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Real Exchange Rate Volatility: Does the Nominal Exchange Rate Regime Matter?

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Abstract

A recent study by Grilli and Kaminsky (1991) argues that real exchange rate (RER) behavior is likely to be dependent on the particular historical period rather than on the nominal exchange rate arrangement itself. This paper reexamines RER behavior using alternative data sets, as well as different econometric methods, over the period 1880-1997. It finds strong evidence supporting the nonneutrality hypothesis of nominal exchange regime on RER volatility. Also, regime shifts play an important role in determining the persistence of shocks to the RER.

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Summary

One of the key issues in the international finance literature has been whether and how nominal exchange rate arrangements affect the behavior of various real macroeconomic variables. Many empirical studies of business cycles in open economies have found that real exchange rates (RER) have experienced dramatic changes in volatility across exchange rate regimes. A recent study by Grilli and Kaminsky (1991), however, challenges the validity of this empirical regularity. They argue that RER behavior is likely to be dependent on the particular historical period rather than the exchange rate arrangements themselves.

Using univariate time-series techniques, this paper seeks to contribute to the literature by reexamining RER behavior using alternative data sets, as well as different econometric methods, over the period 1880-1997. It finds strong evidence supporting the nonneutrality hypothesis of nominal exchange regime on RER volatility. Also, regime shifts play an important role in determining the persistence of shocks to the RER.

From a theoretical standpoint, the confirmation/rejection of regime-dependent behavior of the RER has important implications for the plausibility of various macroeconomic models of exchange rate determination. The findings here support the view that market imperfections may be the reason behind the regime-dependent phenomenon of RER volatility.

I. INTRODUCTION

Before the Bretton Woods fixed exchange rate commitment collapsed, proponents of floating exchange rates claimed that the real exchange rate should be more stable in a floating regime, since the flexibility in nominal rates would offset the effect of different national inflation rates on a country's international competitiveness (Friedman (1953), and Sohmen (1961)).

After major exchange rates were allowed to float in 1973, however, predictions about the desirability of a floating system underwent serious reexamination. One of the key stylized facts of the post-World War II floating experience has been the high volatility of both nominal and real exchange rates. To many analysts, the nominal exchange rates have been *excessively* volatile relative to the fundamentals (although this is still a contentious issue). But it is the higher volatility of RERs during the floating exchange period that is at the heart of the academic debate.

Many empirical studies of business cycles in open economies have found that the nominal and real exchange rates (RER) have experienced dramatic changes in volatility across exchange rate regimes.[1] This observed empirical regularity is often interpreted as evidence against theoretical models that exhibit nominal exchange rate regime neutrality, and in support of sticky prices. Grilli and Kaminsky (GK) (1991), however, challenge the validity of this basic tenet of the international finance literature. By examining the dollar-sterling RER series between 1885-86, they find that only when the post-World War II data are included could they find different volatility behavior across exchange rate regimes. Hence, they argue that real exchange rate behavior is likely to be dependent on the particular historical period rather than the nominal exchange rate arrangements *per se*.

This paper reexamines the volatility of real exchange rates using alternative data sets, as well as different econometric methods, over the period from 1880 to 1997. Contrary to Grilli and Kaminsky (1991), the findings of this paper strongly support the hypothesis of nonneutrality of nominal exchange rate regime on real exchange rate volatility. By employing more rigorous measurement of volatility, the paper finds that the shocks to the dollar-sterling real exchange rates, as well as to the franc-sterling rates, do not seem to come from the same distributions under the flexible and fixed regimes. This conclusion holds for the entire sample period of 1880-97 and for the pre-World War II period. For the post-World War II period, the real exchange rate volatilities of eight industrial countries are found to have increased after the Bretton Woods system broke down. Moreover, in the post-Bretton Woods period, this study finds that the real exchange rates among major European countries have behaved differently when they joined the exchange rate mechanisms (ERM) of the European Monetary System (EMS). Adopting the adjustable-peg system helped the EMS countries to reduce their RER volatilities.

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[1]See Mussa (1986), Baxter and Stockman (1989), Flood and Rose (1995), and Rogers (1995).

The findings of this study are of importance to international macroeconomists because they establish unambiguously the "stylized fact" that the nominal exchange rate regime plays a major role in determining real exchange rate behavior. The shifts in volatility of the real exchange rates can not be attributed to a particular historical period, but rather seem to be systematically linked to the specific nominal exchange rate arrangements in place. The evidence from the two long-run data series, as well as from the post-Bretton Woods period, casts serious doubt on the argument that high RER volatility during certain historical periods has arisen due to factors unrelated to the nominal exchange rate regime, such as the two major oil shocks in the 1970s. The reason (s) behind the nonneutrality phenomenon are still unresolved. Many recent empirical studies, however, have pointed to the sluggishness in the adjustment of national price levels. Besides possible sluggish adjustment in the prices of nontraded goods, economists are increasingly certain that there exist large deviations in the law of one price for traded goods, reflecting the underlying differences in market structures across industries.[2]

The paper is organized as follows. Section II motivates the study with a brief discussion on Grilli and Kaminsky's (1991) evidence regarding RER volatility. An overview of real exchange rate behavior over various historical periods is presented in Section III. In order to examine RER variability around its long-run trend (or average), the issue of alternative trend specification is addressed in Section IV. Section V presents the empirical findings on volatility changes with regime shifts using two long-run annual data sets, and the post-World War II monthly data of eight European countries. In Section VI, seven EMS members' RERs are analyzed by comparing their behaviors inside and outside the EMS. The last section concludes.

II. The Grilli and Kaminsky's (1991) Findings

Grilli and Kaminsky (1991) examine the monthly observations of the real exchange rate between the U.S. dollar and the British pound during 1885-86. Using the Wald-Wolfowitz (W-W (1940)) test, they conclude that the distribution of the monthly rate of change of the real exchange rate is the same under fixed and floating regimes when only the pre-World War II data is used. Therefore, the seemingly regime-dependent volatility behavior of the real exchange rate is only present in the post-World War II period.

Wald and Wolfowitz (1940) develop a nonparametric test designed to check whether two samples are from the same distribution. In the application to the RER study, the observations on the volatility from fixed and flexible exchange rate regimes are considered

[2]See Engel and Rogers (1995), Knetter (1989), and Goldberg and Knetter (1997).

independent observations from two samples. First, a sequence is constructed by merging these observations together and sorting them ascendingly. Then a complementary dummy sequence to the one above is created by assign the value of "0" to observations in the fixed exchange rate period and "1" otherwise. The dummy sequence will be alternating sets of 0's and 1's.

Each such set of 0's and 1's is defined as a run. Let *u* be the total number of runs, *m* be the total number of observations in the fixed exchange rate period, and *n* be the total number of observations in the flexible period. If the observations from each exchange rate period are independent, under the null hypothesis that they are from identical distributions, the mean and the variance of *u* are given by:

$$E(u) = \frac{2mn}{m+n} + 1 \tag{1}$$

$$\sigma^2(u) = \frac{2mn(2mn-m-n)}{(m+n)^2(m+n-1)} \tag{2}$$

Wald and Wolfowitz show that as *m* and *n* goes to infinity and $(m/n) \rightarrow \lambda > 0$, the distribution of $[u-E(u)]/\sigma(u)$ converges to N(0,1).

Two important issues arise when using the W-W test to study the real exchange rate volatility. The first issue is how to measure volatility. GK use the monthly rate of change. A more appropriate measurement, however, should be the deviation of the RER from its long-run trend (mean). Unless the real exchange rate follows a random walk, which GK's study strongly disputes, the monthly rate of change can be a seriously biased estimate of real exchange rate variability. The direction of the bias may go to either direction.

Secondly, without appropriate measurement of the mean and thus volatility of the RER, the assumption of independence of the observations may be significantly violated. By employing Monte Carlo simulations, GK show that the ability of the W-W test to discriminate between different distributions is not affected by the presence of a small degree of serial correlations. The true degree and direction of serial correlation in volatility, however, is unknown without an appropriate measurement of the mean. In particular, the use of the monthly rate of change by GK as a measurement for volatility seems to be at odds with their later claim that the real sterling-dollar exchange rate is not a random walk process. If the RER is indeed mean-reverting, the series of monthly rate of change of it does not only contain potentially large positive serial correlation, but also a unit root in the moving average part of the error process. How these complications are going to affect the power of the W-W test is unknown.

Even if we accept the appropriateness of the W-W test in this case, the test results in regard to the RER volatility across exchange regimes is not very supportive of the main conclusions reached by GK. For the whole sample period, as well as for the sample when the World War II and major devaluation periods are excluded, GK's finding strongly rejects the null hypothesis of no volatility shift across exchange rate regimes.[3] The null is marginally rejected again when the Bretton Woods period, as well as the WORLD WAR II and major devaluation periods, are all excluded. Only when using pre-World War II data alone are GK unable to reject the null. The rationale for using long-run data sets precisely lies in the ability to observe behavior differences across as many different exchange rate regimes as the data can provide. It is difficult to understand the benefits for carrying out the test by excluding the Bretton Woods period. In addition, excluding the post-World War II period leaves about half of the data observations out, with some exchange rate episodes having only a few years of observations. This may seriously reduce the power of the test.

One further limitation of the W-W test is that, even when the test is able to reject the null hypothesis that the two samples are from the same distribution, it is unable to tell which sample contains a higher second moment.

In summary, the results of the GK study on the real exchange rate volatility are not as conclusive as the authors suggest. In light of the limitations of the W-W test outlined above, this study prefers other econometric techniques that are more robust in detecting volatility shifts and their direction. Section V illustrates that, after the mean is appropriately specified, the residual series from different exchange regimes do not seem to come from the same distribution. This conclusion holds for the dollar-sterling RER even when only the pre-World War II data is used. The differences in conclusions between this study and the GK paper demonstrate the importance of an appropriate measurement of the mean when studying the variability around it.

Nevertheless, for the long-run United Kingdom-United States and France-United States RER series, the W-W test is also used later to compare the test results with those obtained by other more robust econometric techniques.

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[3]See Table 2 in Grilli and Kaminsky (1991).

III. A PRELIMINARY LOOK AT THE DATA

A. Description of Data Sets

This study examines two alternative data sets collected by (1) Lothian and Taylor (1996) and (2) the IMF, respectively. The real exchange rate series are defined as:

$$q_t = s_t + p_t^* - p_t \tag{3}$$

where s_t is the natural logarithm of national currency price of foreign exchange. p_t^* and p_t are the natural logarithm of foreign and domestic price levels, respectively.

