and non-EMS currencies, but decreased volatility for both intra-EMS rates *and* intra-non-EMS rates? The answer is no, at least for the major bilateral exchange rates which are the focus of this study. The one-month forward rate prediction error variances for the pound/yen, yen/dollar and pound/dollar exchange rates all rose significantly in the post-EMS period. Using the random walk model for real rates, the real pound/yen,

[14]The F-tests reported in table 3 yield the same results when the random walk prediction error variances are calculated using the sample mean as when the random walk model is assumed to be an unbiased forecaster. An F-test is not valid for the *VAR* estimates, but it is possible to use an asymptotic likelihood ratio test; see Mood, Graybill and Boes (1974). The asymptotic test rejects equality of pre- and post-EMS variances for all the one-month *VAR* estimates.

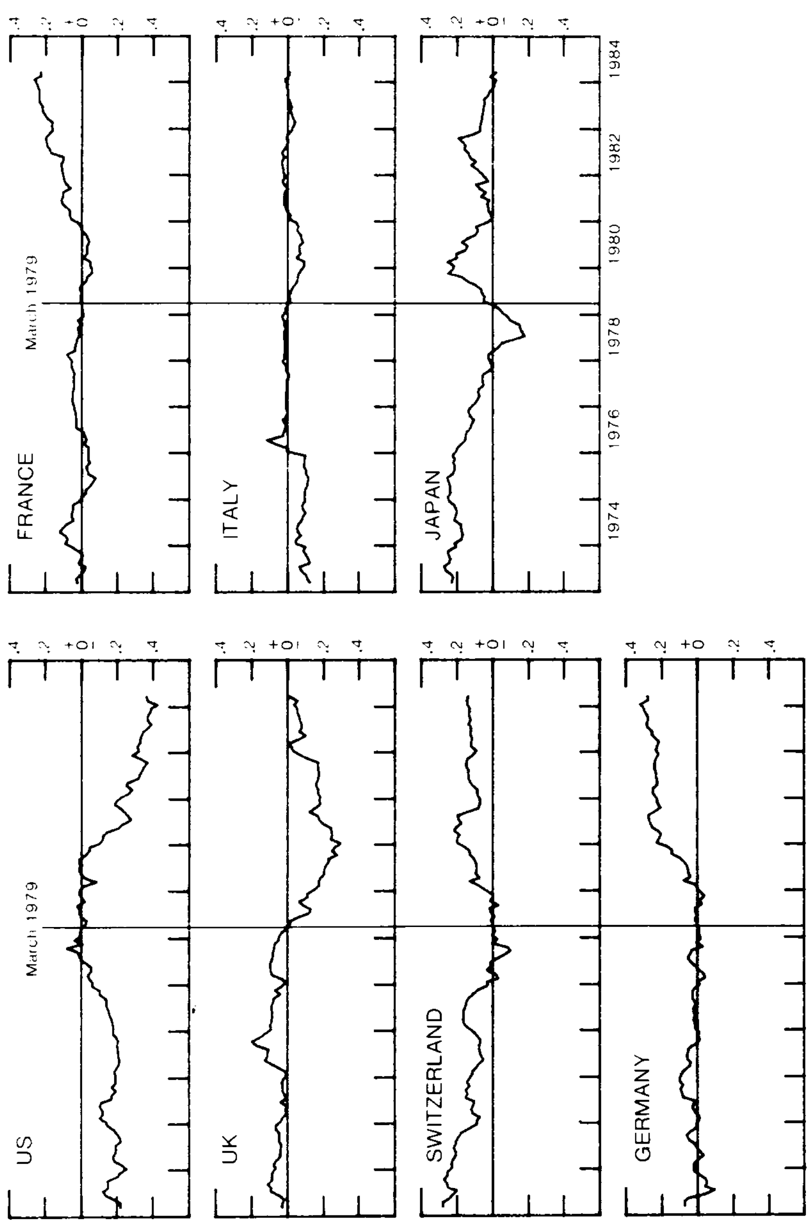

Fig. 2. Monthly movements in the log of trade-weighted real (CPI) exchange rates, March 1973–March 1984. End-of-month (home currency/foreign currency basket) rates, March 1979 = 0.

pound/dollar and yen/dollar one-month variances also rose significantly after March 1979; the pound/dollar variance fell insignificantly.[15]

5. Capital controls or monetary policy coordination?

It would appear that the EMS has indeed coincided with greater real exchange rate stability within Europe, at least in the sense that short-term movements in real exchange rates have become more predictable. How has this 'success' been achieved?

