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A LONG-RUN VIEW OF THE EUROPEAN MONETARY SYSTEM

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Abstract

This paper analyzes the exchange rates and consumer price indices of the six largest countries of the European Monetary System (EMS). The analysis covers the entire period of floating exchange rates. This paper shows that many of the implied real exchange rates have unit roots, even when one allows for the possibility of a structural break occurring at the time of the formation of the EMS. Further, prices and exchange rates are not co-integrated during the EMS period. There is strong evidence that there is a quadratic time trend in these price indices and weak evidence that exchange rates and prices were more highly co-integrated before the advent of the EMS. The data suggest that the eleven realignments of the EMS between 1979 and 1988 have not served fully to offset the member countries' inflation differentials.

A Long-Run View of the European Monetary System

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I. Introduction

The European Monetary System was founded on March 13, 1979. At that time, it was seen as another step in the unification of the economic policies of the countries of the European Community. It established the European Currency Unit (ECU) in an effort to replace the dollar as a reserve currency, in Europe at least. Nearly ten years have lapsed since the EMS' inception; it should now be possible to assess whether the EMS has satisfied a more modest goal: that of providing enough stability in European prices so that intra-European trade can be conducted according to real, not monetary, forces. This paper assesses how the EMS exchange rate arrangement has influenced real exchange rates and relative prices by comparing regimes before and after the advent of the EMS. These issues are fundamentally empirical.

The EMS is a system of quasi-fixed exchange rates. There are ten currencies that participate in the arrangement; these are: the Belgian franc,

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Italian lira, the Luxembourg franc, the Dutch guilder, the pound sterling, and the Greek drachma. The currencies are allowed to fluctuate within bands of 2.25% around an ECU central rate,[2] and the central banks of a pair of member countries are obliged to intervene in the case that their currencies are at the limits of the divergence bands. Although the drachma and the British pound are used to define the ECU, neither the British nor the Greek government has agreed to participate in the intervention mechanism. There have been eleven realignments of the system.[3] When a realignment occurs, a currency's ECU central rate is redefined, and the revaluations or devaluations of the currencies are designed to offset some or all of the changes in relative price levels.[4] [5]

The purpose of this paper is to examine the long-run trends of the real exchange rates of the major currencies participating in this system. Movements in real exchange rates are due either to changes in relative prices or to realignments of the EMS, which have been relatively infrequent. This paper proposes to test the proposition that EMS countries with high inflation have lost competitiveness to low inflation countries. We measure the level of competitiveness by the real exchange rate; thus, a loss of competitiveness is

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2 The lira is allowed to fluctuate within 6% bands, but in practice the Italian central bank intervenes to keep its currency within the narrower divergence limits. The earlier Snake arrangement operated in a similar fashion but without having the ECU as a way of defining the central rates. The following countries participated in the Snake: Belgium, Denmark, France, Germany, Italy, Netherlands, Norway, Sweden and the United Kingdom. Not all of these countries, however, participated continuously in the Snake.

3 The Appendix lists the dates of the initial formation and all realignments of the EMS.

4 We do not intend in this paper to enter into the details of the EMS exchange rate mechanism; for a good treatment of these points, see Grabbe (1986).

5 A re-composition of the ECU is different from a realignment; a re-composition involves changing the actual number of units of each currency in defining the ECU. A re-composition of the ECU occurred on 14 September 1984, the date at which the Greek drachma was added to the definition of the ECU.

an appreciation of the real exchange rate. If each realignment compensated high inflation countries for losses of competitiveness, then the real exchange rate would tend to exhibit no long-run trend. Giavazzi and Pagano (1987) argue that real appreciation of high inflation members is characteristic of the EMS and suggest that this helps to discipline the high inflation countries. The implication of this view is that realignments only partially offset inflation differentials.

To test this idea, we consider three hypotheses. First, we test whether the exchange rates and the relevant price indices are co-integrated. Second, we test whether long-run real exchange rates are stationary when one allows for a structural change in these series at the advent of the EMS; this is the first application in international finance of a unit root test that allows for a structural break. Third, we test if the rate of depreciation of these real exchange rates is different after the advent of the EMS.

Since the dissolution of the Bretton Woods agreement, exchange rates of the major industrial countries have exhibited behavior indistinguishable from that of random walks.[6] Because the EMS is akin to a system of fixed exchange rates, it is perhaps surprising that the fifteen bilateral exchange rates we analyze are also essentially random walks. This fact is plausible, however if one considers that there have been frequent enough realignments of the EMS so that bilateral rates are not sufficiently rigid as to appear stationary to the statistician.[7]

Several other economists have used the co-integration approach in the recent literature on exchange rates. Corbae and Ouliaris (1986) propose a

6 See Meese and Singleton (1982) for an early discussion of unit roots in exchange rate data and Meese and Rogoff (1983) for an analysis of the out-of-sample forecasting properties of several exchange rate models.
7 If there had been no realignments of the EMS, then it would be surprising to find exchange rates and prices co-integrated except in the degenerate case where the relevant price indices themselves were stationary.

robust test for unit roots in the foreign exchange market; they show that weekly spot and forward rates for six currencies have a unit root, whereas the implied risk premia are stationary. Patel (1986) examines quarterly data for currencies and price indices of Canada, Germany, the Netherlands, the United Kingdom, and the United States; he shows that the currencies of Canada, Germany and the United States were co-integrated with their respective price indices. Mark (1986) investigates monthly data covering the currencies and price indices of Belgium, Canada, France, Germany, Italy, Japan, the United Kingdom, and the United States from June 1973 through June 1985. He concludes that the null hypothesis of no co-integration cannot be rejected at standard significance levels. We find similar results for several of these currencies during our entire sample, but there is some contrast in the behavior of the European currencies before and after the advent of the EMS.

Huizinga (1986) examines the spectral properties of ten different exchange rates, using monthly data for the thirteen years between 1974 and 1986. He studies the low frequency properties of many of the implied real exchange rates and argues that these series exhibit mean reversion. He states that he finds no difference in the long-run behavior of the European exchange rates from the exchange rates of other countries. We find some evidence to the contrary simply by breaking the sample into two periods. Kaminsky (1987) analyzes data on the U.S. dollar, yen, pound sterling, mark, and Swiss franc covering the period from April 1973 through April 1986. She finds that the implied real exchange rates exhibit mean reverting behavior; hence they are not random walks.

There have been several other empirical studies of the EMS; Rogoff (1985), Giavazzi and Giovannini (1985), Ungerer, Evans, Mayer, and Young (1986), Collins (1987, 1987a) and Giavazzi and Pagano (1987) are recent examples. None of these papers studies the co-integrative properties of exchange rates.

We structure the rest of the paper as follows. In Section II, we present a simple description of a relationship between price indices and exchange rates. Section III presents a preliminary analysis of the data, analyzing whether these time series have unit roots; we break the sample into two periods, one covering the period before the advent of the EMS and the other covering the time since then. Section IV presents the results of our tests for co-integration between the nominal exchange rates and the relevant consumer price indices. Section V tests whether these European real exchange rates have unit roots, even when one allows for a structural break at the advent of the EMS; we test also whether the real exchange rates show a significantly different rate of depreciation during the EMS period. Section VI presents our conclusions.

II. Models of the Time Series

The usual empirical analyses of the absolute version of purchasing power parity involve running a regression of the logarithm of an exchange rate index on a constant and the difference of the logarithm of two relevant price indices; one traditionally tests the null hypothesis that the constant is zero and the coefficient on the difference of the price indices is unity. Of course, this kind of analysis is not a test of a model of exchange rate determination; it is perhaps best interpreted as the estimate of a reduced form equation relating an asset price and two price levels.[8] We too must

8 See Levich (1985) for a discussion of empirical studies of exchange rates in the traditional genre. The new wave of this literature discusses some of these issues within the framework of an error correction model; see Edison (1985) and Edison and Klovland (1987), for early examples of this work.

admit at the outset that we are not presenting a model of exchange rate determination. Instead, we analyze the statistical properties of consumer prices and exchange rates among the major European economies.

For an arbitrary time series $\{x_t\}$, consider the model

$$x_t = \beta_0 + \beta_1 t + \beta_2 x_{t-1} + u_t. \tag{1}$$

If x_t represents the logarithm of a consumer price index at time t and if we assume that β_2 is unity, then equation (1) states that the logarithm of the price level can be modeled as a random walk with drift and a systematic time component. Using logarithms allows simple economic interpretations of the model's parameters; in particular, β_0 is the rate of consumer price inflation and β_1 is a systematic change in the inflation rate. If x_t represents the logarithm of an index of the number of units of domestic currency necessary to purchase one unit of foreign currency, then assuming that β_2 is unity implies that β_0 is the rate of depreciation of domestic currency and β_1 is a systematic acceleration in that depreciation.

We assume that the errors $\{u_t\}$ in equation (1) satisfy the following statistical properties.

Assumption:

(i) $E\, u_t = 0\ \forall t$;

(ii) $\sup_t E\, |u_t|^{\beta+\epsilon} < \infty$ for some $\beta > 2$ and some $\epsilon > 0$;

(iii) let $S_T = \Sigma\, u_t$; then $\sigma^2 = \lim_{T\to\infty} E(S_T/T)$ exists and $\sigma^2 > 0$; and

(iv) $\{u_t\}$ is a strong mixing with coefficients δ_m satisfying $\Sigma\, \delta_m^{1-2/\beta} < \infty$.

These are the assumptions that Phillips and Perron (1988) have made standard for much of the literature testing for unit roots where the error generating process $\{u_t\}$ is heterogeneous.[9] Much of the current applied work using unit roots in international finance has not imposed this assumption; it assumes an innovation process that is independently and identically distributed. This is particularly inappropriate for price indices because they display strong seasonal components. For example, several of the European countries have annual or biannual administered price increases for basic foodstuffs and housing; these prices, of course, have a strong influence on consumer price indices. Allowing for heterogeneity of the error process is also important for exchange rates because, in a regime of floating exchange rates, these are simply asset prices. There is strong evidence that first differences of these prices seem to be drawn from heteroscedastic distributions.

Let an exchange rate and two relevant price indices be generated by (1). Consider now a generalization of the notion of the real exchange rate r_t defined by

$$(2) \qquad r_t = \alpha_1 e_t + \alpha_2 p^*_t + \alpha_3 p_t$$

where e_t is the logarithm of the exchange rate index, p^*_t is logarithm of the foreign price level, and p_t is that of the domestic price level, all taken at time t. If we assume that $(\alpha_1, \alpha_2, \alpha_3) = (1, 1, -1)$, then r_t is simply an index of the real exchange rate at time t. An increase in r_t is a real depreciation. A decrease in r_t is a real appreciation; this is often called a loss of competitiveness. If we assume that exchange rates and prices are not themselves stationary but that r_t is a stationary stochastic process, then the arbitrary triple $(\alpha_1, \alpha_2, \alpha_3)$ has an interpretation as constants of co-integration.

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9 See Corbae and Ouliaris (1986, p. 376) for a good discussion of the restrictions that this assumption imposes on the stochastic process.

In the Section III, we test certain restrictions on the parameters in equation (1) for data on European exchange rates and prices. In Section IV, we examine the time series r_t defined by setting $(\alpha_1, \alpha_2, \alpha_3) = (1, a_2, a_3)$ where a_2 and a_3 are the ordinary least squares estimates of α_2 and α_3,[10] and in Section V, we examine the time series r_t defined by setting $(\alpha_1, \alpha_2, \alpha_3) = (1, 1, -1)$.

III. Preliminary Analysis of the Data

We examine the currencies of the six largest countries participating in the European Monetary System: the Belgian franc, the French franc, the German mark, the Italian lira, the Dutch guilder, and the pound sterling. The data are observations representing the value on the last business day of the month of these six currencies' noon quotes on the New York foreign exchange market. These six currencies implicitly define fifteen European bilateral exchange rates, only six of which are independent.[11] The price indices are monthly consumer prices, not adjusted for seasonality, whose sources are the six countries in our sample. These indices are typically announced during the third week of the month, but they generally involve sampling a basket of commodities in different cities during the course of the month in which the data are reported. See the appendix for a full description of the data.

For sake of comparison throughout this paper, two sub-samples are considered: the six years of the Snake and the nine years of the EMS.[12] A

10 Stock (1987) has shown that these are consistent estimates of the co-integrating constants.

11 During the rest of this paper, we shall use the terms "exchange rates" and "exchange rate indices" to mean the logarithms of the exchange rate indices, appropriately normalized. The traditional literature testing purchasing power parity almost always uses an exchange rate index; this allows the interpretation of the error in a typical regression as the percentage deviation from a base year relationship.