Lothian and Taylor (1996) construct a data set consisting of almost two century data for the annual dollar/pound (1791-90) and franc/pound (1804-90) real exchange rates. The data series used in this study start from 1880, and are updated through 1997 by the author of this paper using the IMF's International Financial Statistics (IFS).

In the second data set, price and exchange rate series (Jan. 1957-Dec. 1997) are obtained from the IMF's IFS for eight European countries: Belgium, Denmark, France, Germany, Italy, Ireland, the United Kingdom and Netherlands. These data permit a detailed study of the short run volatility patterns during and after the Bretton Woods fixed exchange rate regimes, as well as inside and outside the EMS. If all long run, short run and the EMS period reveal that the behaviors of RERs are regime-dependent, the nonneutrality hypothesis of exchange rate arrangement can be more firmly established.

B. Historical Description of Exchange Rate Regimes

The data series in this study will be grouped under the two alternative nominal exchange rate regimes, fixed and floating. This section provides a brief delineation of the history of the international monetary system by exchange regime between 1880-97.[4]

Over the period studied here, there are three fixed exchange eras. The first is the classical gold standard from 1880 to 1913, a period characterized by fixed nominal exchange rates and essentially no capital controls. The second was the interwar period from 1927 to 1931 when principal countries of the world returned to the gold standard. Finally, there was the Bretton Woods system from 1946-71. During this period, the dollar was pegged to gold while the other countries pegged to the dollar.

[4]See Eichengreen (1994) and Bordo and Schwartz (1996) for a detailed discussion.

There are three flexible exchange rate episodes during the period studied here. The first episode, which was as close to a free float period as the history records, is between the First World War until the mid-1920s. The second floating exchange episode is 1932-38, although government intervention in the foreign exchange market was pervasive. In August 1971, the devaluation of the dollar marked the beginning of the end of the Bretton Woods system. In June 1972, British pound started to float against the dollar. Thus, 1972-96 is the final floating rate episode considered.

Table 1. Exchange Rate Regimes

1880-1913	1914-26	1927-31	1932-38	1946-71	1972-97
Fixed Exchange Regime	Flexible Exchange Regime	Fixed Exchange Regime	Flexible Exchange Regime	Fixed Exchange Regime	Flexible Exchange Regime

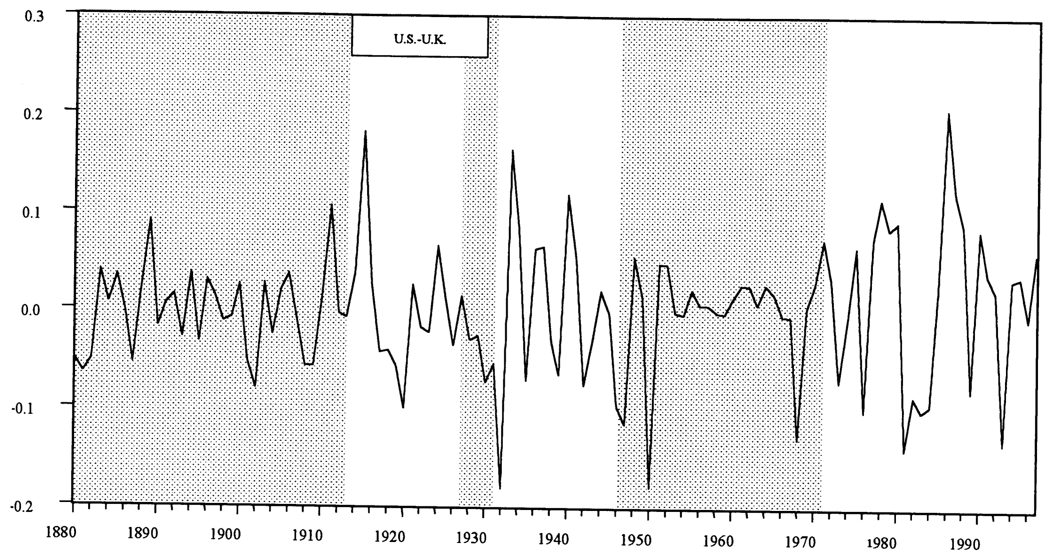
Using Eichengreen (1994) as a guide, the data series will be divided into the six subsamples shown in Table 1. Note that the World War II period between 1939 to 1945 is excluded. It was characterized by strict wartime controls of foreign exchange. Free trading in sterling had ceased by September 1940. It seemed best not to clarify this period as having either fixed or floating exchange rate regime.

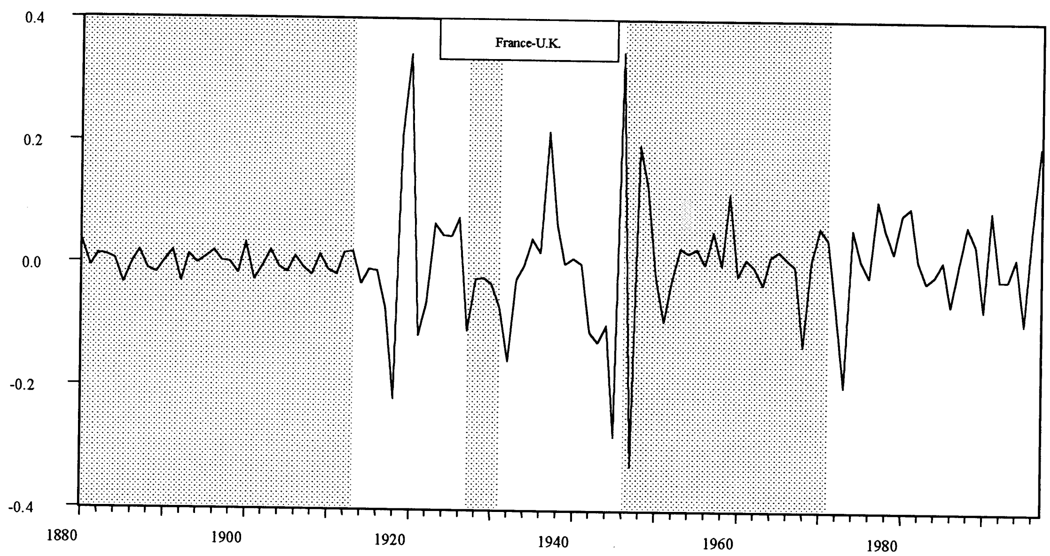
C. Some Visual Evidence of RER Volatility Cross Exchange Regimes

Period over period percentage changes for the various RER series are shown in Figure 1. Shaded areas correspond to the fixed exchange rate episodes. Simple visual examinations of the graphs reveal striking differences in the RER behavior across exchange rate regimes. The two long-run RER series spanning over one hundred years suggest that the rates were relatively tranquil during the fixed exchange episodes and became much more volatile during the flexible exchange periods. The post-World War II monthly RER series against the U.S. dollar show clear differences in volatility during and after the Bretton Woods fixed exchange rate system.

A simple statistical summary of the data series is presented in Table 2. For the dollar-sterling RER, the mean of the absolute yearly change during floating rate episodes was about twice that experienced during fixed exchange period. The same is true for the franc-sterling rate before the World War II. An interesting phenomenon, however, is that the mean and variance of the yearly rate of change was higher during the Bretton Woods period than those

Figure 1. Percentage Change of Real Exchange Rates over Previous Periods





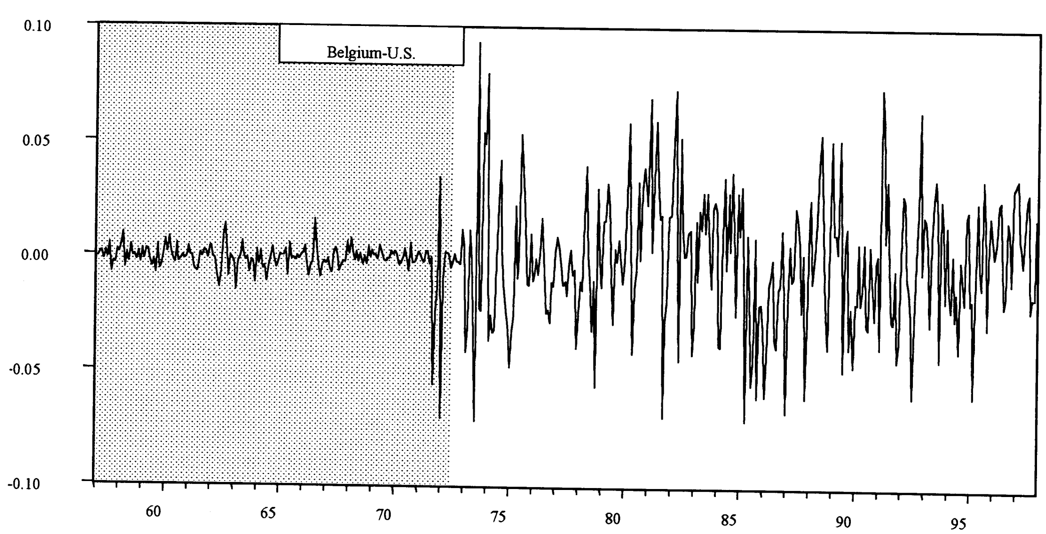
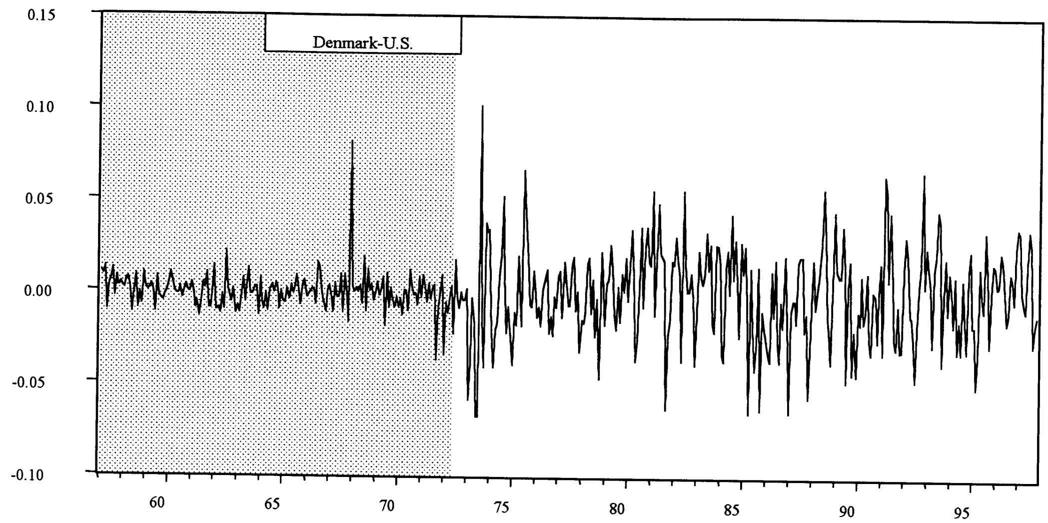
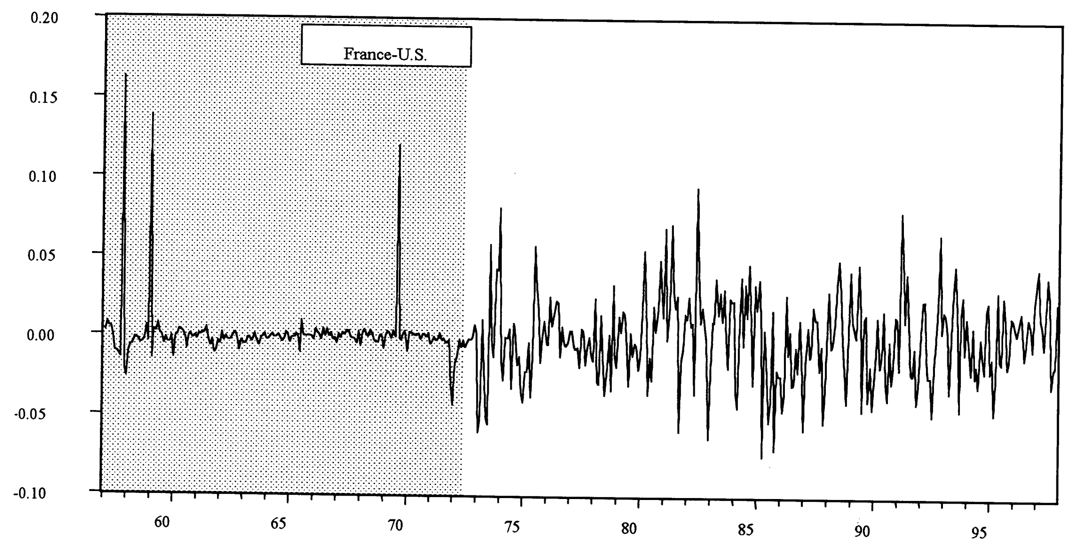