Italian and especially French capital controls have certainly played an important role in the EMS. Table 4 illustrates the significant differentials between Euromarket and domestic interest rates for franc and lira assets. The higher offshore rates probably cannot be attributed to a risk premium or taxes, since both these factors would presumably tend to go in the other direction. As table 4 illustrates, capital controls existed before the EMS, and indeed the analysis presented below does not rest on the assumption that the intensity of controls has varied over time. (Though French capital controls are generally thought to have become tighter after May 1981.[16]) The effects of capital controls on the volatility of any given exchange rate depend in part on the objectives of the monetary authorities.

Table 4
Offshore/onshore three-month interest differentials for France and Italy.

	Jan. 1975–Feb. 1979			March 1979–March 1984		
	Mean	Maximum	Standard deviation	Mean	Maximum	Standard deviation
France[a]	2.0	4.6	1.2	3.2	14.3	4.1
Italy[b]	5.0	24.8	5.5	3.8	18.5	4.6

[a]Three-month Eurofranc rate minus the three-month Paris interbank rate (end-of-month data, annual percentage rates).

[b]Three-month Eurolira rate minus the three-month Milan interbank rate (end-of-month data, annual percentage rates).

[15]For forward rate prediction errors, the pre-EMS one-month conditional variances for the (pound/yen, yen/dollar, pound/dollar) are (8.4, 8.8, 7.8). Post-EMS, they are (15.3, 15.5, 11.5). For real rates, the pre-EMS random walk mean-squared prediction errors are (8.4, 8.2, 8.0), and post-EMS they are (15.2, 13.7, 11.6). At twelve-month horizons, the differences are much less pronounced. I am grateful to John Flemming for suggesting that I present results for rates between non-EMS currencies.

[16]Frankel (1982) also analyzes the effects of French capital controls on offshore/onshore interest rate differentials. Some of the details of French and Italian capital controls are described in International Monetary Fund (1983).

If capital controls were insignificant, and if monetary policy coordination were solely responsible for reducing the conditional variance of real exchange rates under the EMS, one might well expect to observe a reduction in the conditional variances of intra-EMS real interest rate differentials. In the presence of money demand disturbances, for example, a money supply policy directed at stabilizing real exchange rates is equivalent to a policy of stabilizing real interest rate differentials. A similar result holds if a shift in private portfolio preferences between domestic- and foreign-currency denominated bonds is offset by a sterilized intervention operation.[17] It is true that the optimal (output stabilizing) response to real disturbances does not necessarily involve stabilizing real interest rate differentials, but then it does not typically involve stabilizing real exchange rates, either. Thus if one observes that intra-European real interest rate differentials have become more volatile since the formation of the EMS, it would suggest that either (a) capital controls have been a major factor in the stability of EMS exchange rates, or (b) the EMS has been (suboptimally) stabilizing real exchange rates in the face of real disturbances (though this normative conclusion is weakened if the real shocks are fiscal policy shocks). It is true that uncertainty about the size and timing of EMS realignments would be a third candidate explanation. But a supporter of the EMS would hardly want to argue that its main contribution was an increase in uncertainty about policy.