12 We have normalized each time series so that its initial logarithmic value is zero; this occurs in March 1973. Each time we run a regression involving a sub-sample from the EMS period, we re-normalize so that each series has a

central question we want to address is whether exchange rate behavior has changed between the two arrangements. It may not be entirely proper to draw a sharp distinction between periods before and after the advent of the EMS because the Snake was quite similar in practice to the EMS arrangement. On the other hand, it has been argued that there has been much more coordination of policy between countries during the EMS than during the Snake.

Figure 1 shows the plots of the logarithms of the fifteen EMS bilateral nominal exchange rates over the entire sample. Consider the model expressed by (1). For each series in each sub-sample, we test first the null hypothesis H_o: $\beta_2 = 1$ against the general alternative. We use the test statistic $Z_p(t(\hat{\alpha}))$ described by Park and Choi (1988) because of its computational elegance.[13] The critical values for this test are given in Fuller (1976, Table 8.5.2). The first sub-samples for each exchange rate covers the months from April 1973 through February 1979. The test statistics are reported in Table 1.

Table 1 shows that the null hypothesis that these bilateral exchange rates have a unit root can be rejected in only two cases: those of the Belgian franc-mark and Belgian franc-guilder bilateral rates. This is evidence that the Belgian monetary authorities may have used the Snake to control the rates of depreciation of their currency against the mark and the guilder.

We have also run the same test for these currencies in the period from March 1979 through May 1988. Table 2 reports these results. During the

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logarithmic value of zero in March 1979. None of the test statistics used in this paper depend upon this re-normalization.

13 All the tests based upon the work of Phillips (1987) and Phillips and Perron (1988) use an estimate of the ratio of the short-run variance to the long-run variance of the random variable S_t described in the Assumption above; throughout this paper have we have used estimates of the first four auto-correlations of the processes generating the data in constructing this ratio. The values of the test statistics seem not to vary much with the inclusion of more auto-correlations. Corbae and Ouliaris (1986) report a similar finding.

period of the European Monetary System, we could reject the null that these nominal exchange rates have unit roots in only two cases. These were the Belgian franc-French franc rate and the mark-guilder rate. Many observers of the EMS have noted that its realignments tend to maintain rough parity between the German and Dutch currencies; this sample also shows that these realignments maintain a systematic relationship between the Belgian and French currencies.

We performed the same tests on first differences of these time series, and in every case it was possible to reject soundly the null that these differenced time series had unit roots. For the sake of brevity, we do not report all thirty statistics. These tests present very strong evidence for the conclusion that none of the exchange rates is integrated of an order higher than one.

Figure 2 depicts the six consumer price indices. These series trend upward; moreover, there seems to be a systematic change in the rate of inflation in each of these countries over the entire sample period. This reduction in the rate of inflation has to do with the coincidence of the advent of the regime of floating exchange rates with the first oil price shock. Further, the EMS was started about six months before the second oil price shock. We show below that it is important to include a time trend in analyzing first differences of these series.[14] The economic interpretation of this is that the decline in inflation is an integral part of the data.

Again, we model the time series according to the process in equation (1), and we test the null hypothesis H_o: $\beta_2 = 1$ against the general

14 Patel (1986) conjectured that several of these consumer price indices were integrated of order two for the period he was examining. He did not pursue this insight by including a time trend in his unit root tests.

alternative. We used again the statistic given by Park and Choi (1988). This statistic is very useful for price indices because they have seasonal components, and the Parzen window of four lags seemed to be able to capture the process generating the data quite adequately.[15] The test statistics for the consumer price indices are presented in table 3. In only one case can we reject the null for these time series. The value of the test statistic for the case of the United Kingdom during the EMS period is strong evidence of the monetary policy of the Thatcher government. We ran the same tests on first differences of the series; to keep the number of tables manageable, we do not report these statistics. They allow us to reject soundly the hypothesis that the differenced price series have a unit root. We conclude that the price indices are integrated of order at most one.

Consider equation (1) again. For all series in each sub-sample, we now test the joint hypothesis H_o: $(\beta_0, \beta_1, \beta_2) = (\beta_0, 0, 1)$ against the general alternative. We are interested in determining whether the exchange rates or price indices show systematic changes in the rates of depreciation or inflation respectively. Again, we use a statistic developed by Park and Choi (1988) because it is robust with respect to the process generating the innovations in (1) and because of its computational elegance. This statistic is essentially an adjustment to the standard Wald statistic for H_o that takes into account the heterogeneity of the innovation process for the series. Its critical values are given in Dickey and Fuller (1981, Table VI). Table 4 presents the values of the test statistics for the nominal exchange rates in

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15 In an earlier draft of this paper, we used the Augmented Dickey Fuller Statistic in performing these tests. That statistic was quite sensitive to the filter applied to the data, and we had to apply complicated filters in order to make the errors in the estimates of equation (1) appear to be white noise.

the earlier period, and Table 5 presents those in the later period. Neither of these tables is very surprising. They confirm essentially the evidence given in tables 1 and 2 that these exchange rates have unit roots; there is some marginal evidence that Italy's rate of depreciation against Germany and the Netherlands accelerated during the first sub-sample. Italy experienced high inflation after the first oil price shock, and the other two countries had relatively low inflation during the last decade.

Tables 6 presents these test statistics for the price indices in the two sub-samples. Here the evidence is strikingly different from that of table 3. We can reject the null for two of the six indices in the first sub-sample and for every index in the later sub-sample. This is very strong evidence of the systematic decline in inflation of this decade; price indices have showed significant decelerations over much of the regime of floating exchange rates. The earlier tests seemed to indicate that most of these price indices had unit roots; we interpret the values in table 6 as indicating that the deterministic part of the time series has much explanatory power. Co-integration has to do, of course, with the stochastic element of the time series involved. This has strong implications for the kinds of co-integration tests one might run. In particular, it is probably necessary to include a quadratic time trend to capture the systematic changes in inflation if one hopes not to mis-specify the reduced form between exchange rates and prices in the last two decades. No applied work in international finance has noticed this fact.

IV. The Co-integrating Regressions

We now continue the analysis by examining whether the bilateral exchange rates are co-integrated with the appropriate pairs of price indices. If an exchange rate and a pair of price indices are co-integrated, then we interpret this as a loose form of purchasing power parity.[16] In particular, although all three time series have unit roots, there is a long-run equilibrium relationship among them. Consider equation (2) again; a traditional test of purchasing power parity constrains this relationship to be given by the triple (1, 1, -1), but a test of cointegration allows for the arbitrary triple (1, α_2, α_3), where we have imposed an implicit normalization. Hence, we are not constraining the relationship to the traditional view of purchasing power parity, but we are looking for empirical evidence of the notion that the three series cannot drift infinitely far apart in time.

There are two kinds of tests for co-integration: those that are not based upon analyses regressions residuals and those that are. We conduct one test of each kind in this section. The first is a test proposed by Park, Ouliaris, and Choi (1988). Consider running a regression based upon the model

$$e_t = \beta_0 + \beta_1 t + \beta_2 t^2 + \beta_3 p_t + \beta_4 p^*_t + u_t \tag{3}$$

where the variables are defined as above. If these exchange rate and price indices are not stationary but the error process is, then the reduced form given by (3) has the simple interpretation that these series are co-integrated, with the coefficients of co-integration defined up to the normalization factor that we have imposed implicitly on the term e_t. This test uses the J(p,q) statistic proposed by Park, Ouliaris, and Choi; the null hypothesis is that the three series are not co-integrated. This test is based upon the simple intuition that if exchange rates and two price series are indeed co-integrated, then adding higher order polynomial time trends in

16 See Mark (1986), Patel (1986), and Corbae and Ouliaris (1988) about testing purchasing power parity in both the traditional and co-integrative framework.

estimating (3) will not reduce this regression's residual sum of squares drastically.[17] We have good evidence indicating that it is necessary to include a quadratic time trend in (3); hence, we include four extra polynomial time trends in our test. Park, Ouliaris, and Choi do not give critical values for J(2,6); we had to estimate them from Monte Carlo simulations.[18] The test statistics are reported in tables 7 and 8. These tables seem to indicate that there is little co-integration of the exchange rates and relevant prices in either of the periods. We notice, however, that the J(p,q) statistics seem to be lower in general in the earlier period; this is weak evidence that there may be a higher degree of co-integration between the series in the first sub-sample.

The second test is based upon the residuals from the regression implied by

$$e_t = \beta_0 + \beta_1 p_t + \beta_3 p^*_t + u_t. \tag{4}$$

This test uses the Dickey-Fuller statistic on the residuals from this regression; we follow the two step procedure outlined in Engle and Yoo (1987).[19] For these data, this test may be unsatisfactory because it ignores

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17 One way of illustrating this intuition is to note that one can fit estimates of a time series arbitrarily well by adding increasingly higher order polynomial time trends as regressors. If the addition of several such trends does not give rise to a large relative reduction in the residual sum of squares, then the parsimonious specification of the regression given by (3) is indeed an adequate representation of the relationship between exchange rates and price indices.

18 No other work in applied econometrics has included a quadratic time trend. Hence, these test statistics are the first reported which include a second degree polynomial and heterogeneously distributed innovations as the null hypothesis.

19 This procedure is an extension of the seminal work of Engle and Granger (1987).

the deterministic time trends that seem to be an important part of many of the price series. We include it to illustrate the possibility of finding spurious co-integration if the data generating process is mis-specified. This test uses the residuals from the regression based on (4); the null hypothesis is that the exchange rate and the two price indices are not co-integrated. Table 9 reports the Dickey Fuller Statistic for the residuals of the regression based on (4) for the period before the EMS. The critical values for this test are given by Engle and Yoo (1987, Table 2). Table 10 presents the test statistics for the period of the EMS. We cannot reject the null of no co-integration for any regressions in either sub-sample. Many other empirical analyses of exchange rates have had difficulty in finding evidence of co-integration; it is perhaps surprising that even European exchange rates do not seem to satisfy the weak form of purchasing power parity that a co-integrative relationship entails. We see, again, that the test statistics generally seem to have slightly higher significance levels in the early period.

V. Analyses of the Real Exchange Rates within Europe

In this section we analyze whether the European real exchange rates are stationary time series. Recall that we define real exchange rates as $r \equiv e + p^* - p$, where the symbols are as above. An increase in r represents a real depreciation; this is often called a gain in competitiveness. We are interested in whether the EMS realignments of the nominal exchange rates have maintained competitiveness between European countries. Figure 3 graphs the fifteen European real exchange rates.

It is important to emphasize that the hypothesis that a real exchange rate is stationary is strictly nested within the hypothesis that the relevant nominal exchange rate and price indices are co-integrated. We describe some evidence below that several of the real exchange rates in the earlier period are stationary, but there is only weak evidence that the relevant series are cointegrated. This is may be an artifact of the testing procedure because the Engle and Yoo critical values and those of Park, Choi, and Ouliaris are quite conservative in the case of three co-integrated series. We do analyze, however, a hypothesis that is not nested within the first. We examine whether these real exchange rates can be modeled as stationary series with a single structural break occurring at the time of the formation of the EMS.

As a preliminary step in this analysis, we tested whether the real exchange rates were themselves stationary over the whole sample and in each sub-sample. We did this in exactly the same way that we analyzed the nominal exchange rates in Section III. We could reject the null of a unit root for only one real exchange rate over the entire sample: that of Germany-Italy. We found also that we could reject the null of a unit root in only five cases in the earlier period; these were for the real exchange rates involving these pairs of countries: Belgium-Germany, Belgium-Italy, Germany-Italy, Italy-Netherlands, and Italy-United Kingdom. In the later period, we rejected the null of a unit root for the real exchange rate of this country pair only: Germany-Netherlands. This is no surprise because the EMS realignments have treated these two countries' currencies similarly, and they are the low inflation countries of Europe. Again for the sake of brevity, we do not report all forty-five statistics.

As the last step in the analysis, we are interested in determining whether the advent of the EMS actually represents a structural break in the data generating process for these real exchange rates. The test we present below was developed by Park (1988), and Perron (1987) first noticed the importance of distinguishing between time series with unit roots and stationary time series that exhibit a structural break. Now let t_b represent the period at the end of which there is a structural break in a time series under investigation. Further, define

$$d = \begin{cases} 0 & \text{if } t \leq t_b \\ 1 & \text{otherwise} \end{cases}$$

and

$$s = \begin{cases} 0 & \text{if } t \leq t_b \\ t - t_b & \text{otherwise.} \end{cases}$$

Consider this model

$$x_t = \beta_0 + \beta_1 t + \beta_2 d + \beta_3 s + \beta_4 x_{t-1} + u_t \tag{5}$$

where the innovation sequence $\{u_t\}$ satisfies the Assumption in Section II. The interpretation of the parameters β_2 and β_3 is that they model respectively a shift in the intercept and a shift in the deterministic time trend describing the series. If we consider equation (5) as a model of a real exchange rate, then β_2 describes a change in the rate of depreciation at the advent of a new regime and β_3 describes a shift in the acceleration of this depreciation.