Figure 1 (Continued). Percentage Change of Real Exchange Rates over Previous Periods





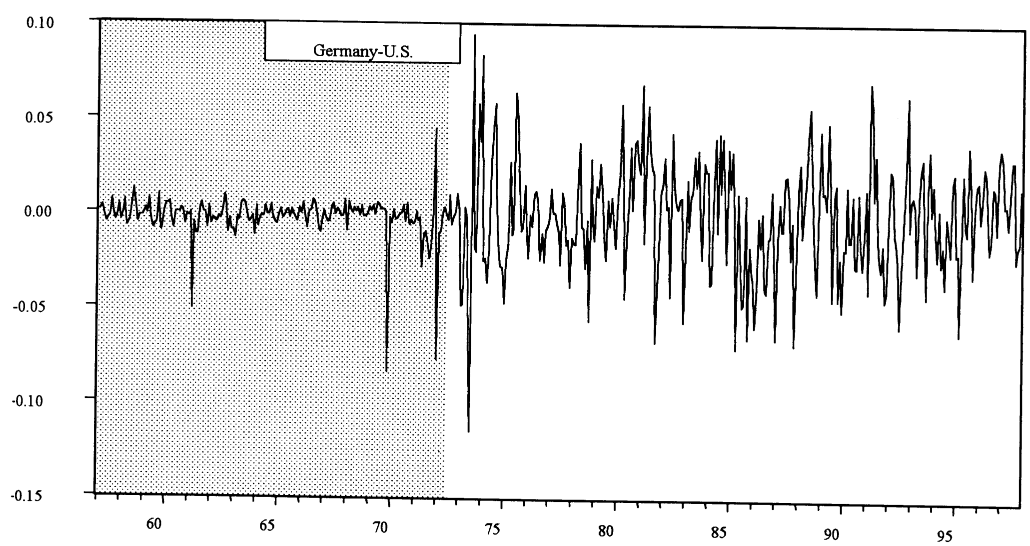
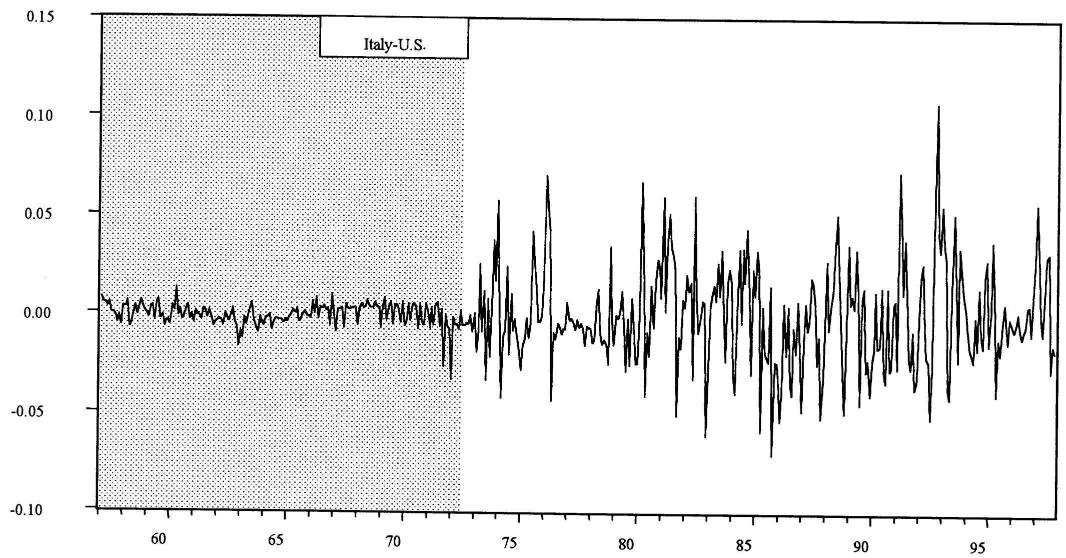


Figure 1 (Continued). Percentage Change of Real Exchange Rates over Previous Periods



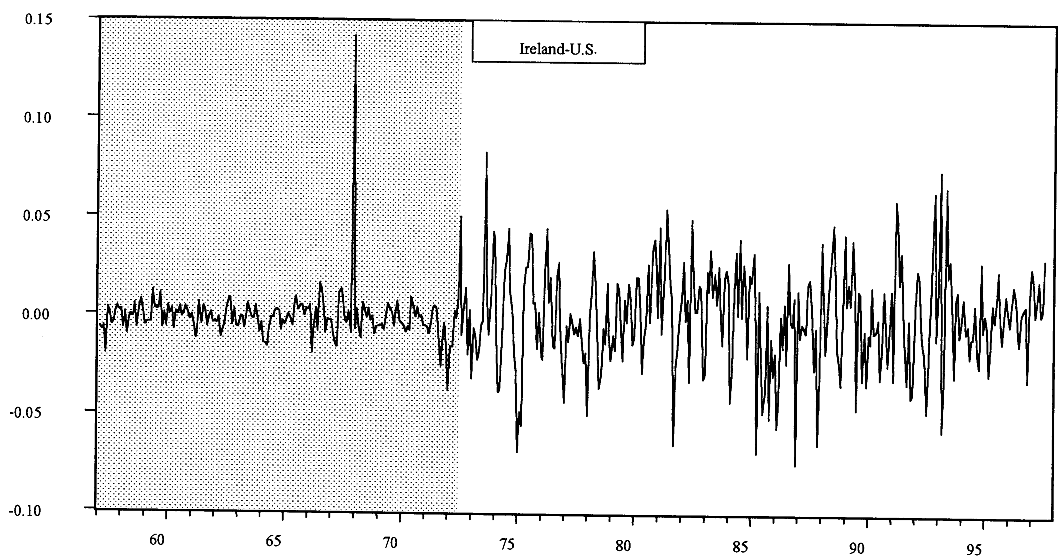
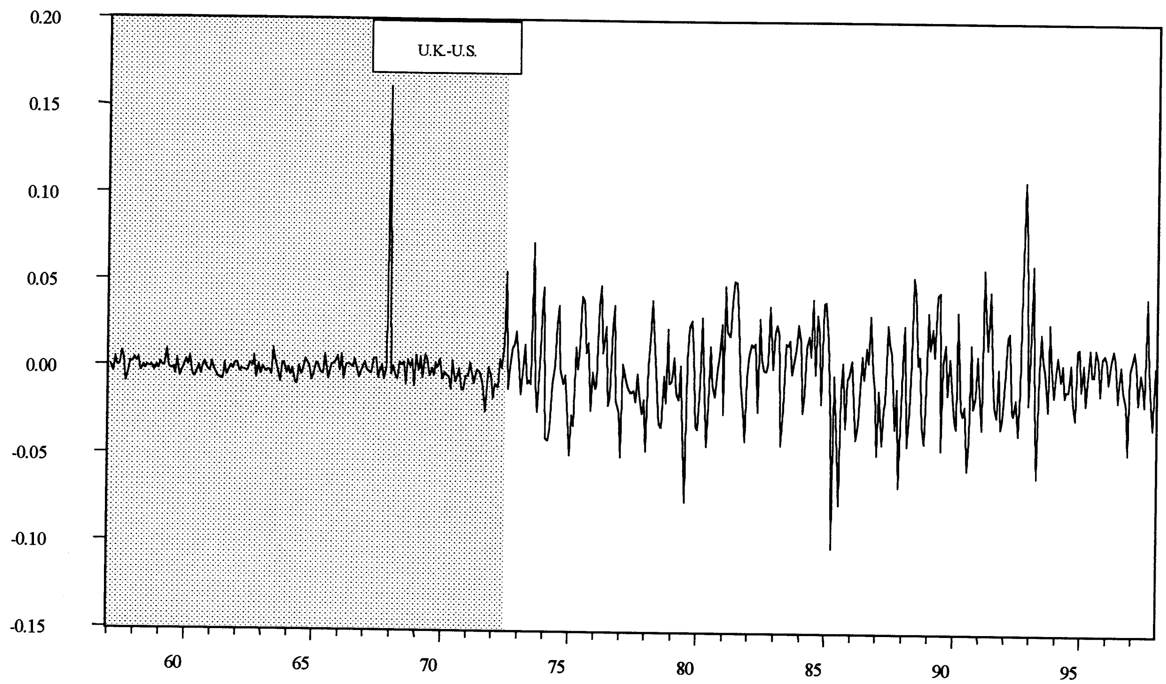
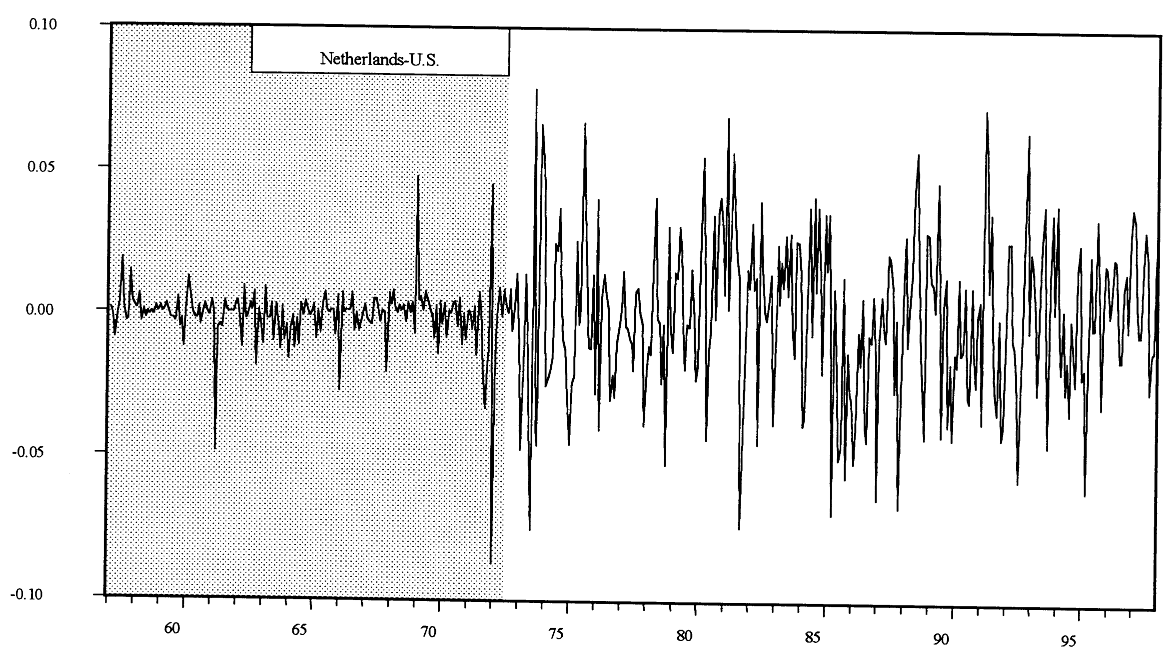