A direct relationship between the real exchange rate and the real interest rate differential may be obtained by manipulating the uncovered interest parity equation, which holds when capital mobility and asset substitutability are perfect,[18]

$$1+r_t(k)=[1+r_t^*(k)]({}_tS_{t+k}/S_t). \tag{1}$$

In eq. (1), $r_t(k)$ is the domestic k-period nominal interest rate at time t; $r_t^*(k)$ is the foreign rate. The exchange rate at time t is S_t, and ${}_tS_{t+k}$ is the expectation at time t of the exchange rate at time $t+k$. Denoting the domestic CPI by P and the foreign CPI by P^*, we can multiply both sides of eq. (1) by a common factor to obtain

$$\begin{aligned}&[1+r_t(k)](P_t/{}_tP_{t+k})/[1+r_t^*(k)](P_t^*/{}_tP_{t+k}^*)\\&\quad=(P_t/S_tP_t^*)({}_tP_{t+k}^*\cdot{}_tS_{t+k}^*/{}_tP_{t+k}).\end{aligned} \tag{2}$$

Taking logarithms of both sides of eq. (2) yields a relationship between the

[17]For a discussion of monetary stabilization policy in an open economy, see Henderson (1984).

[18]Isard (1983) stresses the usefulness of the 'identities' approach in analyzing the relationship between real exchange rates and real interest rates.

real interest differential and the expected rate of change of the logarithm of the real exchange rate,

$$ {}_tR_{t+k} - {}_tR^*_{t+k} = q_t - {}_tq_{t+k}, \tag{3} $$

where ${}_tR_{t+k} \equiv \log\{[1+r_t(k)](P_t/{}_tP_{t+k})\}$, and $q_t \equiv \log(P_t/S_tP^*_t)$. To derive a relationship in terms of conditional variances, take $t-1$ expectations across eq. (4) and subtract the resulting equation from (4) to obtain

$$ [{}_tR_{t+k} - {}_tR^*_{t+k}]' = q'_t - {}_tq'_{t+k}, \tag{4} $$

where ${}_tX'_t \equiv {}_tX_t - {}_{t-1}X_t$. Referring to eq. (4), consider the effect of a money demand disturbance. Because purely nominal disturbances have no real effects in the long run, the effect of the disturbance on the expected future real exchange rate should be smaller than its effect on the current real rate, at least for large enough k. [In standard sticky-price exchange-rate models such as Frankel (1979) or Mussa (1977), the effect on q_t would be greater than on ${}_tq_{t+k}$ for all k.] Thus if monetary policy is used to offset the effects of financial disturbances on current and expected future real exchange rates, one would expect to see a decline in the conditional variance of real interest rate differentials.

In table 5, short-term real interest rates are measured using two alternative proxies for expected inflation differentials: a three-month moving average of past inflation differentials, and the inflation forecasts of a vector autoregression. The *VAR* is of the same general form described in section 4, except that three-month domestic interbank rates are used in place of Euromarket rates. Note that by either expected inflation measure, short-term German/French and German/Italian real interest differentials have been higher on average under the EMS. The opposite is true for long-term differentials, which are constructed using a twelve-month moving average of past inflation as a proxy for expected inflation. What matters for comparison with our earlier results on the conditional variance of the real exchange rate, however, is the *conditional variance* of the real interest differential. For the short-term differential with lagged inflation proxy and for the long-term differential, second-order autoregressive (*AR*) processes are used to generate predictions of next month's real interest rate. (The results reported in table 5 are quite robust to inclusion of a time trend and to using first- or third-order *AR*'s.) For the short-term differential with *VAR* inflation proxy, the one-month-ahead real interest rate prediction errors are formed as follows: The *VAR* is used to generate time t expectations of (a) the three-month nominal interest differential in $t+1$, and (b) $t+1$ expectations of the inflation differential in periods $t+2$ through $t+4$. These are combined to form a time t forecast of the $t+1$ three-month real interest differential. The realized (ex-

Table 5

The month-to-month conditional variance of DM/franc and DM/lira real interest rate differentials.[a]

	Three-month real interest rate differential				Long-term real interest rate differential	
(Proxy for expected inflation):	(Three-month *MA* of lagged inflation)		(*VAR* model forecasts)		(Twelve-month *MA* of lagged inflation)	
	Mean	Standard deviation of one-month forecast error[c]	Mean	Standard deviation of one-month forecast error[c]	Mean	Standard deviation of one-month forecast error[b]
DM/franc						
Pre-EMS	1.3	1.4	1.2	0.8	2.8	0.45
Post-EMS	1.9	1.6	2.0	0.9	0.8	0.62
DM/lira						
Pre-EMS	1.2	2.2	1.2	1.6	6.3	0.79
Post-EMS	2.1	2.3	2.3	1.0	2.6	0.85

[a]Means and standard deviations are in percent/year. The data is monthly. Short-term nominal interest rates are end-of-month domestic three-month interbank rates; the series begin only in January 1975. The long rate series begin in February 1974.