We used Park's H(1,1) statistic to test this null hypothesis H_o: $\beta_4 = 1$ against the alternative that $|\beta_4| < 1$. The critical values for this statistic are given in Park (1988, Appendix B). We report the results from this test in table 11. We reject the null in six of the fifteen cases, and the test statistic is marginally significant in three other cases. This is a reassuring finding for proponents of the notion of purchasing power parity.

It implies that several of these real exchange rates may actually have been stationary over the entire period of floating exchange rates if one allows for a structural change at the advent of the EMS. The evidence suggests that there are several cases where the behavior of real exchange rates in relationship to the pound sterling have undergone a structural change at the time of the advent of the EMS.[20] Further, there is strong evidence that lira-mark and guilder-lira real rates have undergone structural breaks. This lends credence to the notion that the EMS has served to "discipline" Italy by allowing for a real appreciation against the low inflation European countries in this decade.

We conclude the analysis of these real exchange rates by testing this hypothesis directly. Consider the following model of a real exchange rate

$$r_t = \beta_0 + \beta_1 t + \beta_2 s + u_t \tag{6}$$

where r_t is the logarithm of the real exchange rate and the other variables are as above. Equation (6) models the notion that the rate of depreciation of the real exchange rate is systematically different after the advent of the EMS. If we assume that the innovation process $\{u_t\}$ is integrated of order one, then it is natural to consider the first differences of (6) given by

$$\Delta r_t = \beta_0 + \beta_1 d + \Delta u_t \tag{7}$$

where $\Delta r_t = r_t - r_{t-1}$, Δu_t is analogous, and the dummy variable is as defined above. Of course, the process $\{\Delta u_t\}$ is integrated of order zero, but we still

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20 The Bank of England does not participate in the EMS intervention mechanism; since the pound sterling is used to define the ecu, the European central banks often define the par value of this currency so that there is less apparent appreciation of the mark and guilder against the other European currencies.

must assume that this process is not necessarily independently and identically distributed. For equation (7), we are interested in this null hypothesis H_o: $\beta_1 = 0$ against the general alternative. Park and Choi propose a test statistic for models like (7) that is a modification of the Wald statistic for this hypothesis; this modification takes into account the ratio of the short-run and long-run variances in the innovation process. This statistic is distributed $\chi^2(1)$ for the test based on (7). We report the test statistics in table 12. We reject the null in two cases: those involving the real exchange rates of Belgium-Italy and Italy-Netherlands. In both cases, Italy's real exchange rate appreciated against these countries. This is evidence that Italy actually did suffer a loss of competitiveness against Belgium and the Netherlands under the EMS.

VI. Conclusion

This paper has explored two themes. The first has had to do with an assessment of the European Monetary System, and the second has been an illustration of the applications of new econometric techniques to the old issue of purchasing power parity. We summarize our findings by discussing each of these themes separately.

Our first theme speaks to international economists interested in the European monetary affairs. We have presented very strong evidence that the realignments of the European Monetary System have not served to maintain competitiveness between its largest member countries. We have presented weaker evidence that there was a higher likelihood of co-integration between exchange rates and prices before the EMS than after it. More importantly, the evidence on real exchange rates shows that before the EMS it was more common for countries to realign their nominal exchange rates to offset inflation differentials. These observations lead us to conclude that there may be an asymmetry in the realignments of the System. Italy, the highest inflation country in the System, does seem to have lost competitiveness during the EMS period. Moreover, there seems to be a structural break in the behavior of several of the real exchange rates involving Italy. This may indicate that the EMS has treated the lira differently from the way the Snake treated it.

Our second theme is of practical importance for both applied international economists and for applied econometricians. We would like to emphasize three points that should serve to guide further applied research in this field. First, floating exchange rates are asset prices; there are good theoretical and empirical reasons to model them as stochastic processes, integrated of order one, whose innovations are heterogeneously distributed. Second, the price indices of many industrial countries in the last two decades exhibit marked declines in their rates of inflation; it is therefore necessary to include a quadratic time trend in any co-integrating regression if one hopes not to mis-specify a long-run relationship between prices and other macroeconomic variables. Third, we have provided the first application in international economics of a unit root test that allows for a structural break. Econometricians studying changes in monetary regimes, such as the advent of a hyper-inflation, would do well to apply this kind of a test in their own research.

Table 1
The Nominal Exchange Rate Before the EMS

$Z_p(t(\hat{\alpha}))$ Statistic

	Fr	Ge	It	Ne	UK
Be	-2.005	-3.911**	-2.996	-3.567**	-2.178
Fr	--	-2.069	-2.925	-2.144	-1.690
Ge	--	--	-3.016	-2.558	-2.421
It	--	--	--	-3.143	-2.807
Ne	--	--	--	--	-2.196

Notes to table:
The sample period is from April 1973 to February 1979 (71 observations). The critical values are given in Fuller (1976, Table 8.5.2); they are -3.18 at the 10% significance level and -3.50 at the 5% significance level.
* indicates significant at the 10% level
** indicates significant at 5% level

Table 2
The Nominal Exchange Rate During the EMS

$Z_p(t(\hat{\alpha}))$ Statistic

	Fr	Ge	It	Ne	UK
Be	-3.441*	-0.886	-2.181	-1.113	-2.309
Fr	--	-1.216	-2.159	-1.582	-2.357
Ge	--	--	-2.161	-3.267*	-2.734
It	--	--	--	-2.701	-2.405
Ne	--	--	--	--	-2.784

Notes to table:
The sample period is from March 1979 to May 1988 (111 observations). The critical values are given in Fuller (1976, Table 8.5.2); they are -3.15 at the 10% significance level and -3.45 at the 5% significance level. See the notes to table 1 for an explanation of the symbols.

Table 3
Price Indices

$Z_p(t(\hat{\alpha}))$ Statistic

	Be	Fr	Ge	It	Ne	UK
pre-EMS	0.314	-1.269	-1.473	-1.558	0.748	-0.327
EMS	1.554	-0.091	-0.841	-0.366	-0.604	-4.153**

Notes to table:
The pre-EMS period is from April 1973 to February 1979 (71 observations). The EMS period is from March 1979 to May 1988 (111 observations). The critical values are as in tables 1 and 2 for the relevant sample sizes. See the notes to table 1 for an explanation of the symbols.

Table 4
The Nominal Exchange Rate Before the EMS

$Z_p(\hat{\alpha},\hat{\beta}_p)$ Statistic

	Fr	Ge	It	Ne	UK
Be	3.208	9.431**	5.440	4.666	2.674
Fr	--	3.258	5.004	3.542	1.888
Ge	--	--	5.910*	3.742	3.769
It	--	--	--	5.958*	5.441
Ne	--	--	--	--	2.981

Notes to table:
The sample period is from April 1973 to February 1979 (71 observations). The critical values are given in Dickey and Fuller (1981, Table VI); they are 5.61 at the 10% significance level and 6.73 at the 5% significance level. See the notes to table 1 for an explanation of the symbols.

Table 5
The Nominal Exchange Rate During the EMS

$Z_p(\hat{\alpha},\hat{\beta}_p)$ Statistic

	Fr	Ge	It	Ne	UK
Be	6.884**	1.315	2.630	1.348	4.278
Fr	--	0.913	2.876	1.509	4.029
Ge	--	--	1.408	7.358**	5.202
It	--	--	--	3.540	4.217
Ne	--	--	--	--	5.607*

Notes to table:
The sample period is from March 1979 to May 1988 (111 observations). The critical values are given in Dickey and Fuller (1981, Table VI); they are 5.47 at the 10% significance level and 6.49 at the 5% significance level. See the notes to table 1 for an explanation of the symbols.

Table 6
Price Indices

$Z_p(\hat{\alpha},\hat{\beta}_p)$ Statistic

	Be	Fr	Ge	It	Ne	UK
pre-EMS	6.345*	4.271	4.973	2.450	8.593**	1.535
EMS	14.253**	56.913**	16.511**	50.109**	24.843**	26.779**

Notes to table:
The pre-EMS period is from April 1973 to February 1979 (71 observations). The EMS period is from March 1979 to May 1988 (111 observations). The critical values are as in tables 4 and 5 for the relevant sample sizes. See the notes to table 1 for an explanation of the symbols.

Table 7
Co-integration Tests for the Period Before the EMS

J(2,6) Statistic

	Fr	Ge	It	Ne	UK
Be	1.878	0.317	0.529	0.243*	0.784
Fr	--	1.970	1.267	2.039	1.278
Ge	--	--	0.832	0.496	0.813
It	--	--	--	0.741	0.066**
Ne	--	--	--	--	0.771

Notes to table:
The sample period is from March 1973 through February 1979 (72 observations). The critical value at the 10% level 0.254 is and that at the 5% level is 0.163. These values were estimated by Monte Carlo methods using 5000 replications. One rejects the null hypothesis of no co-integration for low values of the statistic. See the notes to table 1 for an explanation of the symbols.

Table 8
Co-integration Tests for the Period of the EMS

J(2,6) Statistic

	Fr	Ge	It	Ne	UK
Be	0.149**	1.794	0.938	1.648	0.584
Fr	--	5.630	2.561	3.913	0.371
Ge	--	--	0.522	0.443	0.609
It	--	--	--	0.503	0.478
Ne	--	--	--	--	0.601

Notes to table:
The sample period is from March 1979 to May 1988 (111 observations)
The critical values at the 10% level is 0.237 and that at the 5% level is 0.155. These values were again estimated by Monte Carlo methods with 5000 replications. See the notes to table 1 for an explanation of the symbols.

Table 9
Residual-Based Co-integration Tests for the Period Before the EMS

Dickey-Fuller Statistic

	Fr	Ge	It	Ne	UK
Be	-1.672	-3.125	-3.150	-3.527	-2.456
Fr	--	-1.971	-2.757	-1.791	-2.695
Ge	--	--	-3.051	-2.457	-2.609
It	--	--	--	-3.176	-3.292
Ne	--	--	--	--	-2.486

Notes to table:
The sample period is from March 1973 through February 1979 (72 observations). The critical values are given in Engle and Yoo (1987, Table 2); they are -3.73 at the 10% level and -4.11 at the 5% level.

Table 10
Residual-Based Co-integration Tests for the Period of the EMS

Dickey-Fuller Statistic

	Fr	Ge	It	Ne	UK
Be	-2.183	-1.758	-1.786	-1.758	-1.544
Fr	--	-2.148	-1.945	-2.507	-1.987
Ge	--	--	-1.716	-3.073	-3.205
It	--	--	--	-3.313	-1.927
Ne	--	--	--	--	-2.811

Notes to table:
The sample period is from March 1979 to May 1988 (111 observations). The critical values are given in Engle and Yoo (1987, Table2); they are -3.59 at the 10% and -3.93 at the 5% level.

Table 11
Tests for Structural Change In the Real Exchange Rates

H(1,1) Statistic

	Fr	Ge	It	Ne	UK
Be	-3.234	-2.386	-3.466	-2.683	-3.611
Fr	--	-3.144	-3.778	-3.387	-3.918*
Ge	--	--	-4.775**	-3.512	-3.973*
It	--	--	--	-4.800**	-4.288**
Ne	--	--	--	--	-4.070*

Notes to table:
The sample period is from April 1973 to May 1988 (182 observations); the time of the structural break is March 1979 (the 72nd observation). The critical values are given in Park (1988, Appendix B); they are -3.87 at the 10% level and -4.15 at the 5% level. See the notes to table 1 for an explanation of the symbols.

Table 12
Tests for Change in Rate of Real Depreciation

G_{Δ} Statistic

	Fr	Ge	It	Ne	UK
Be	1.362	1.378	4.445**	0.031	0.894
Fr	--	0.109	1.547	1.129	0.108
Ge	--	--	1.634	2.024	0.235
It	--	--	--	4.133**	0.178
Ne	--	--	--	--	0.777

Notes to table:
The sample period is from April 1973 to May 1988 (182 observations); the time of the structural break is March 1979 (the 72nd observation). This is a $\chi^2(1)$ test; the critical values are 2.71 at the 10% level and 3.84 at the 5% level. See the notes to table 1 for an explanation of the symbols.

Appendix

Dates of Realignments of the European Monetary System

13 March 1979
24 September 1979
30 November 1979
23 March 1981
5 October 1981
22 February 1982
14 June 1982
21 March 1983
22 July 1985
7 April 1986
4 August 1986
12 January 1987

The Data

The six currencies examined are end of month exchange rates as maintained by the Federal Reserve Board. They are the currencies' rates at noon on the last working day of the month in New York. The implicit bilateral rates e^{ij} have been defined so that the rates are quoted in units of currency j necessary to obtain one unit of currency i; of course, an increase in e^{ij} is a weakening of currency j relative to currency i. The consumer price indices are raw monthly data from national sources and generally represent a sample of consumer goods taken in geographically dispersed cities in each of the six countries.