Figure 1 (Concluded). Percentage Change of Real Exchange Rates over Previous Periods





after it broke down. The reason behind this "irregularity" may have been that floating had been less common among European countries than in non-European countries after Bretton Woods. Section VI will formally test the effects of the ERM arrangement on the RER volatility inside and outside the EMS.

Table 2. Statistical Summary of Real Exchange Rate Volatility Across Exchange Regimes 1/

	Fixed 1880-1913	*Free Float* 1914-26	*Fixed* 1927-31	*Controlled Float* 1932-38	*Fixed* 1946-71	*Controlled Float* 1972-97
The United States-United Kingdom RER						
	(33 periods)	(12 periods)	(4 periods)	(6 periods)	(25 periods)	(25 periods)
μ	3.37	5.12	3.96	9.45	3.69	7.68
σ	2.58	4.55	2.38	5.60	4.62	4.59
The France-United Kingdom RER						
μ	1.45	10.09	5.22	7.55	6.69	5.88
σ	0.96	9.8	3.77	8.10	9.39	5.19

IFS Index: January 1957—December 1997

	Fixed: Jan. 1957-Jul. 1971		*Flexible: Aug. 1971-Dec. 1997*	
	μ	σ	μ	σ
Belgium-United States	0.44	0.75	2.19	1.77
Denmark-United States	0.62	0.78	2.03	1.58
France-United States	0.61	1.83	2.09	1.71
Germany-United States	0.56	0.99	2.25	1.82
Ireland-United States	0.64	1.12	1.98	1.61
Italy-United States	0.43	0.37	1.95	1.65
Netherlands-United States	0.57	0.96	2.18	1.71
United Kingdom-United States	0.48	1.21	2.07	1.71

1/ μ and σ are the mean and standard deviation of the absolute percentage change over previous period of the various RERs.

To test the volatility shift conjecture, a more systematic study of the issue is presented in Section V.

IV. Unit Root Tests and Trends in the RER Series

Before measuring the temporal variations in the second moments of the real exchange rates, it is crucial to consider alternative specifications of their trends. One disagreement among scholars of international finance has been whether the purchasing power parity (PPP) holds for the real exchange rates, even for the very long-run data. In the last ten to fifteen years, a large literature has emerged on testing the validity of PPP, or equivalently the stationarity of the RER.[5] Nonstationarity in the RER may take the form of a unit root process, a deterministic trend and/or structural breaks. Using their long-run data series, Lothian and Taylor (1996) are able to reject the unit root hypothesis for the dollar-sterling and franc-sterling real exchange rates. Cuddington and Liang (1998), however, reach a different conclusion when re-examining the dollar-sterling rates, and they conclude that long-run PPP does not hold.[6] The difference in conclusions between the two studies is due to different procedures used in implementing the unit root tests.

In light of these conflicting results, the various series are first tested to see if they contain unit roots. The model is specified as follows:

$$\log q_t = \omega + \theta * time + e_t \tag{4}$$

$$(1-\rho L)A(L)e_t = B(L)\varepsilon_t \tag{5}$$

When testing the presence of unit roots, the Phillips-Perron test is used in this study to estimate the following regression equation[7]:

$$d\log q_t = \mu + \phi \log q_{t-1} + b * time + u_t \tag{6}$$

where $\phi = \rho - 1$ and $dlogq_t$ represents the first difference of the logarithm of the underlying RERs.

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[5]See Rogoff (1996) for recent references.

[6]Cuddington and Liang (1998) show that, contrary to the stationary AR(1) specification chosen by Lothian and Taylor (1996), the dollar-sterling RER during 1791-1990 is better modeled as either a difference stationary process with an MA(5) error, or a trend stationary process with an AR(1) and MA(5) error.

[7]Since u_t is shown to be heteroskedastic later in the study, Phillips-Perron unit root test is more appropriate than the augmented Dickey-Fuller test.

The equation is estimated using ordinary least squares, and then the t-statistic for the null hypothesis $\phi = 0$ (i.e., that there is a unit root in the underlying data generating process) is corrected for both heteroskedasticity and autocorrelations using Newey-West (1987) procedure.[8]

It is now well-known that unit root tests have low power, and that whether an intercept and time trend are included in the regression used to obtain the PP statistics is critical in interpreting the results. In general, the appropriate procedure is to use the general-to-specific methodology by first including both a constant and a time trend in the estimation. If the null is not reject in the most general version of the specification, the significance of the trend and intercept can then be tested in turn to see if they can be omitted, thereby increasing the power of the unit root test (Enders (1995)).

The results of the Phillips-Perron tests are reported in Table 3. Except for the franc-sterling long-run real exchange rate, the null of $\phi = 0$ can not be rejected at the conventional statistical significance levels. The third column of Table 3 indicates whether a time trend (T), or a constant (C), or neither term (N) is included when estimating equation (6), using the method described in the previous paragraph.

Table 3. Estimated t-Statistics for the Unit Root Hypothesis

	τ	*Specification in (6)*
Long-run RER series (1880-1997):		
United States-United Kingdom	-0.018	N
France-United Kingdom	-3.923 1/	T&C
Monthly RER series (1957.1-1997.12)		
Belgium-United States	-0.937	N
Denmark-United States	-0.700	N
France-United States	-0.741	N
Germany-United States	-1.512	N
Ireland-United States	-0.407	N
Italy-United States	-0.901	N
Netherlands-United States	-0.859	N
United Kingdom-United States	0.142	N

1/ Indicates rejection of the unit root hypothesis at the 5 percent significance level.

[8]The truncation lag is set equal to $4(T/100)^{2/9}$ as recommended by Newey and West (1987).

V. Econometric Evidence on Volatility Shifts in RER Across Exchange Rate Regimes

A. The Long-Run Data

The variance test

The long-run data series permit investigation of RER volatility across three different fixed and three flexible exchange rate regimes. The empirical strategy used to test for volatility shifts in RERs is to add dummy variables in the residual series:

$$Var\ (\varepsilon_t) = \mu_1 + \mu_2 * DumFixed \tag{7}$$

where $Var\ (\varepsilon_t)$ is the error variance in equation (5). *DumFixed* is the dummy variable that takes the value 1 for fixed exchange period and 0 for flexible exchange period. The regime shift dates are chosen according to Table 1.

Since Cuddington and Liang (1998) and Lothian and Taylor (1996) reach conflicting conclusions regarding the unit root test for the dollar-sterling RER series, both the stationary and difference stationary (DS) specifications are considered for this RER series. It will be interesting to see whether the test results on volatility shifts depend on how the trend is modeled. On the other hand, the franc-sterling RER is modeled only as a stationary AR(1) process.

To obtain the series of ε_t, the appropriate length in A(L) and B(L) is determined for equation (5) by examining the autocorrelations and partial autocorrelations of the squared residuals from the models defined in (4) and (5).[9] The Ljung-Box Q-statistic is checked to make sure that there are no further serial correlations in the ε_t series.

The null hypothesis is that the variances of ε_t are the same across different exchange regimes, that is, H_o: $\mu_2=0$ in model defined by (4), (5) and (7). The test results are reported in Table 4. The t-statistics are in parentheses, and are calculated using the Newey-West heteroskedasticity consistent covariance. The test results show that, regardless of whether the stationary or difference stationary model is specified for the dollar-sterling RER, μ_2 is

[9] Although it is possible that the coefficients for A(L) and B(L) are also regime dependent, the short length of some exchange regime episodes makes it very difficult to test this conjecture. For example, there are only six observations in the 1932-38 fixed exchange rate period. Hence, in the subsequent test, the coefficients for A(L) and B(L) are assumed to be the same across exchange regimes.

significantly different from zero at the 5 percent level. For the franc-sterling rate, the rejection of the null is at the 10 percent significance level.[10] The rejection still holds strongly when the post-World War II period is excluded for both RERs.