[b]One-month conditional forecast error based on second-order autoregressive model.

[c]One-month conditional forecast error based on vector autoregressive model.

ante) real interest differential is then formed using the realized $t+1$ nominal interest differential together with time $t+1$ expectations of inflation in periods $t+2$ through $t+4$. The *VAR* method is consistent with rational expectations. Obviously, there is not enough data to apply the *VAR* approach to the long-term differentials.

The results are reported in table 5. For the German–French real interest differential, the conditional variance is higher in the EMS subperiod by any of our three measures. The evidence is mixed for the German–Italian differential, since the variance for the short-term differential with *VAR* inflation proxy is lower under the EMS. But the other two measures do not yield lower variances for the EMS period. Taken together, the results of tables 1, 4 and 5 suggest that the success of the EMS in stabilizing real exchange rates cannot be attributed to monetary policy coordination alone.

The issue of sterilized intervention has not been given much attention thus far. One reason is that to the extent it is used to offset portfolio disturbances, sterilized intervention should also stabilize real interest differentials. But the main reason is the growing body of empirical evidence that sterilized intervention has very little effect in the absence of capital controls.[19] This does not rule out the possibility that sterilized intervention can be 'effective' in an economy with capital controls. Consider a (not atypical) regime where foreigners may go to the home central bank and demand foreign currency in exchange for domestic currency, but domestic residents are not permitted to acquire foreign assets from abroad. In such an economy, sterilized intervention essentially mops up any domestic currency which leaks abroad and is not willingly held. Obviously, such intervention can continue only as long as the central bank has adequate reserves. Leakage may occur through illegal capital flight, or through current account deficits which are not financed by foreign-currency borrowing. If the controls are sufficiently effective to slow the pace of illegal capital flight, they can serve to (temporarily) protect the domestic currency from sudden portfolio shifts, thereby stabilizing day-to-day movements in the exchange rate.

Though my presumption is that the microeconomic inefficiencies caused by capital controls outweight any possible macroeconomic benefits, I will not try to argue the case here. What is clear is that if capital controls have been substantially responsible for the success of the EMS in stabilizing exchange rates, then the EMS experience has only limited relevance as a model for coordination between large countries with open capital markets.

6. Conclusion

Rather than repeat the main findings of this study, which are already

[19]For a recent survey of the empirical literature on sterilized intervention, see Rogoff (1984).

summarized in the introduction, I shall conclude by stressing two major qualifications of my results. First, it should be recognized that without a complete structural model and without knowledge of the processes governing the exogenous disturbances, one cannot really be sure what effect the EMS has had on the predictability of exchange rates. The 1979 oil price increase is but one example of a disturbance which makes the results here more difficult to interpret. By drawing on the pre-EMS experience and by drawing on the experiences of non-EMS countries, one can provide only a very imperfect counterfactual. Second, the EMS has not been in existence nearly long enough for one to develop powerful statistical tests of its effects. Is the trend rise in the real exchange rate of the lira against the Deutsche Mark an equilibrium phenomenon, or an EMS-generated bubble which we would see burst in a large sample? This is certainly one issue best left to future research.

Data appendix

All financial market data are end-of-month, seasonally unadjusted observations. Spot and three-month forward exchange rates are from the Federal Reserve Board data base. Forward rates for other maturities and Eurocurrency interest rates are from Data Resources Inc. [Forward rates and realized spot rates are matched according to the procedure described in Riehl and Rodriguez (1977). Three- and twelve-month non-dollar Euro-interest rates were constructed using covered interest parity.] Except for trade balances, which are formed from lines 70 and 71.v of *International Financial Statistics*, all the remaining data described below are from the FRB data base.