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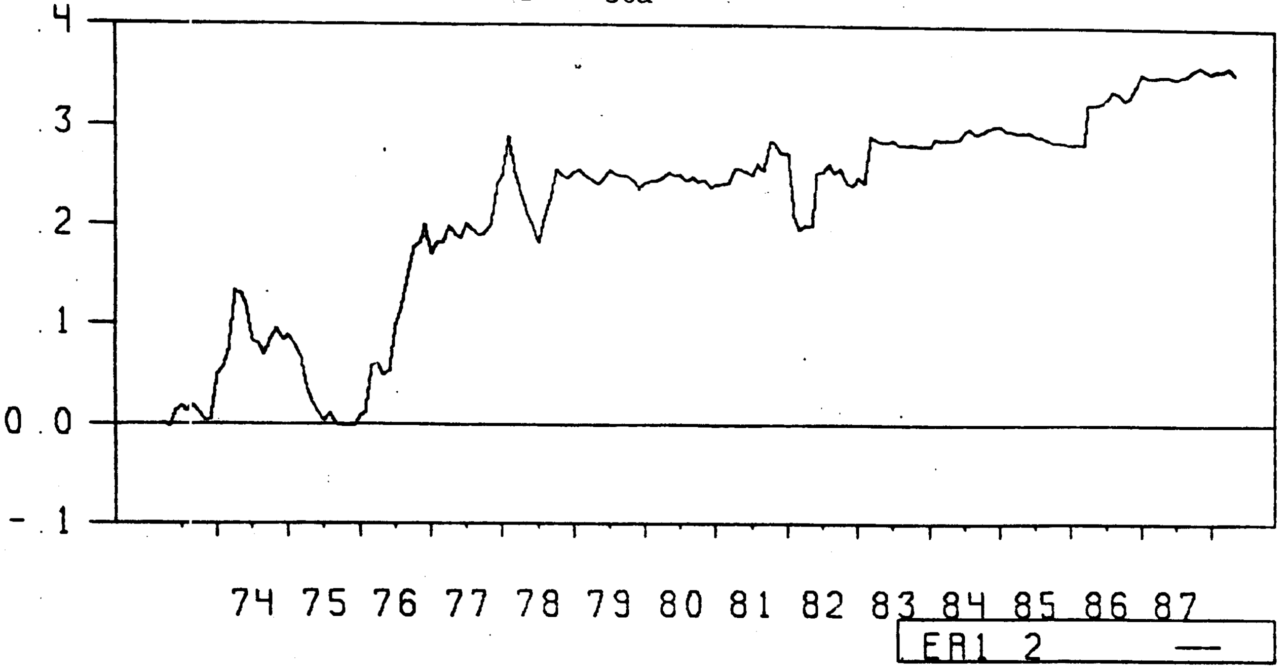
Figure 1: Logarithms of the Nominal Exchange Rates

The next five pages present graphs of the fifteen bilateral nominal exchange rate indices; the series have been normalized so that their initial values are zero. The set of countries is given by {Belgium, France, Germany, Italy, Netherlands, United Kingdom} and it is identified with {1,2,3,4,5,6}. A currency eri_j is quoted in units of currency j necessary to obtain a unit of currency i. Hence an upward movement in the graph represents a devaluation of the j-th currency.