The magnitude of μ_2 measures how much the volatility associated with the flexible exchange regimes differs from that of fixed exchange regimes. Hence, it can also tell us the direction of volatility shifts. For the two long-run RER series examined here, μ_2 is negative in all the specifications considered. Therefore, there is strong evidence that the flexible exchange rate regimes have been associated with a much higher RER volatility than the fixed exchange regimes.

Table 4. Estimated Values of μ_2 and Relevant t-Statistics

	$\mu_1 \times 10^2$	$\mu_2 \times 10^2$	*Constant*	*Error Process*
		The United States-United Kingdom RER		
The Stationary Model				
1880-97	0.66 (4.85)	-0.43 (-2.84) 1/	1.52 (38.05)	$(1-0.81L)e_t=(1+0.30L)\varepsilon_t$ (15.04) (3.46)
1880-39	0.63 (4.22)	-0.48 (2.61) 2/		
The DS Model				
1880-97	0.70 (7.07)	-0.46 (-3.51) 1/	1.02E-03 (0.24)	$e_t=(1-0.26L^5)\varepsilon_t$ (-2.66)
1880-39	0.64 (4.42)	-0.45 (2.51) 2/		
		The France-United Kingdom RER		
1880-97	1.01 (3.82)	-0.60 (-1.73) 2/	-1.36 (-36.98)	$(1-0.77L)e_t= \varepsilon_t$ (-11.28)
1880-39	1.48 (4.62)	-1.44 (-3.66) 1/		

1/ Indicates that the null hypothesis can be rejected at the 5 percent significance level.
2/ Indicates that the null hypothesis can be rejected at the 10 percent significance level.

[10] The weaker rejection for the franc-sterling RER during the whole sample period may be because of the fact that floating has been much less free in Europe than in other parts of the industrial world after the breakdown of the Bretton Woods system. Analysis in Section VI shows that during the brief period when Britain and France joined the ERM, the volatility of their bilateral RER against the deutsche mark was significantly lower.

Wald-Wolfowitz test

The analysis in Section II suggests that the Wald-Wolfowitz test should be used with caution when testing whether the observed volatilities from two samples are from the same distribution. The main lesson is that the volatility should be measured against an appropriately specified mean and the deviation from the mean should be independent. This subsection will apply the W-W test to the ε_t series obtained in the previous section. The null hypothesis is that the observations of the residuals during the fixed and floating regimes belong to the same population.

The results of the test are summarized in Table 5. If the dollar-sterling RER is a random walk process, the null can be rejected regardless of whether the whole sample period (1880-97) or only the pre-World War II period is examined. If instead, the dollar-sterling RER is a mean-reverting stationary process, the correct procedure of applying the W-W test is to obtain the ε_t series as the innovations from the statistical models in equation (4) and (5). For the dollar-sterling RER, the null can be strongly rejected for the whole sample period, but not for the pre-World War II period. For the franc-sterling RER series, however, the null can be strongly rejected regardless of which sample period is used.

The unit root test result indicates that for the period of 1880-97, the dollar-sterling RER is better modeled as a difference stationary process. Hence, the results in Table 5 may be safely taken as rejection of the null for the dollar-sterling RER. Besides the dollar-sterling RER, the rejection of the null also holds strongly for the franc-sterling RER over more than a century. Therefore, it seems that the conclusion of this study is robust to sample selection.

Table 5. Wald-Wolfowitz Test for the RER Volatility

		U	*Significance Level*
The Dollar-Sterling RER			
The DS Model	1880-97	-1.94 1/	0.027
	1880-38	-1.30 2/	0.099
The Stationary Model	1880-97	-3.32 1/	0.001
	1880-38	-0.72	0.237
The Franc-Sterling RER			
	1880-97	-2.53 1/	0.006
	1880-38	-4.24 1/	0.000

1/ Indicates that the null hypothesis can be rejected at the 5 percent significance level.
2/ Indicates that the null hypothesis can be rejected at the 10 percent significance level.

Although the test results suggest the rejection of the null that the underlying distributions of the shocks are the same across exchange rate regime, they still leave the question open as to which direction the volatility has shifted. The analysis in the previous subsection indicates that short-term fluctuations in the RERs increased under flexible exchange rate periods.

B. The RER Behavior and the Bretton Woods System

The last section provided some strong evidence of volatility shifts across exchange rate regimes using long-run annual data series. There are, however, only a few observations during some of the brief exchange regime episodes between the two World Wars. This may leave the skeptical readers unconvinced. Therefore, in this section the **monthly** RER series are constructed by using the IMF's *International Financial Statistics* database. These series are examined to see whether there is a change in variance associated with the breakdown of Bretton Woods fixed exchange rate commitment. The sample of the countries studied extends to include another six European countries, and later analysis intends to investigate whether the EMS arrangement had also systematically influenced the RER behavior. The RER series are constructed by using the price level of the United States as p^* in equation (3).

A univariate generalized autoregressive conditional heteroskedastic model (GARCH) is used here to account for time varying variance and covariance. The aim of the exercises is to examine the extent to which different nominal exchange rate regimes influence the parameters in the GARCH process for the RER series. If changes in the exchange regimes lead to parameter shifts in GARCH, it implies changes in the degree to which shocks to the RER volatility persist over time. First, the RER series are examined to see whether GARCH provides a good fit for the behavior of various monthly RER series, ignoring the possible regime shift effects on the conditional variances. Then, the second part of the exercise introduces the exchange regimes shift factor into the GARCH specification.

Application of the GARCH model to the RER series

It has long been observed that changes in the nominal and real exchange rates tend to be leptokurtic, that is, they exhibit "fat tails." Visual inspections of Figure 1 reveal volatility clustering: large changes tend to be followed by large changes, and small changes tend to be followed by small changes. After an appropriate specification of the conditional mean using equation (4) and (5), the residual series is modeled as a GARCH (p,q) process:

$$\varepsilon_t | I_{t-1} \sim N(0, h_t) \tag{8}$$

$$h_t = \delta + \sum_{i=1}^{p} \alpha_i \varepsilon_{t-i}^2 + \sum_{i=1}^{q} \beta_i h_{t-i} \tag{9}$$

The variance today depends on past news about volatility (the ε_t^2 terms) and past forecast variance (the h_t terms). The inclusion of lagged conditional variances might capture some sort of adaptive learning mechanism.

For most financial time series, GARCH (1,1) provides a sufficiently good fit. This is also true for the variables studied here. Sufficient conditions for well-defined variance and covariance only require $\alpha>0$, $\beta>0$, and $\alpha+\beta<1$.

Table 6 presents the test statistics for both the mean and variance estimation. For the monthly RER series, GARCH (1,1) provides a sufficiently good fit. All the ARCH and GARCH terms are significantly greater than zero, which is a strong indication of the appropriateness of the GARCH specification for the RER series. Diagnostic checks (Ljung-Box statistics and ARCH LM tests) have been carried out to insure that the resulting innovations in the estimated error processes have neither serial correlation nor ARCH effects.

The squared residuals of GARCH (1,1) can be written as an ARMA (1,1) process. Specifically, ε_t can be decomposed into its conditional expectation (h_t) plus an innovation (v_t) term. The latter by definition is unpredictable based on the past:

$$\varepsilon_t^{\,2} = h_t + v_t \tag{10}$$

Substituting equation (10) into (9), an alternative expression for the squared residuals is as follows:

$$\varepsilon_t^2 = \delta + (\alpha+\beta)\varepsilon_{t-1}^2 + v_t - \beta v_{t-1} \qquad (11)^{11}$$

The conditional variance of ε_t is $E_{t-1}\varepsilon_t^2 = h_t$. The unconditional variance exists when $\alpha+\beta<1$, and is defined as:

$$\sigma^2 = \tfrac{\delta}{1-\alpha-\beta} \tag{12}$$

When σ^2 exist and is independent of time, the GARCH process is stationary. The sum $\alpha+\beta$ measures the persistence of volatility shocks. For many financial time series, this sum is very close to 1, that is, shocks die out very slowly. For the RER series of Belgium, Denmark, Ireland, Italy and United Kingdom, the sum is greater than one.

[11]Note that v_t is not a white noise innovation.

Table 6. Estimation of the GARCH (1,1) Model

(The real exchange rates against the U.S. dollar between Jan.1957-Dec.1997)

	$\delta \times 10^4$	α	β	*Log Likelihood*
Belgium	0.03 (6.45)	0.10 (7.13)	0.90 (72.92)	1328.59
	$Dlogq_t=-0.0013+e_t;\ e_t=(1+0.23L)\varepsilon_t$ (-1.70) (4.57)			
Denmark	0.05 (4.14)	0.16 (6.81)	0.86 (51.25)	1271.89
	$Dlogq_t= -0.0001+e_t;\ e_t=(1-0.15L)\varepsilon_t$ (-0.12) (2.72)			
France	2.99 (4.96)	0.30 (4.04)	0.27 (2.00)	1139.21
	$Dlogq_t= -0.0003+e_t;\ (1-0.20L)\ e_t=\varepsilon_t$ (-0.22) (3.50)			
Germany	0.06 (4.65)	0.06 (5.90)	0.93 (75.46)	1232.58
	$dlogq_t= -0.0020+e_t$ (-2.26)			
Ireland 1/	8.97E-03 (1.63)	0.13 (5.52)	0.89 (52.30)	1346.03
	$Dlogq_t= -0.0015+0.15dum1967+e_t;\ e_t=(1+0.28L)\varepsilon_t$ (-2.91) (4.65) (5.47)			
Italy	5.10E-03 (1.13)	0.18 (7.78)	0.84 (58.27)	1385.45
	$Dlogq_t =-0.0018+e_t;\ e_t=(1+0.16L)\varepsilon_t$ (-10.13) (4.58)			
Netherlands	0.05 (5.29)	0.08 (6.96)	0.91 (70.01)	1249.21
	$Dlogq_t =-0.0007+e_t;\ (1-0.13L)e_t=\varepsilon_t$ (-0.89) (2.74)			
United Kingdom[1]	3.91E-03 (1.32)	0.13 (5.80)	0.89 (50.17)	1391.43
	$Dlogq_t = -0.0009+0.16dum1967+e_t;\ e_t=(1+0.26L)\ \varepsilon_t$ (-2.08) (16.56) (5.63)			

1/ Note a "spike" dummy for observation in December 1967 (*dum1967*) is included in the conditional mean equation for Ireland and United Kingdom to account for the one-time dramatic realignment of the nominal exchange rate during the Bretton Woods period.