Wholesale and consumer price indices (*WPI*, *CPI*) are: France: wholesale prices (industrial goods), *CPI*; Germany: general *WPI*, cost of living index; Italy: general *WPI*, *CPI*; Japan: wholesale prices (all commodities),*CPI* (all items); Switzerland: *WPI* (total), *CPI* (total); U.K.: producer price index (manufactured goods, home sales), retail price index; U.S.: producer price index (finished goods), *CPI* (all items).

Domestic three-month and long-term interest rates are France: Interbank rate (Paris), long-term public sector bond yield; Germany: Interbank loan rates, long-term public authority loan rate; Italy: Interbank rate (Milan), long-term government bond yield (source: OECD, *Main Economic Indicators*).

Multilateral trade weights are based on each country's share of total trade (measured by the sum of exports plus imports) of the G-11 countries over the period 1972 through 1976. The weights are France (0.132), Germany (0.21), Italy (0.091), Japan (0.137), Switzerland (0.036), U.K. (0.121), U.S. (0.273).

Statistical appendix

One can obtain consistent estimates of k-month forward rate prediction error variances by sampling every kth monthly observation. When $k>1$, however, more efficient estimates may be found by employing all the data. Under the more efficient approach, the multi-step prediction errors overlap and thus follow a moving average process. Here we demonstrate that the sample variance of an MA process, although still a consistent estimator of the unconditional variance of the series, is asymptotically less efficient than the maximum likelihood estimator (MLE). This section also develops a test for comparing overlapping multi-step forecast error variances across independent subsamples.

Suppose forward rates are for two-month horizons and the data is sampled monthly. Let e_t be the difference between the logarithm of the realized time $t+2$ spot rate, and the logarithm of the time t two-month forward rate. Under the joint assumptions of rational expectations and risk neutrality, e_t follows the MA process,

$$e_t = a_t - \theta a_{t-1}, \tag{A.1}$$

where the a's are serially uncorrelated; for convenience we will make the further assumption that $a_t \sim N(0, \sigma_a^2)$, though our results only require asymptotic normality. We are interested in forming an estimate of $\sigma_e^2 = (1+\theta^2)\sigma_a^2$. The sample variance of e, $\mathring{\sigma}_e^2$, is given by

$$\mathring{\sigma}_e^2 = \frac{1}{n} \sum_{i=1}^{n} e_t^2. \tag{A.2}$$

(Note that mean of e_t is known and equal to zero under the null hypothesis.) The consistency of $\mathring{\sigma}_e^2$ is easily demonstrated;[20] the variance of the asymptotic distribution of $\mathring{\sigma}_e^2$ is given by

$$\mathring{V} = = n^{-1} \operatorname{plim} [n^{-\frac{1}{2}} \textstyle\sum (e_t^2 - \sigma_e^2)]^2, \tag{A.3}$$

$$\mathring{V} = n^{-1} \operatorname{plim} \left\{ n^{-1} \left[\sum_t (a_t - \theta a_{t-1})^2 - \sigma_e^2 \right] \left[\sum_s (a_s - \theta a_{s-1})^2 - \sigma_e^2 \right] \right\}, \tag{A.4}$$

$$\begin{aligned} \mathring{V} = n^{-1} \operatorname{plim} \Bigg\{ & n^{-1} \sum_t [(a_t - \theta a_{t-1})^2 - \sigma_e^2]^2 \\ & + 2n^{-1} \sum_t [(a_t - \theta a_{t-1})^2 - \sigma_e^2][(a_{t-1} - \theta a_{t-2})^2 - \sigma_e^2] \Bigg\}. \end{aligned} \tag{A.5}$$

[20]The fact that $\mathring{\sigma}_e^2$ is a consistent estimate of σ_e^2 may be viewed as a special case of theorem 14 in Hannan (1970, p. 228). Hannan's theorem demonstrates very general conditions under which sample moments are consistent estimates of population moments.