30a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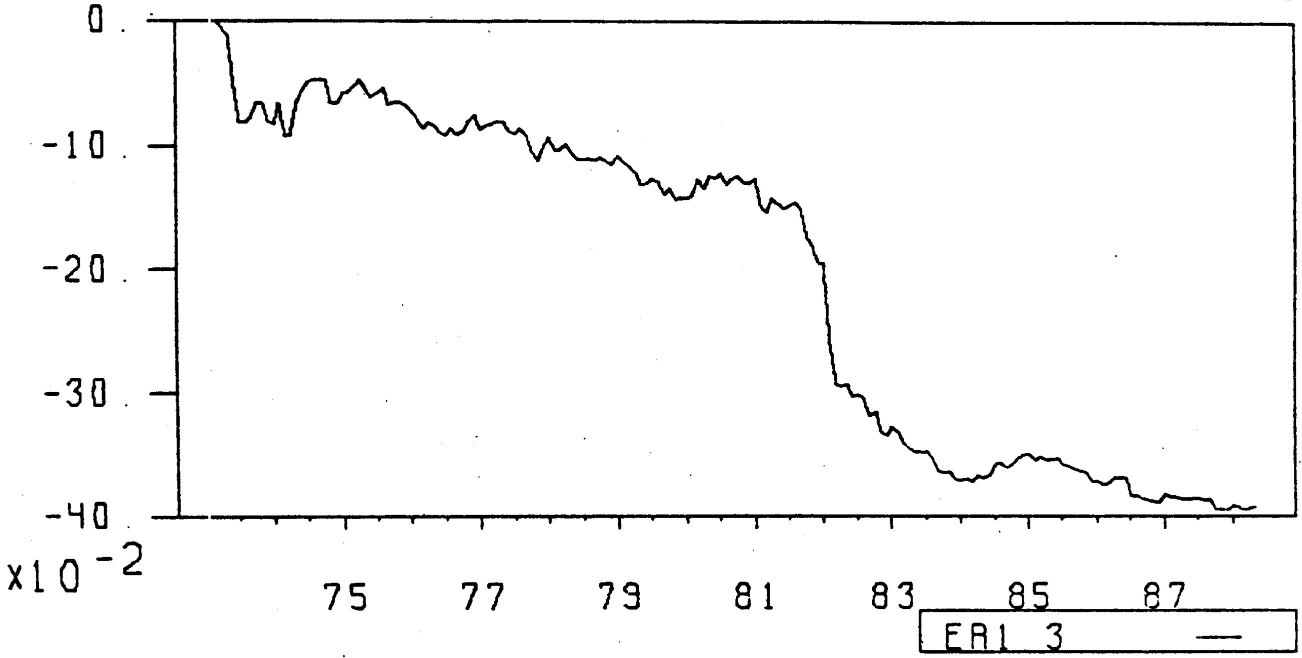
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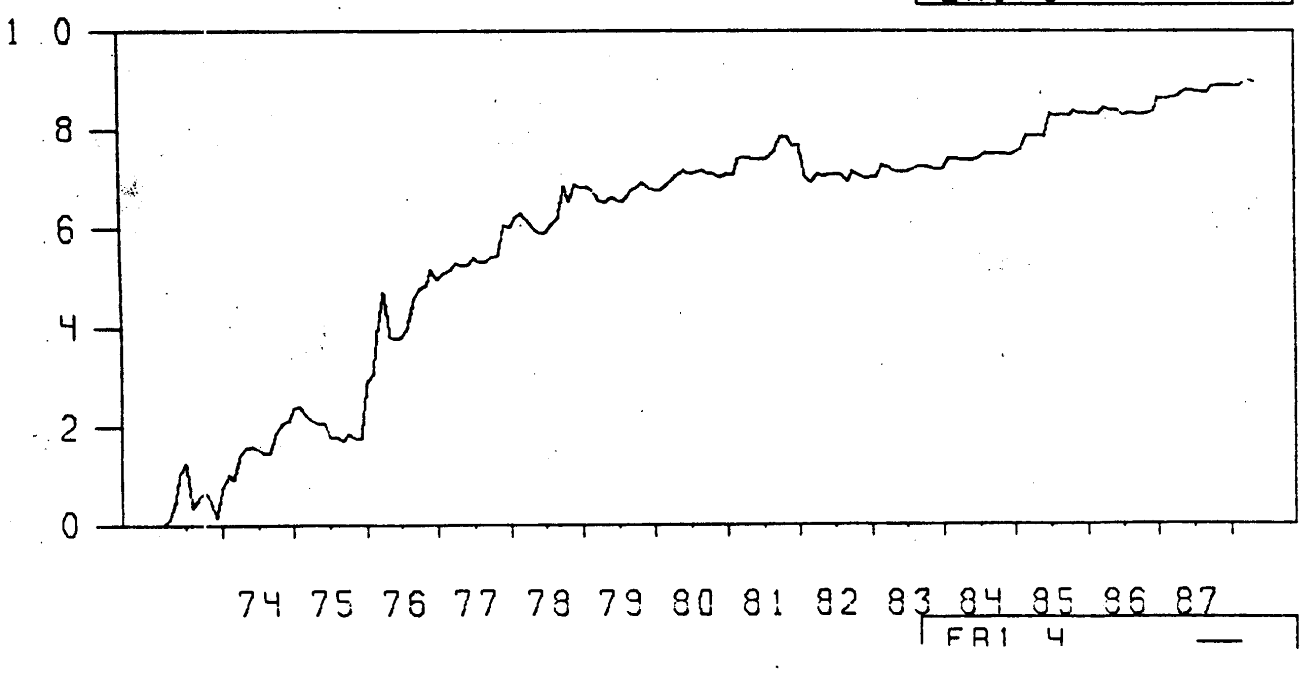
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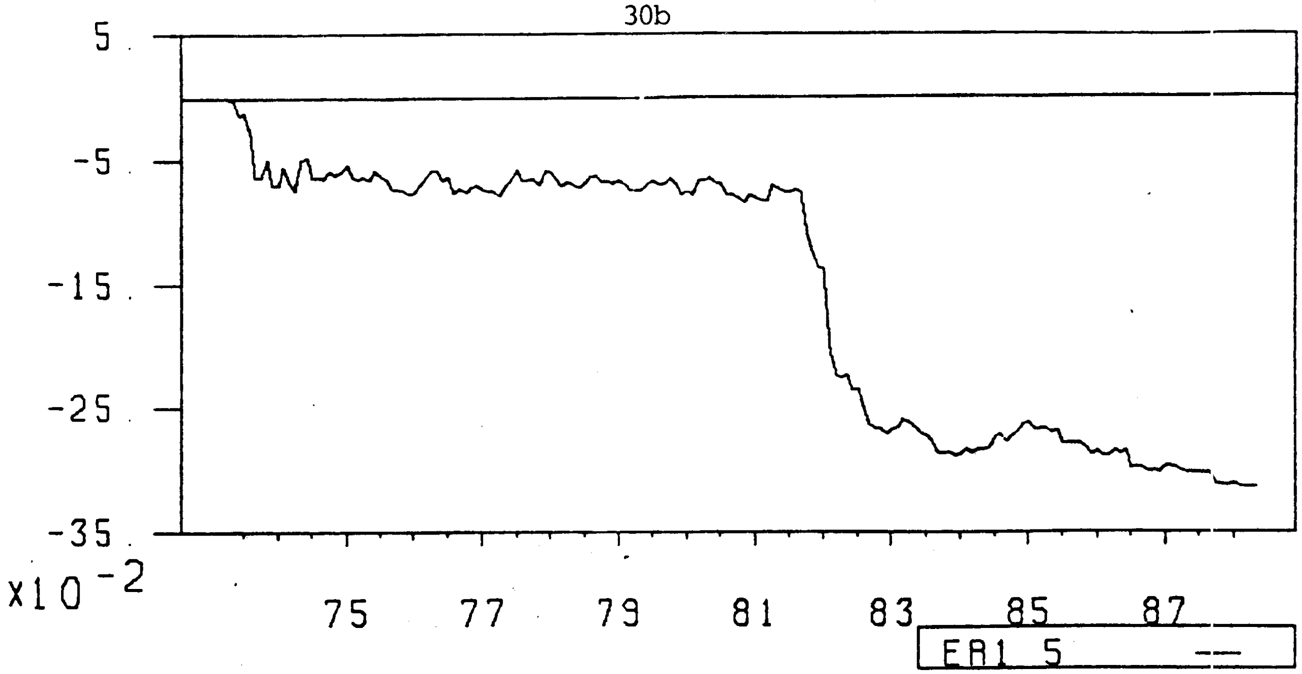
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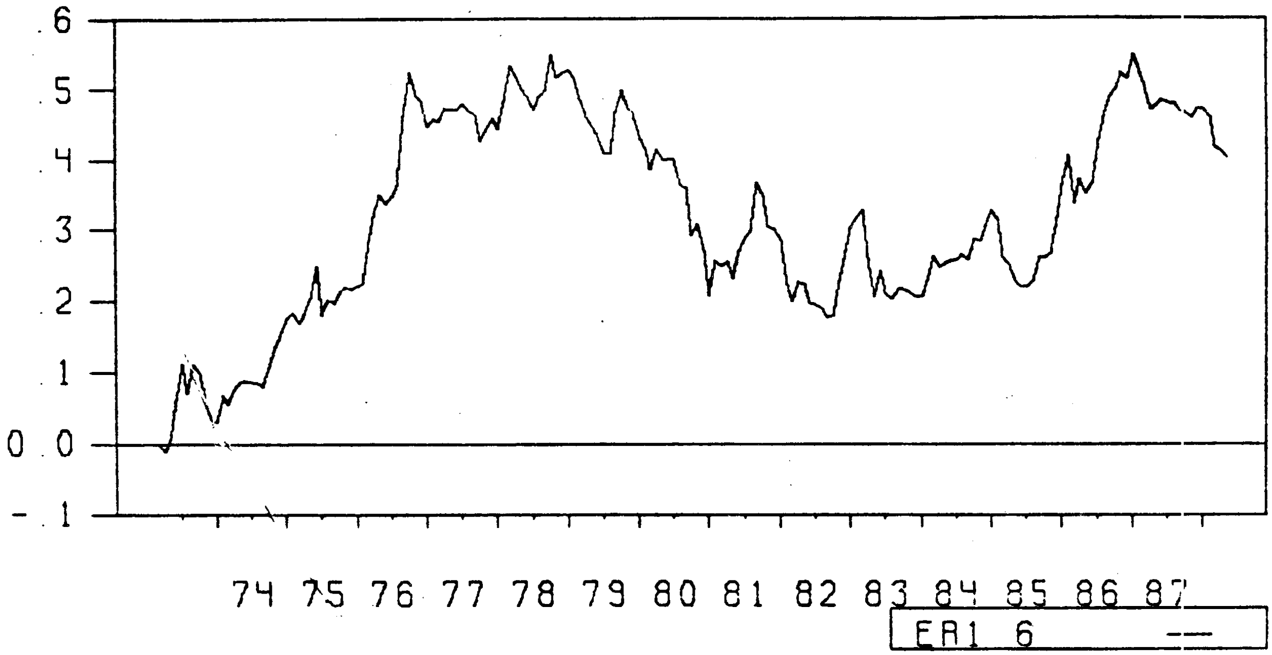
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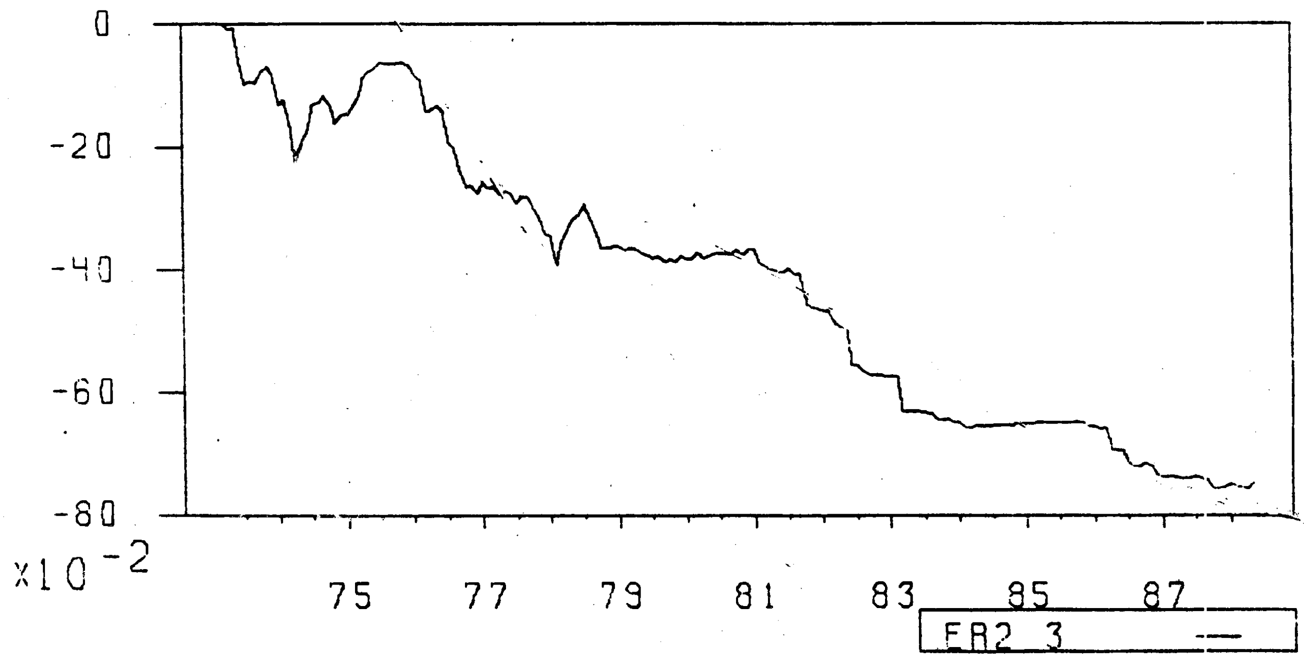
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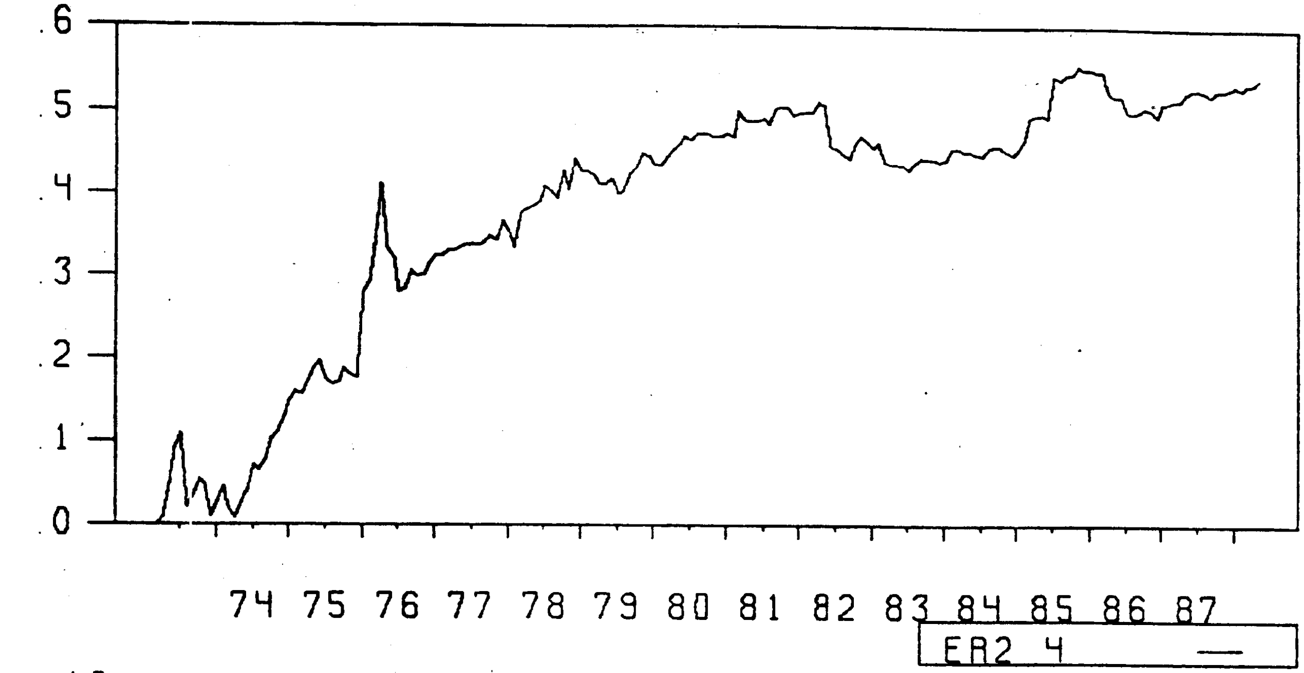
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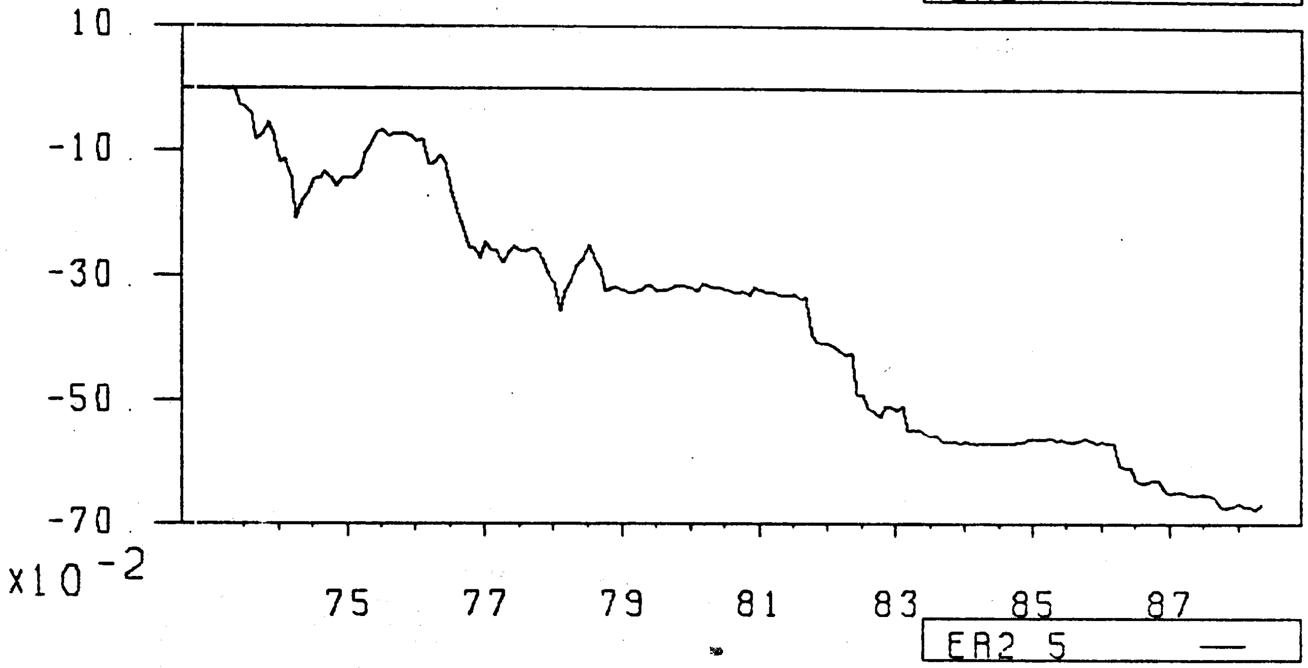