When $\alpha+\beta=1$, the ARMA process for ε_t^2 would have a unit root and the GARCH process is said to be integrated in variance (IGARCH) (Engle and Bollerslev (1986)). In this case the unconditional variance of ε_t is infinite, even though it is still possible for ε_t itself to be a strictly stationary process (Nelson (1990)). For IGARCH process, "current information remains important for the forecasts of the conditional variances for all horizons" (Engle and Bollerslev (1986), p27).

The empirical analysis in the next section will examine whether σ^2 is constant over time or it depends on the shifts in the exchange rate regimes.

Accounting for regime shifts in GARCH

It is potentially restrictive to assume that δ is constant over time. Shifts in exchange regimes may lead to shifts in policy parameters of national governments and/or shifts in the optimal response functions of economic agents. Lastrapes (1989) demonstrates that changes in the U.S. monetary policy had significant impact on the U.S. nominal exchange rate volatility. Therefore, it is interesting to see whether the breakdown of the Bretton Woods system caused shifts in the conditional and unconditional variances of the RERs.

To test this conjecture, a dummy variable, *Fixed*, indicating the presence of a fixed exchange regime, is added to equation (9):

$$h_t = \delta_1 + \delta_2 * Fixed + \alpha\varepsilon_{t-1}^2 + \beta h_{t-1} \tag{13}$$

where $Fixed=1$ for observations between January 1957 and June 1972, and 0 otherwise.

Equation (13) reduces to equation (9) if $\delta_2=0$, that is, when the intercept of the conditional variance equation does not depend on the exchange rate regimes. Likelihood ratio tests can be employed to examine parameter shifts in the GARCH process by nesting the restricted specification within a general unrestricted specification.

The estimation results in Table 7 show that the null of $\delta_2=0$ can be statistically rejected for all RER series. Adding the regime shift dummy greatly improves the fit of the model. The likelihood ratio statistic rejects the restriction of $\delta_2=0$ at less than 5 percent level in all cases, except for the deutsche mark-dollar RER where the rejection is at the 10 percent level. Moreover, the Bretton Woods period was associated with a much lower conditional variance and hence a lower unconditional variance. These conclusions also hold for the franc-dollar RER series. Unlike the frank/sterling rate, the French RER against the U.S. dollar clearly shows increasing volatility after the Bretton Woods fixed exchange rate commitment was abandoned.

In addition, the sum of α and β dramatically declines, and is no longer greater than one for the five RER series.[12] The evidence reported here supports Diebold's (1986, p.55) conjecture that regime shifts may cause the appearance of IGARCH. Hence, if there are regime shifts in the underlying data generating process, failing to account for it may lead to conclusions that shocks to RER volatility are more persistent than they actually are.

Table 7. Accounting for Exchange Regime Shift in the GARCH (1,1) Model

	$\delta_1 \times 10^3$	$\delta_2 \times 10^3$	α	β	*log likelihood*	*Likelihood ratio*[13] *Ho:* $\delta_2 = 0$
Belgium	0.12 (7.22)	-0.11 (-2.6E+101) 1/	0.15 (5.86)	0.69 (15.08)	1337.56	17.94 1/
Denmark	0.11 (7.95)	-0.09 (-3.3E+101) 1/	0.25 (5.24)	0.61 (14.88)	1281.88	19.99 1/
France	0.63 (8.10)	-0.26 (-3.83) 1/	0.26 (3.58)	-0.10 (-1.37)	1145.36	12.30 1/
Germany	0.04 (2.63)	-0.03 (-1.2E+101) 1/	0.11 (7.22)	0.83 (34.52)	1235.12	5.08 2/
Ireland	0.17 (2.46)	-0.16 (-2.45) 1/	0.11 (2.76)	0.62 (4.44)	1358.14	24.21 1/
Italy	0.058 (8.37)	-0.055 (-6.1E+101) 1/	0.15 (3.69)	0.77 (21.64)	1394.20	17.49 1/
Netherlands	0.60 (5.44)	-0.53 (-5.21) 1/	0.28 (6.19)	-0.02 (-0.17)	1268.53	38.64 1/
United Kingdom	0.21 (3.60)	-0.20 (-3.58) 1/	0.15 (3.29)	0.59 (6.06)	1400.80	18.74 1/

1/ Significant at the 5 percent level.
2/ Significant at the 10 percent level.

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[12] For the RER series of France and Netherlands, when adding the regime shift factor into the conditional variance equation, the GARCH term becomes negative. Although it is not significant, a negative GARCH term can not ensure the conditional and unconditional variances to be positive for all realizations of ε_t.

[13] The statistic has a chi-square distribution with one degree of freedom under the null hypothesis.

VI. RER VOLATILITY: DOES THE EMS MATTER?

After the breakdown of Bretton Woods system, the resistance to floating exchange rates remained intense in Europe. In March 1979, the European Monetary System became effective as a new initiative to stabilize exchange rates among the principal members of the European Community.[14] The main technical detail of the EMS included that the bilateral exchange rates between each pair of participating countries were allowed to fluctuate within ±2.25 percent of their parity rates.[15] Parity adjustments required mutual agreement by EMS members. The EMS evolved through several stages towards more stability before September 1992, when it underwent serious speculative attacks. Several EMS currencies widened the intervention band to ±15 percent afterwards.

The effects of the EMS on monetary convergence among its member countries have been studied quite intensively.[16] Empirical evidence also suggests that the EMS had a smoothing effect on the adjustment in trade prices.[17] Figure 2 plots the monthly rate of change of six major EMS-members' RERs against the deutsche mark since 1972.[18] Contrary to Figure 1, which is the plot of these countries' RER against the U.S. dollar, the RERs within the EMS seem to be much more stable in the 1980s.

This section will formally investigate two hypotheses: (1) RER volatility between six major EMS-member currencies against the deutsche mark were significantly lower during the 1979-92 period, and (2) the same is true when the U.S. dollar is used as the numerator country. If both hypotheses are true, then the observed lower RER volatility in Europe is likely to be specific to the particular sample period of 1980s. If only the first contention is valid, however, it can be taken as strong evidence supporting the main conjecture of this paper, that is, the nominal exchange rate arrangement does matter for RER behavior.

Before undertaking the formal empirical analysis on volatility patterns inside and outside the EMS, the unit root tests and a preliminary statistical summary of the sample period is presented in Table 8. During the period between 1972:06 to 1997:12, the unit root

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[14]The initial members included Belgium, Denmark, France, Germany, Ireland, Italy, and Netherlands. Britain only joined in October 1990.

[15]Some weaker currencies were allowed to fluctuate within 6 percent of their parity rates.

[16]See Giavazzi and Giovannini (1989), MacDonald and Taylor (1990), and Ungerer et al. (1990).

[17]See Sapir and Sekkat (1995), and Bourdet (1996).

[18]Shaded areas correspond to the period when the ERM was effective.

hypothesis is not rejected for the EMS-RERs studied here. This implies a failure of the long-run PPP hypothesis. The rejection may be due to low power of the tests because of the relatively short-horizon of the data. On the other hand, PPP may not hold, as several nonprice shocks to the EMS-members occurred during this period (Classen and Peree, 1988).

Table 8. Unit Root Tests and Volatility Statistics for the RERs (1972:6-1997:12)

	Phillips-Perron Unit Root Test[19] *(τ)*	$\mu x10^3$	$\sigma x10^3$
RERs Against the Deutsche mark			
Belgium	0.11	2.20	1.70
Denmark	0.27	2.04	1.61
France	0.09	2.13	1.69
Ireland	-1.01	1.23	1.42
Italy	0.19	1.20	1.49
Netherlands	0.77	0.51	0.70
United Kingdom	-0.19	2.09	1.76
RERs Against the U.S. dollar			
Belgium	-0.48	2.19	1.77
Denmark	-0.57	2.03	1.58
France	-0.54	2.09	1.71
Germany	-0.88	2.25	1.82
Ireland	-0.63	1.98	1.61
Italy	-0.34	1.95	1.65
Netherlands	-0.51	2.18	1.71
United Kingdom	0.04	2.07	1.71

To test hypotheses (1) and (2), exercises similar to those performed in the previous section are applied to the different RER series. GARCH (1,1) again provides a sufficiently good fit for the RER series.

The role of the EMS on the RER volatility is examined by estimating the following conditional variance equation:

$$h_t = \lambda_1 + \lambda_2 * dumEMS + \alpha\varepsilon_{t-1}^2 + \beta h_{t-1} \tag{14}$$

where *dumEMS=1* for observations between March 1979 (October 1990 for Britain) and August 1992, and 0 otherwise.

[19]Neither the trend nor the constant term is significantly different from zero in the PP regressions. Hence they were excluded in the unit root test specification.