[To obtain expression (A.5) from (A.4), note that crossproduct terms involving $s \neq t+1$, t, or $t-1$, are all zero.] Expression (A.5) can be evaluated to obtain

$$\overset{*}{V} = \frac{2\sigma_a^4}{n}[(1+\theta^2)^2 + 2\theta^2]. \tag{A.6}$$

To derive the *MLE* of σ_e^2, $\tilde{\sigma}_e^2$, and its asymptotic variance, $\tilde{V}$, we employ the likelihood function for the *MA* process defined in (A.1),

$$L(\theta, \sigma_a | e_n) = (\sigma_a^2)^{-n/2}[(1-\theta^{2n+2})/(1-\theta^2)]^{-\frac{1}{2}} \exp\left\{\frac{-1}{2\sigma_a^2}\sum_{t=0}^{n} a_t^2\right\} \tag{A.7}$$

[see Box and Jenkins (1976, p. 272)]. By maximizing $\log L$ with respect to θ and σ_a^2, one can obtain the non-linear equations in $\{e\}$ for the *ML* estimates of θ and σ_a^2. The asymptotic covariance matrix of $\tilde{\theta}$ and $\tilde{\sigma}_a^2$, $H(\tilde{\theta}, \tilde{\sigma}_a^2)$, is given by the inverse of the asymptotic information matrix,

$$H(\tilde{\theta}, \tilde{\sigma}_a^2) = n^{-1}\begin{bmatrix} 1-\theta^2 & 0 \\ 0 & 2\sigma_a^4 \end{bmatrix}. \tag{A.8}$$

[The term $1-\theta^2$ is given in Box and Jenkins (1976, p. 245). Using the likelihood function (A.7), it is straightforward to prove that the information matrix is diagonal and to derive the asymptotic variance of $\tilde{\sigma}_a^2$.]

By the invariance property of maximum likelihood estimators [see Mood, Graybill and Boes (1974)], the *MLE* of σ_e^2 is given by

$$\tilde{\sigma}_e^2 = (1+\tilde{\theta}^2)\tilde{\sigma}_a^2. \tag{A.9}$$

To derive the asymptotic distribution of $\tilde{\sigma}_e^2$, we make use of (a) the fact that $[(\tilde{\theta}, \tilde{\sigma}_a^2) - (\theta, \sigma_a^2)] \overset{a}{\sim} N(0, H(\tilde{\theta}, \tilde{\sigma}_a^2))$, and (b) the Mann–Wald theorem [see Mann and Wald (1943)]. Together, (a) and (b) imply that $(\tilde{\sigma}_e^2 - \sigma_e^2) \overset{a}{\sim} N(0, \tilde{V})$, where

$$\tilde{V} = \frac{2\sigma_a^4}{n}[(1+\theta^2)^2 + 2\theta^2(1-\theta^2)] = rHr', \qquad r \equiv \left(\frac{\partial \sigma_e^2}{\partial \theta}, \frac{\partial \sigma_e^2}{\partial \sigma_a^2}\right). \tag{A.10}$$

(A consistent estimate of $\tilde{V}$ may be obtained using $\tilde{\sigma}_a^2$ and $\tilde{\theta}$.) Comparison of (A.10) and (A.6) reveals that for $(|\theta| < 1)$, $\tilde{V} < \overset{*}{V}$. Thus the *MLE* of σ_e^2 is more efficient than the sample variance. It is straightforward to extend to above analysis to the case of higher-order *MA* process.

We can make use of the asymptotic distribution of $\tilde{\sigma}_e^2$ to obtain an asymptotic test of whether, for example, the twelve-month forward rate prediction error variance was higher during the EMS period than during the pre-EMS period. By constructing the two subperiods so that the final forecast error of the first subperiod has no overlap with the first forecast error of the second subperiod, we can insure that the two ML estimates of the error variances, $\tilde{\sigma}_1$ and $\tilde{\sigma}_2$, are independently distributed with asymptotic variances $\tilde{V}_1$ and $\tilde{V}_2$. Therefore $[(\tilde{\sigma}_1-\tilde{\sigma}_2)-(\sigma_1-\sigma_2)] \overset{a}{\sim} N(0, \tilde{V}_1+\tilde{V}_2)$.

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