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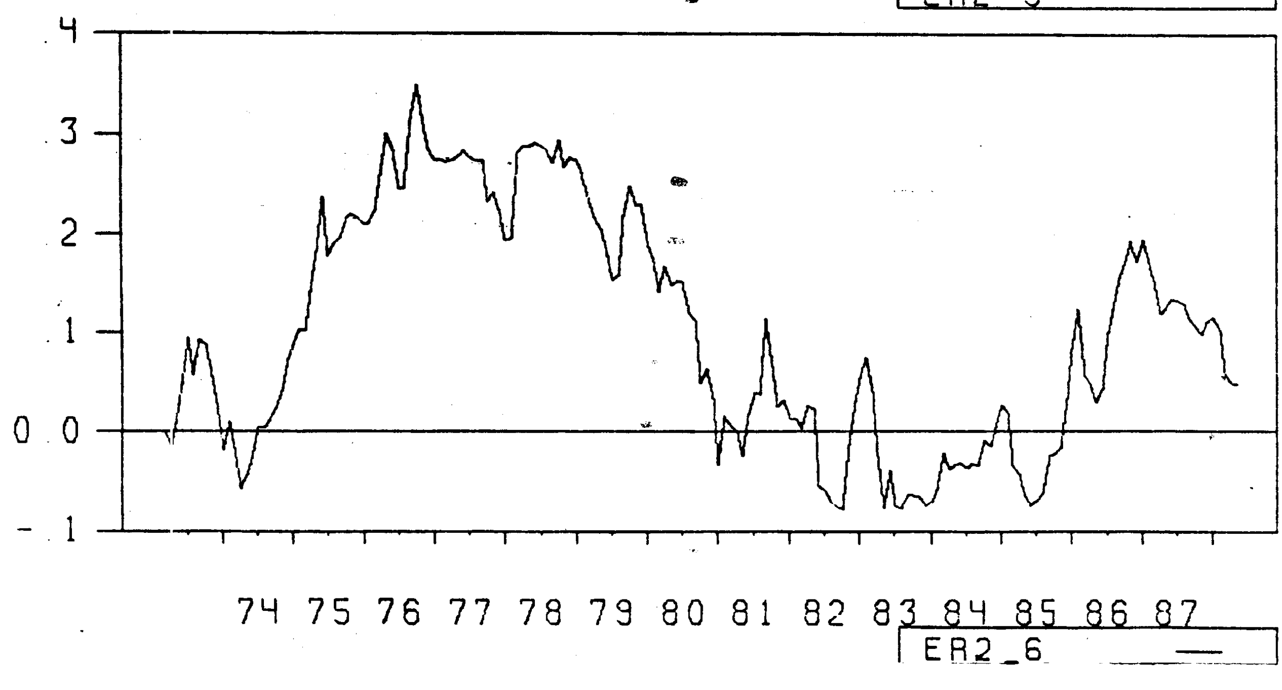
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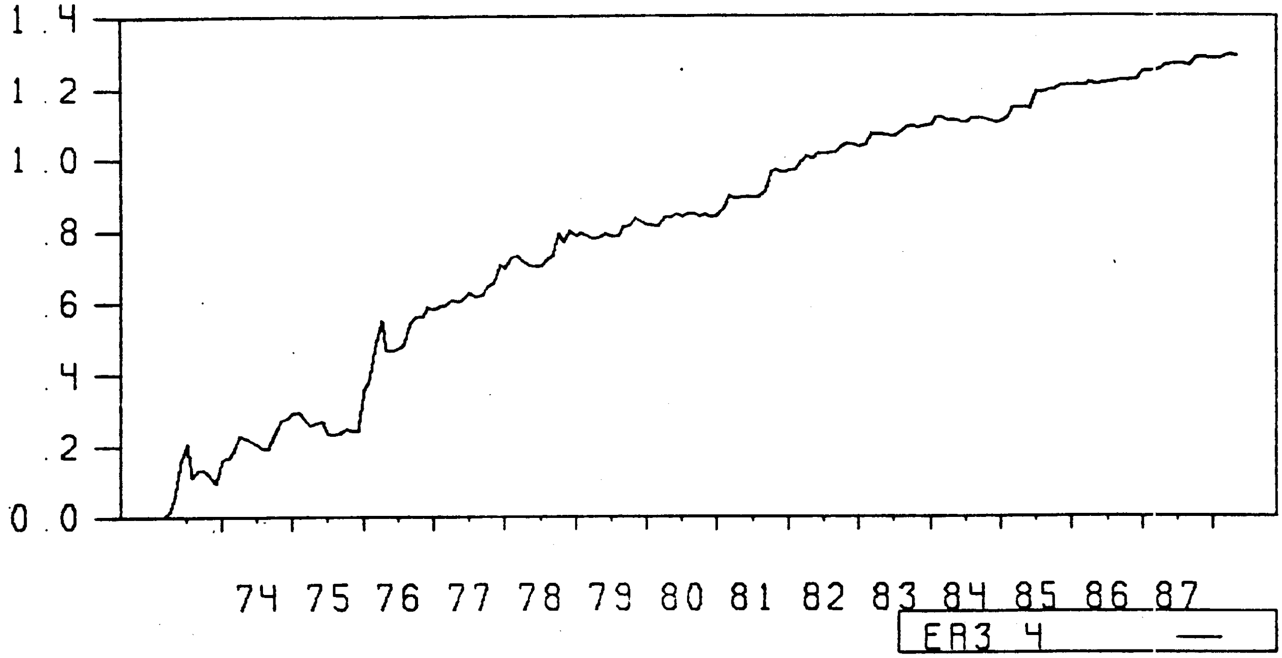
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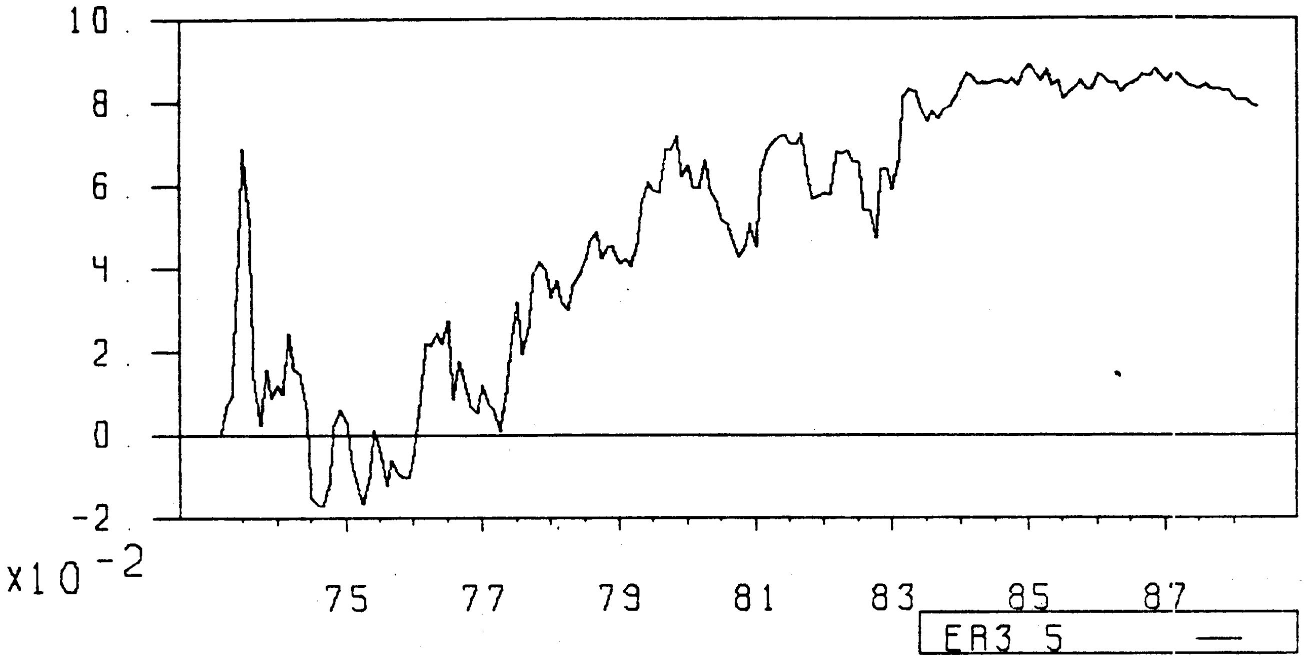
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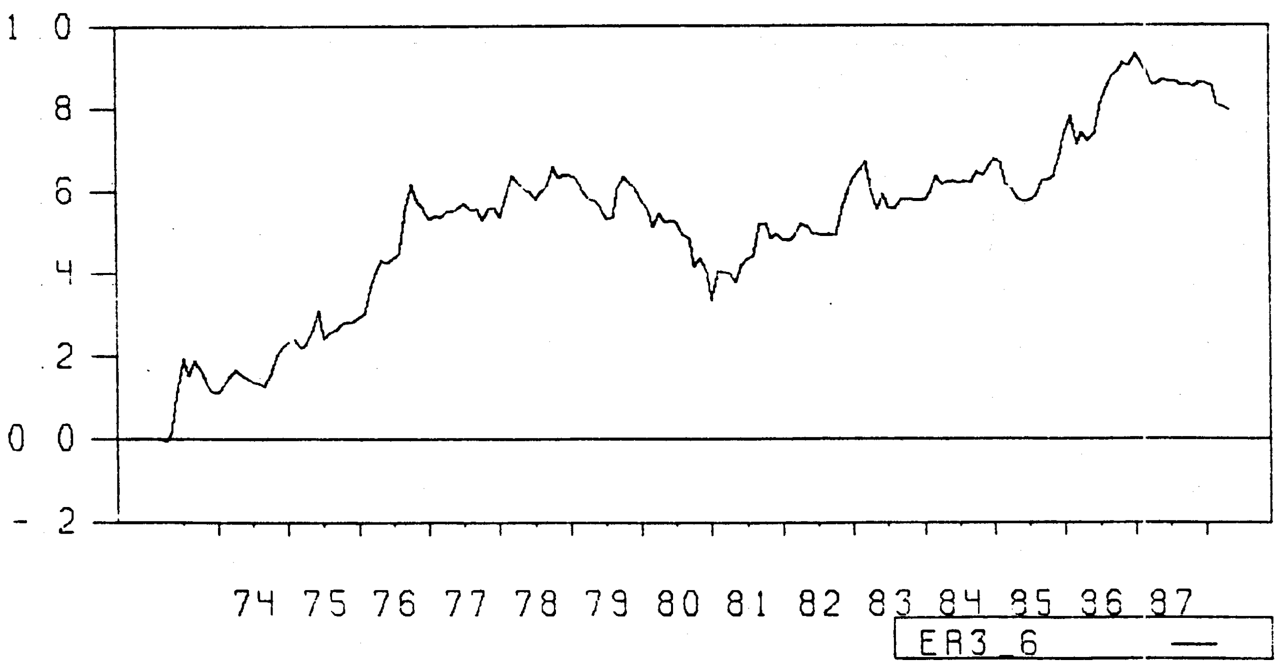
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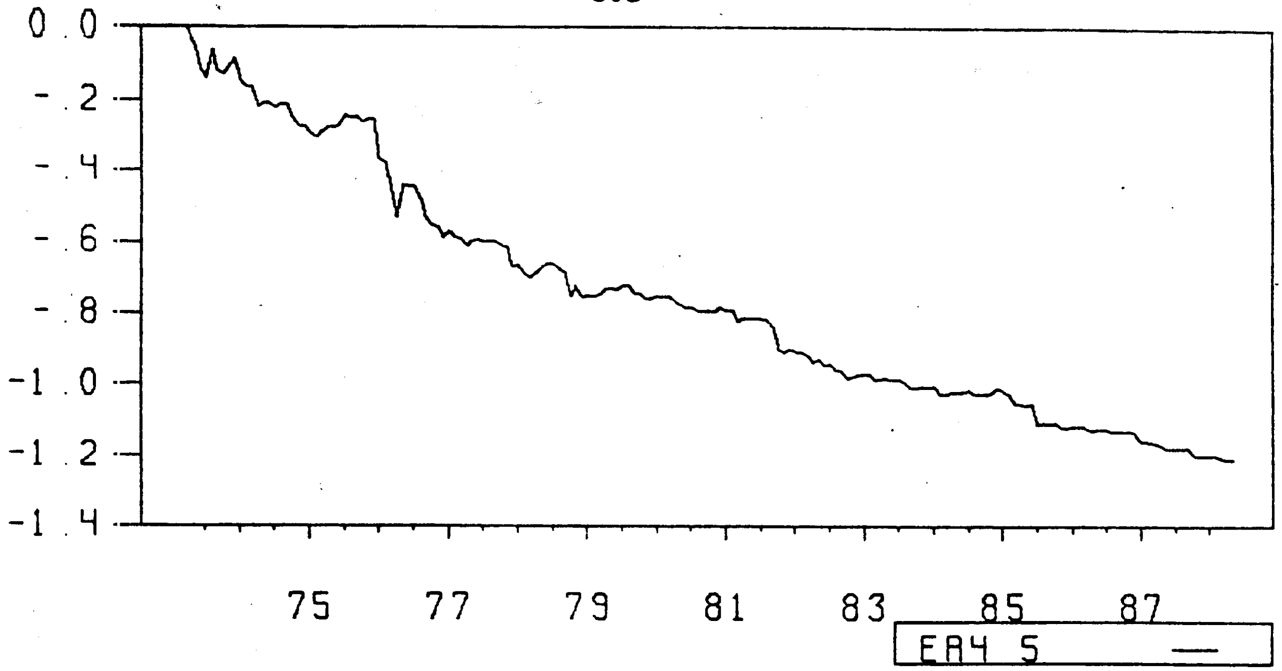
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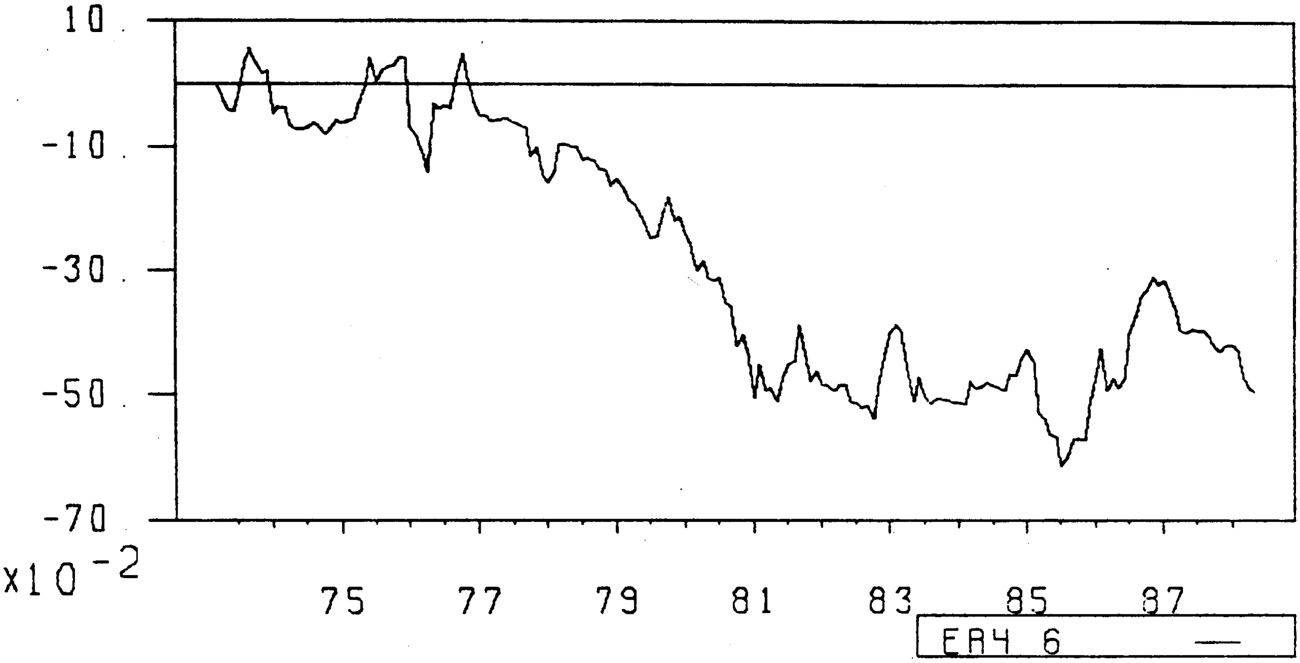
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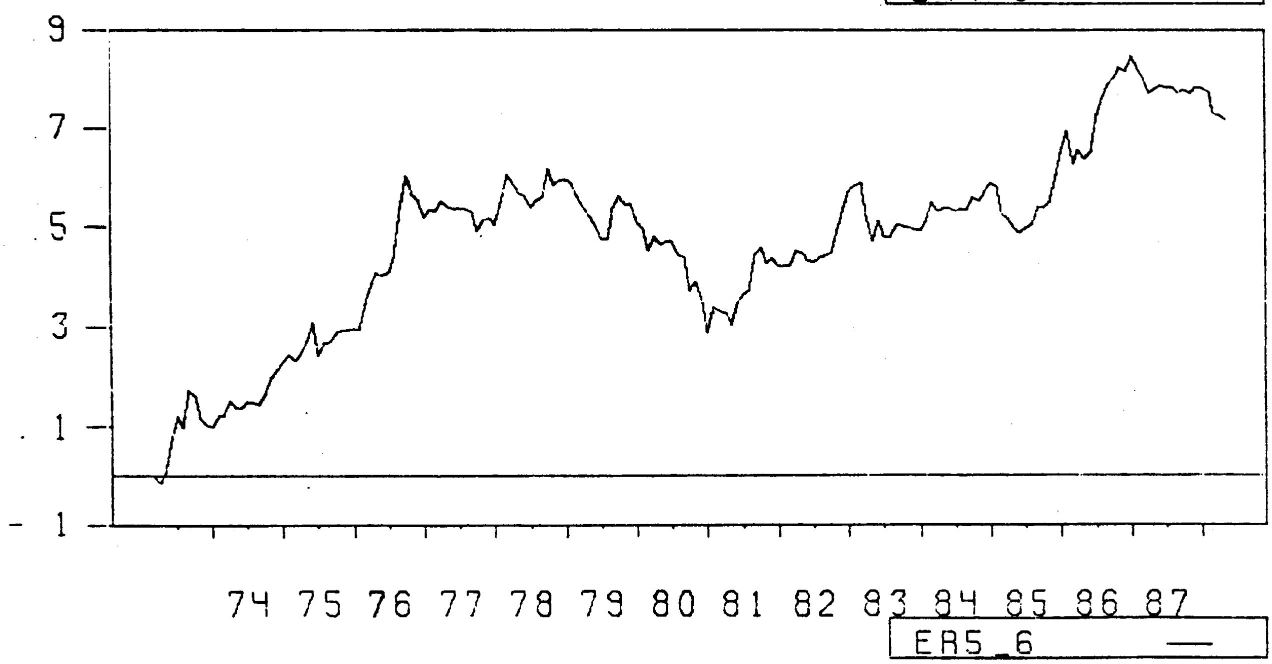
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Logarithms of the Consumer Price Indices