Figure 2. Monthly Rate of Change in EMS-Members' RERs Against the Deutsche mark

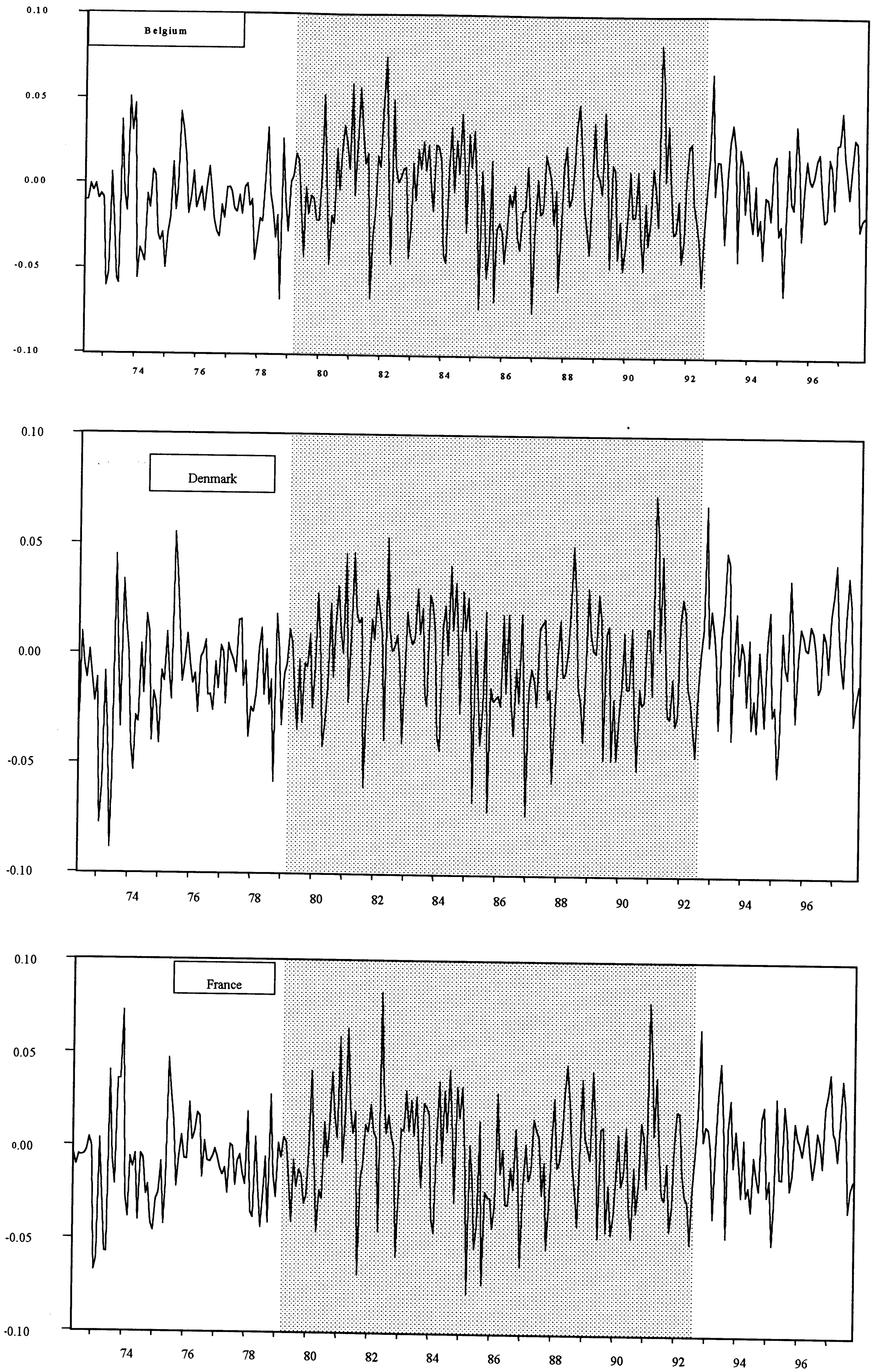
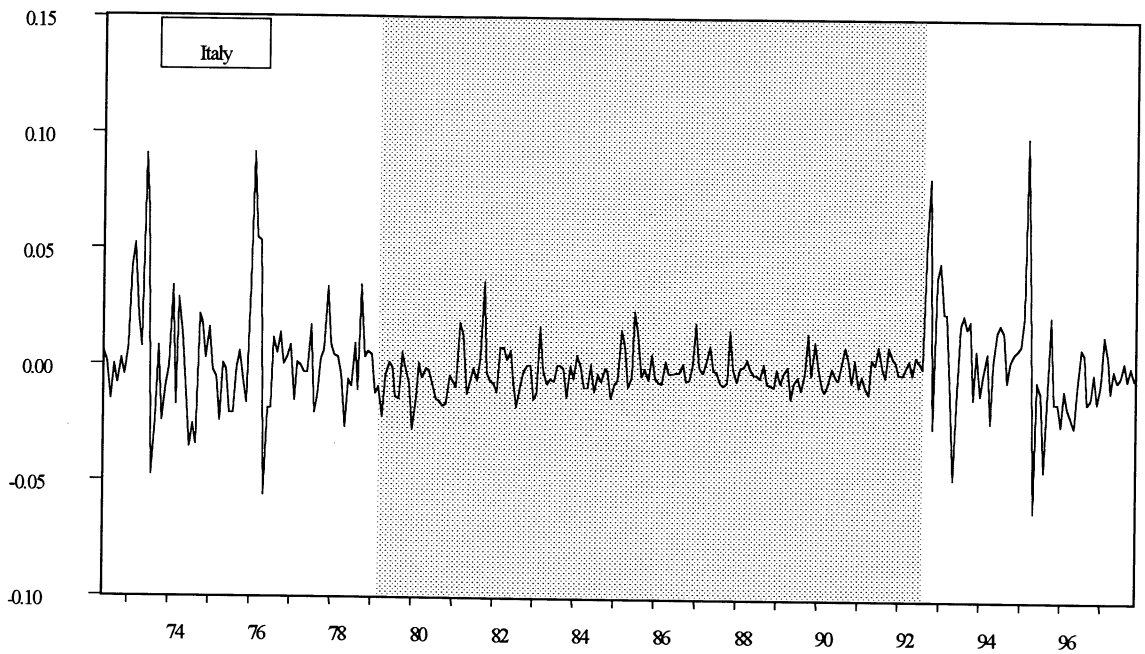


Figure 2 (Continued). Monthly Rate of Change in EMS-Members' RERs Against the Deutsche mark



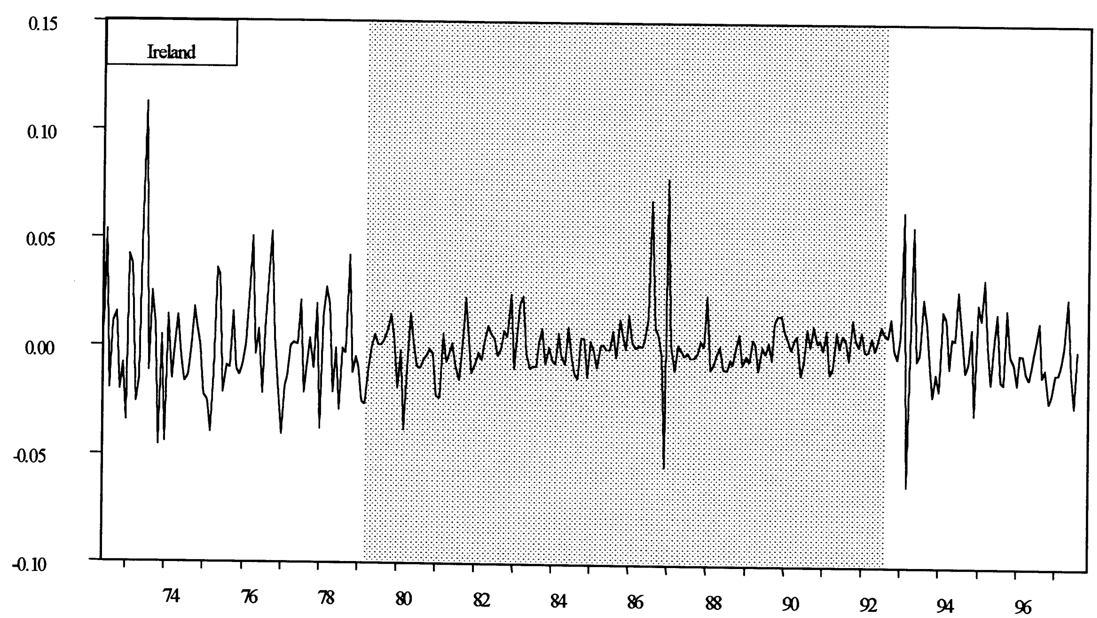
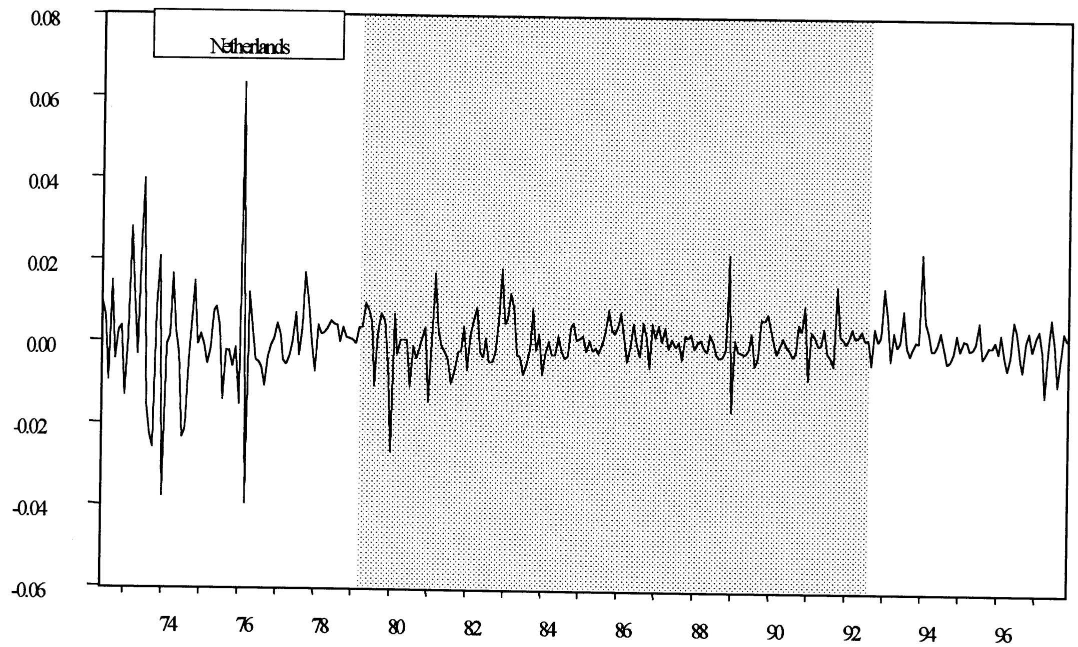
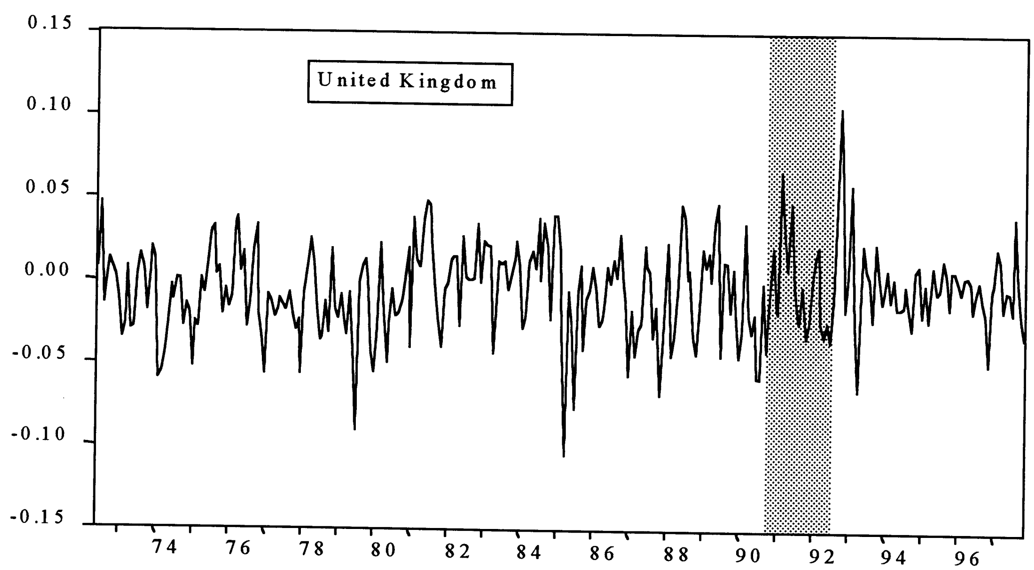


Figure 2 (Concluded). Monthly Rate of Change in EMS-Members' RERs Against the Deutsche mark





The ERM period of 1979-92 provides a controlled experiment in investigating the impact of nominal exchange rate arrangements on the exchange rates. The test statistics for both the mean and variance estimation of equation (9) and (11) are given in the Appendix I. Table 9 presents the value and t-statistics of λ_2.