The next page presents a graph of the consumer price indices. They have been normalized so that their initial values are zero.

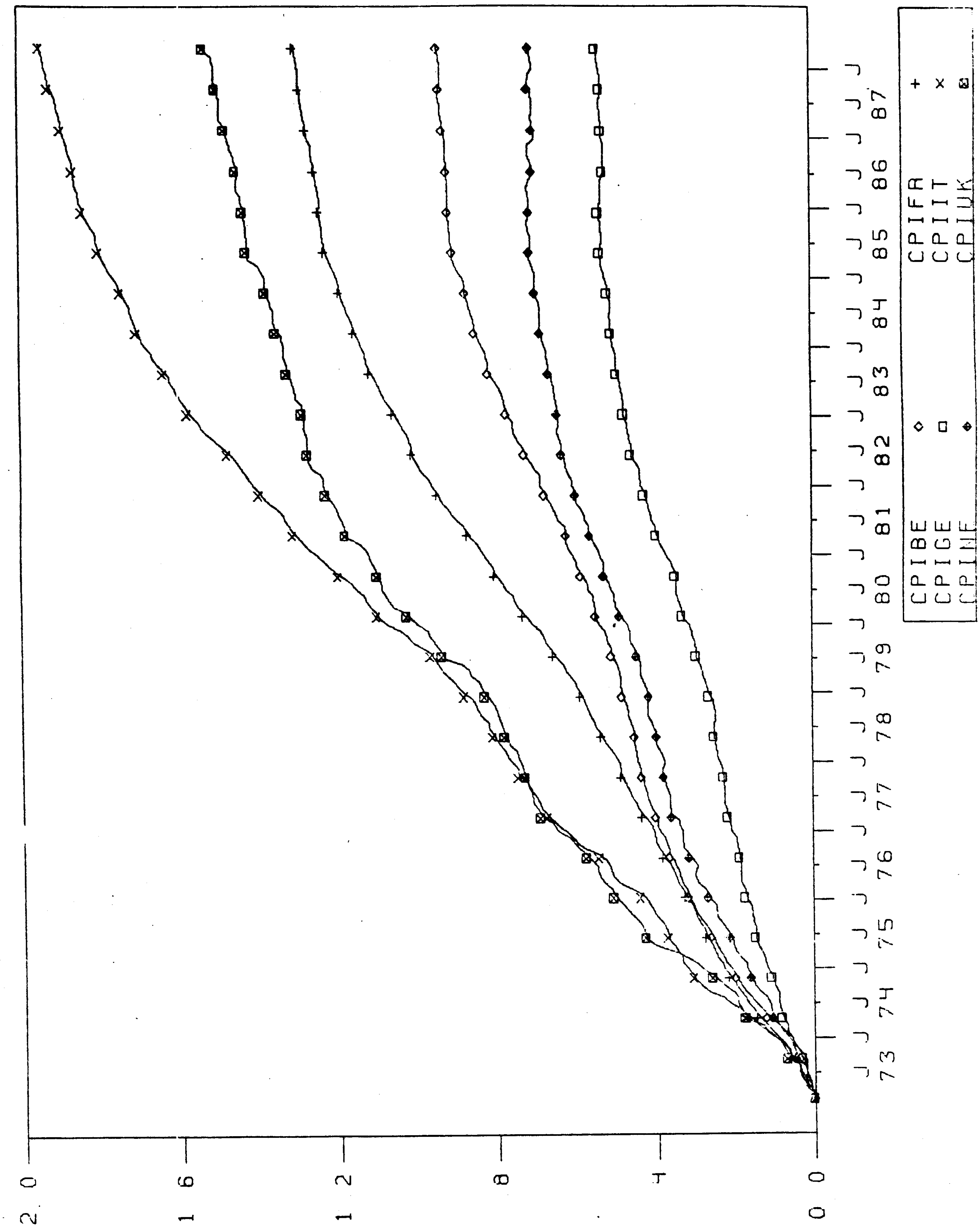

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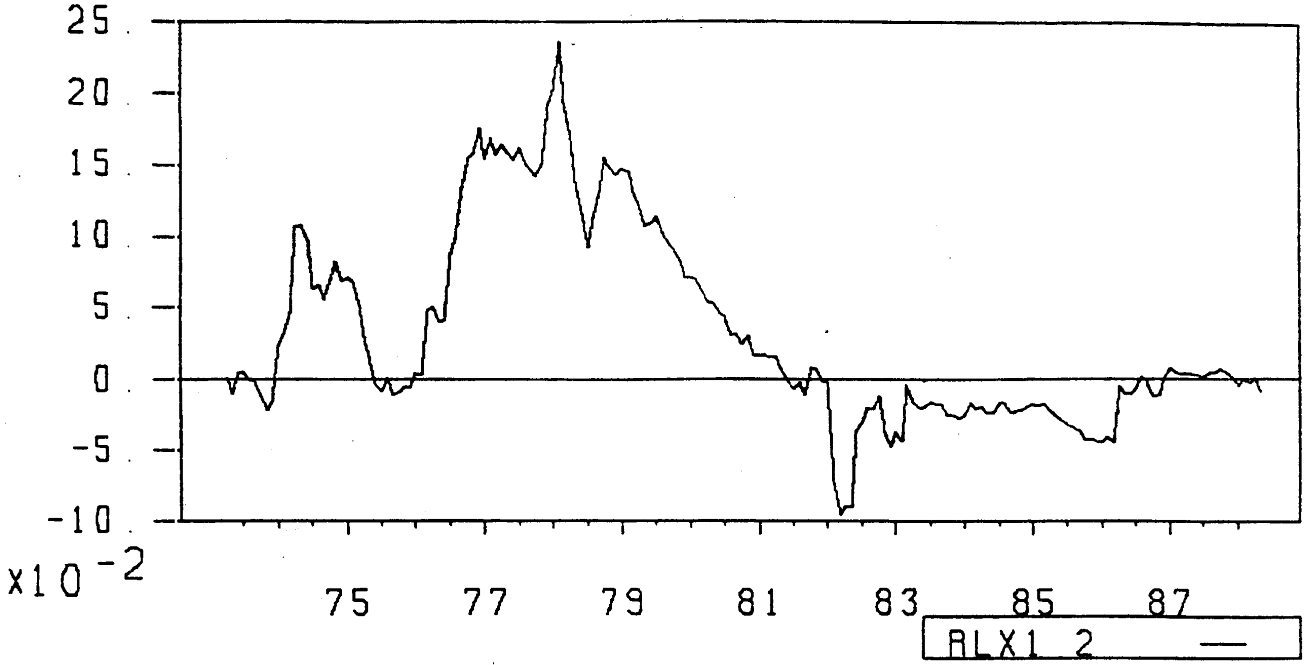
Figure 3: Logarithms of the Real Exchange Rates