The real exchange rate volatility between EMS-members' currencies was significantly lower during the time when the ERM was operated effectively. λ_2 is statistically different from zero for all the RERs between the EMS-members, except for Denmark. In each case, λ_2 is negative. On the other hand, the EMS-members' currencies jointly float against other non-EMS currencies, such as the U.S. dollar. For the eight EMS-members studied here, λ_2 is significantly different from zero only for two RER series against the dollar. In addition, in each case, λ_2 is positive, indicating higher RER volatility associated with the EMS-period. The simultaneous occurrence of decreased volatility in RERs inside the EMS and unchanged/increased volatility outside the EMS is difficult to be explained by the specific conditions of the world economy during the sample period. It provides additional strong support for the hypothesis that the RER behavior seems to be systematically linked to the nominal exchange rate arrangements.

Table 9. RER Volatility and the EMS

RERs Against the Deutsche mark		***RERs Against the U.S. dollar***	
	λ_2		λ_2
Belgium	-1.22E-05 (-3.67) 1/	*Belgium*	6.74E-05 (1.36)
Denmark	-2.11E-07 (-0.21)	*Denmark*	3.50E-05 (0.78)
France	-3.35E-06 (-1.91) 2/	*France*	6.79E-05 (1.50)
Ireland	-1.15E-04 (-4.76) 1/	*Ireland*	1.21E-05 (2.68) 1/
Italy	-2.67E-04 (-5.77) 1/	*Italy*	1.07E-04 (2.45) 1/
Netherlands	-5.74E-06 (-3.14) 1/	*Netherlands*	6.45E-05 (1.01)
United Kingdom	-1.01E-04 (-1.78) 2/	*United Kingdom*	1.06E-04 (0.85)

1/ Significant at the 5 percent level.
2/ Significant at the 10 percent level.

VII. CONCLUSIONS

There has been renewed interest in the relative merits of alternative exchange rate regimes in light of the Asian currency crisis beginning in July 1997, and the fact that a common currency area in eleven European countries is scheduled to begin in January 1999. One of the key issues in the heated debate about regimes is whether and how nominal exchange rate arrangement affects the behavior of real macroeconomic variables.

Following upon the controversial results reported by Grilli and Kaminsky (1992), this paper re-investigates the differences in real exchange rate volatility across fixed and flexible exchange rate regimes. Using two long-run RER series dating back to 1880, as well as monthly RER observations from 1957, this study finds strong evidence supporting the long-held suspicion of international macroeconomists that flexible exchange rate periods have been associated with a higher RER volatility than fixed exchange periods. Also, regime shifts play an important role in determining the persistence of shocks to the RER.

From a theoretical standpoint, this confirmation of regime-dependent behavior of the real exchange rates has important implications for the plausibility of various macroeconomic models of exchange rate determination. As Stockman (1983) highlights, two types of models have been offered by economists to study alternative exchange rate regimes. Sticky-price models show that adopting floating exchange rates implies higher nominal and real exchange rate variability.[20] In contrast, a class of equilibrium models satisfies "the nominal exchange regime neutrality proposition," i.e. the time series properties of all real variables are invariant to the choice of exchange regime.[21] Stockman (1983) illustrates, however, that some equilibrium models, for example those with nontradable as well as tradable goods—not just sticky price models—may also exhibit nonneutrailities with respect to the exchange rate regime.

Interestingly, Cuddington and Liang (1998) show that the relative prices of primary commodities in terms of manufactures have time series characteristics that also vary across exchange rate regimes. The nonneutrality of nominal exchange regimes on two types of internationally *traded* goods, primary commodities and manufactures, is hard to explain in the context of equilibrium models, including those described by Stockman (1983).

Research on real exchange rates has recently shifted to examine the importance of imperfect competition in international markets. Several empirical studies considered the price-exchange rate relationships since the 1970s.[22] Economists are increasingly convinced that

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[20]See, e.g., Dornbusch (1976), Frenkel (1981), and Mussa (1982).

[21]See, e.g., Helpman (1981), Lucas (1982) and Stockman (1980).

[22]See Goldberg and Knetter's (1997) excellent survey on goods prices and exchange rates.

prices of different goods respond differently to exchange rate changes, reflecting the underlying market structure across industries. The regime-dependent behavior of the real exchange rate appears to be largely the result of market imperfections. This is also the reason behind persistent deviations from law of one price, and hence deviations from purchasing power parity.

The findings of this study shed light on the on-going debate on the desirability of alternative exchange systems. One of the main reasons why people care about the regime-dependent behavior of real macroeconomic variables is due to the so-called *social inefficiency* associated with floating exchange rates (Hallwood and MacDonald, 1994). If there are nominal rigidities in goods prices, nominal exchange rate changes lead to changes in real exchange rates. Moreover, if goods with different underlying market structures adjust to exchange rates changes differently, the relative prices of two traded goods will also change. Such changes may result in inefficient allocation of goods and services.

Analysis of RER Volatility Inside and Outside the EMS

This appendix gives the test results of applying the GARCH model to the RER series inside and outside the EMS during 1972:6-1997:12.

Table 10 shows that GARCH (1,1) provides a good fit for the RER series. It contains the estimation results for the conditional mean and variance for the EMS-members' RER against the deutsche mark , without taking into consideration the possible effects of the EMS on the RER volatility:

$$h_t = \lambda + \alpha \varepsilon_{t-1}^2 + \beta h_{t-1} \tag{9}$$

Table 10. GARCH (1,1) Estimation for the EMS-Member's RER Against the Deutsche mark (Jun. 1972-Dec. 1997)

	$\lambda x10^4$	α	β	Log likelihood
Belgium	0.16 (7.78)	0.52 (7.37)	0.31 (5.55)	1097.10
	$Dlogq_t$=-0.0005+e_t; e_t=(1+0.14L)ε_t *(-1.33)* *(1.64)*			
Denmark	0.03 (3.84)	0.003 (0.22)	0.95 (55.16)	995.70
	$Dlogq_t$= -0.0005+e_t; e_t=(1-0.12L)ε_t *(-0.70)* *(-1.93)*			
France	0.04 (3.85)	0.17 (4.51)	0.80 (26.67)	1004.75
	$dlogq_t$= -0.0005+e_t; (1-0.31L) e_t=ε_t *(-0.62)* *(5.11)*			
Ireland	0.52 (6.11)	0.41 (5.46)	0.49 (8.89)	818.17
	$dlogq_t$= -0.0002+e_t *(-0.23)*			
Italy	0.42 (7.82)	0.42 (8.51)	0.52 (11.19)	858.27
	$dlogq_t$ =-0.0022+e_t; e_t=(1+0.28L)ε_t (-2.39) (4.13)			
Netherlands	0.10 (5.90)	0.48 (9.26)	0.46 (8.54)	1091.74
	$dlogq_t$ =0.0010+e_t; e_t=(1+0.17L)ε_t *(2.90)* *(2.61)*			
United Kingdom	0.62 (32.77)	0.14 (11.27)	0.74 (80.17)	742.42
	$dlogq_t$ = -0.0006+e_t; (1-0.33L) e_t=ε_t *(-0.84)* *(10.43)*			

The ERM period of 1979-92 provides a controlled experiment in investigating the impact of nominal exchange rate arrangements on the exchange rates. The test statistics for both the mean and variance estimation of equation (14) are given in the Table 11. It shows how the GARCH parameters have changed when the EMS factor is included in the conditional variance specification.

$$h_t = \lambda_1 + \lambda_2 * dumEMS + \alpha\varepsilon_{t-1}^2 + \beta h_{t-1} \qquad (14)$$

Table 11. Conditional Volatility of RERs and the EMS

	$\lambda_1 \times 10^3$	$\lambda_2 \times 10^3$	α	β	***Log likelihood***	***Likelihood Ratio Ho: $\delta_2 = 0$***
Belgium	0.04 (14.02)	-0.01 (-3.67) 1/	0.47 (7.70)	-0.04 (-1.43)	1102.30	10.37 1/
Denmark	3.08E-03 (2.22)	-2.11E-04 (-0.21)	-0.005 (-0.53)	0.96 (58.20)	996.44	1.48
France	6.86E-03 (3.53)	-3.35E-03 (-1.91) 2/	0.14 (3.50)	0.82 (23.85)	1005.68	1.87
Ireland	0.15 (6.12)	-0.12 (-4.76) 1/	0.38 (5.91)	0.41 (11.45)	830.64	24.95 1/
Italy	0.31 (5.34)	-0.27 (-5.77) 1/	0.17 (2.47)	0.28 (2.50)	884.43	52.32 1/
Netherlands	0.01 (5.45)	-5.74E-03 (-3.14) 1/	0.50 (9.08)	0.44 (7.47)	1093.35	3.22
United Kingdom	0.15 (2.19)	0.10 (-1.78) 2/	0.18 (2.86)	0.52 (3.67)	744.49	4.15

1/ Significant at the 5 percent level.
2/ Significant at the 10 percent level.

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