The next five pages present graphs of the fifteen bilateral real exchange rate indices; the series have been normalized so that their initial values are zero. The set of countries is given by {Belgium, France, Germany, Italy, Netherlands, United Kingdom} and it is identified with {1,2,3,4,5,6}. A currency rlxi_j is quoted in units of currency j necessary to obtain a unit of currency i. Hence an upward movement in the graph represents a real devaluation of the j-th currency.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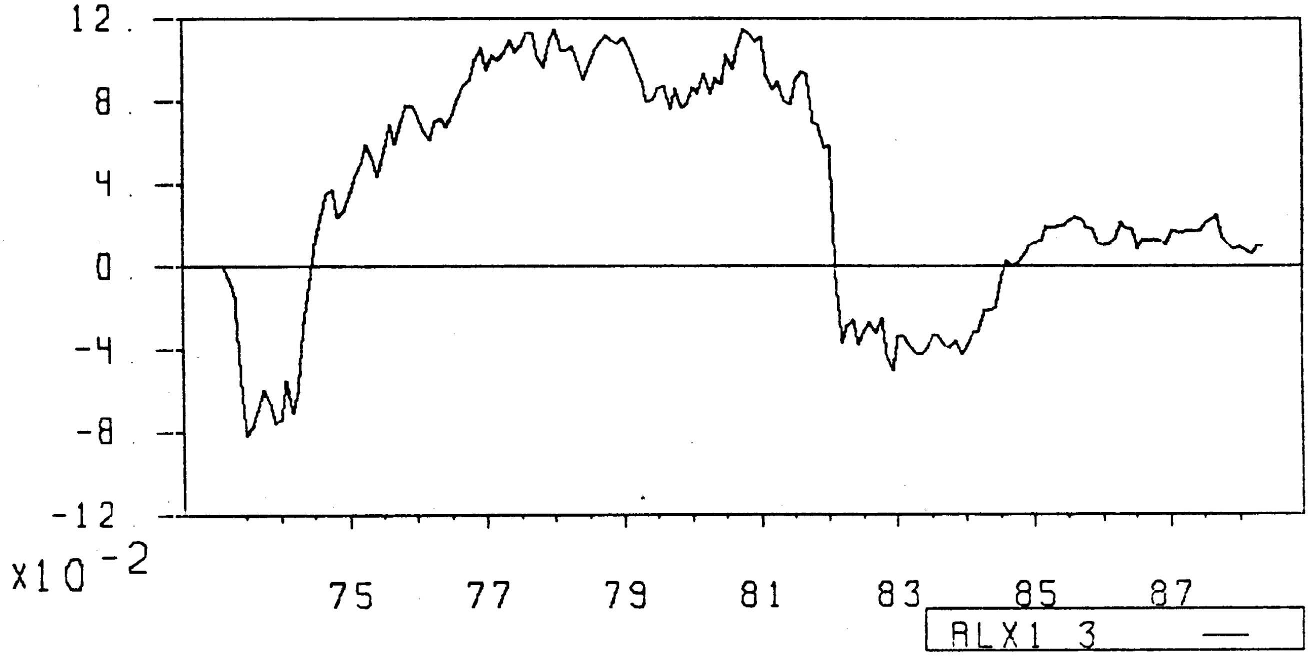
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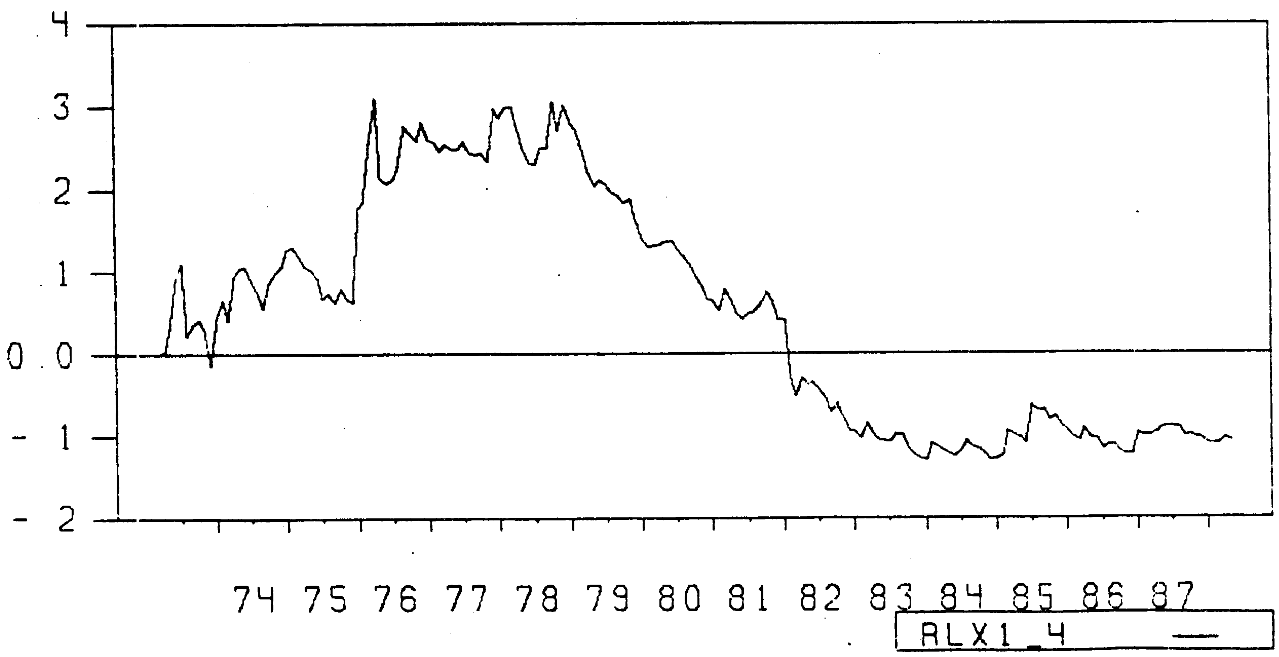
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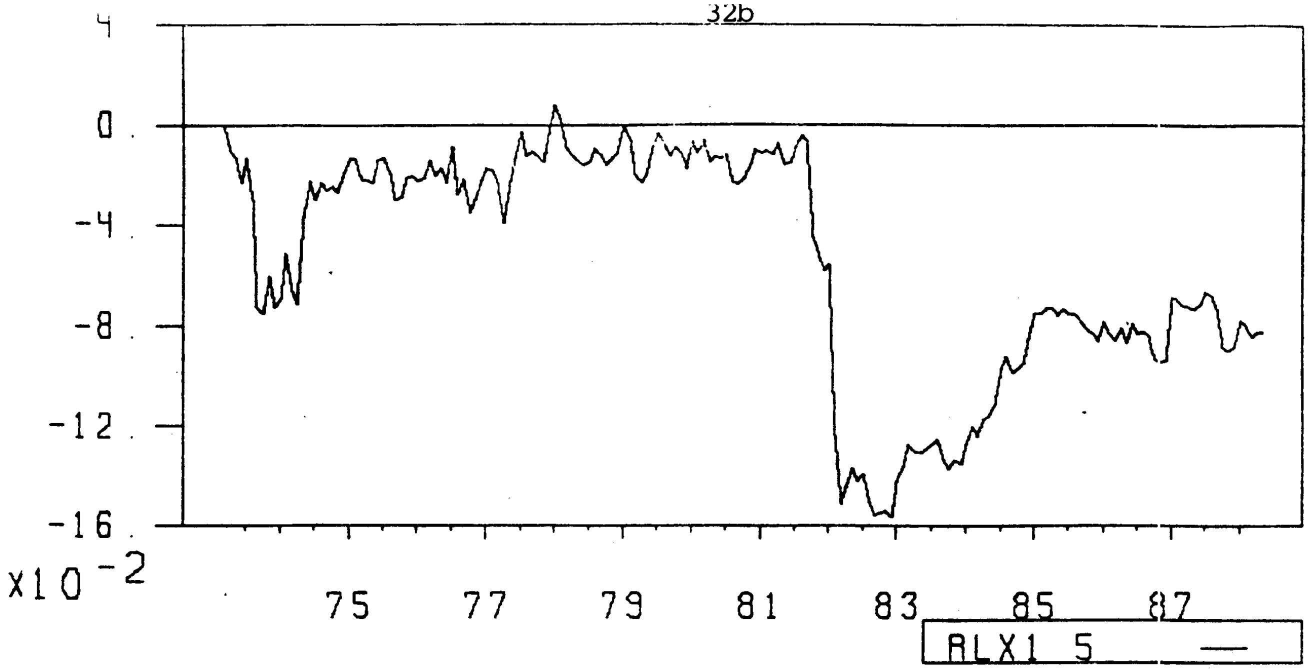

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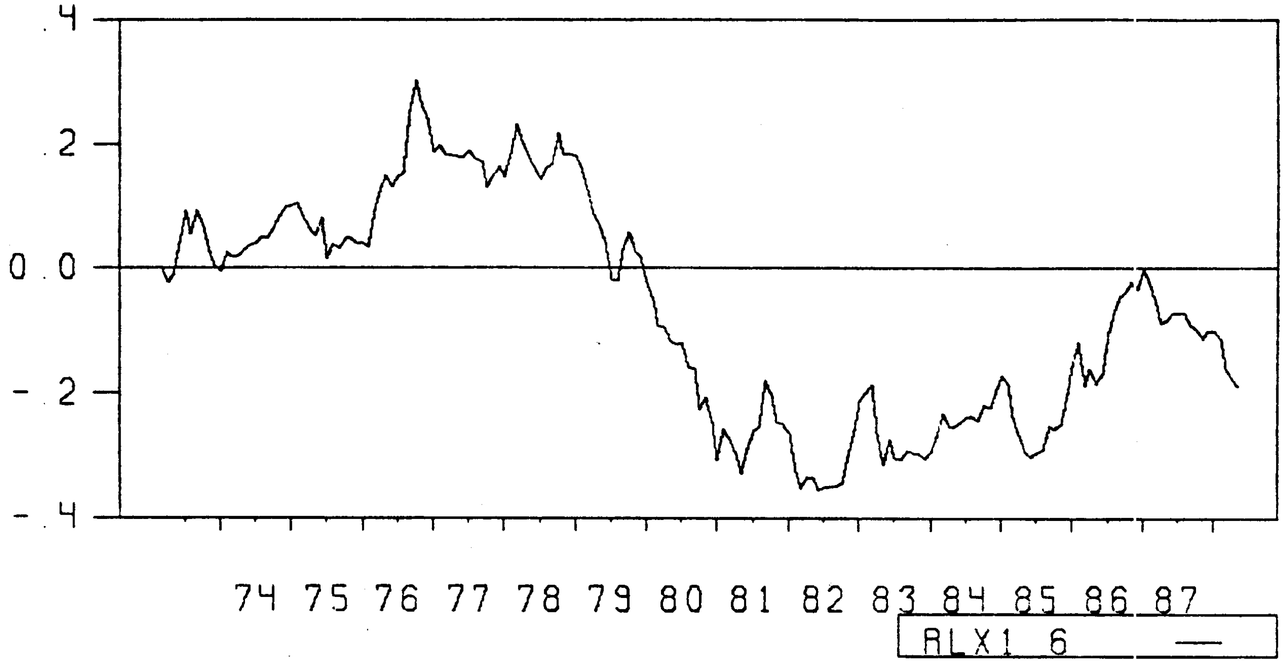

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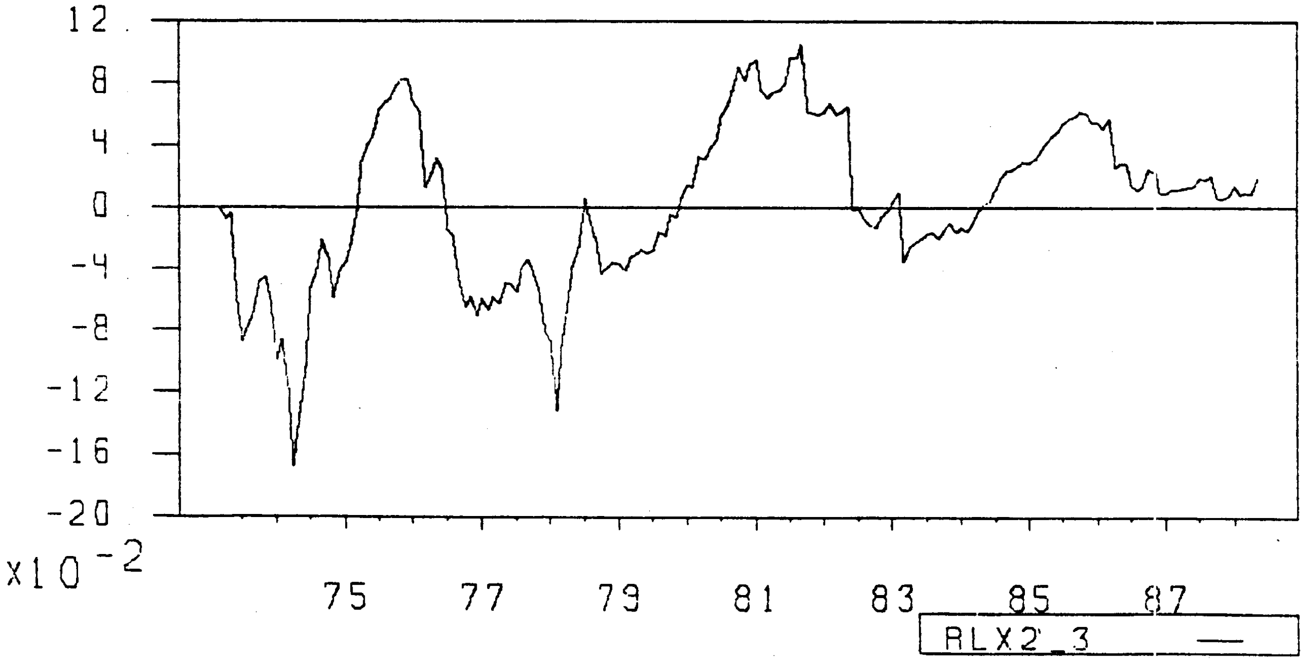

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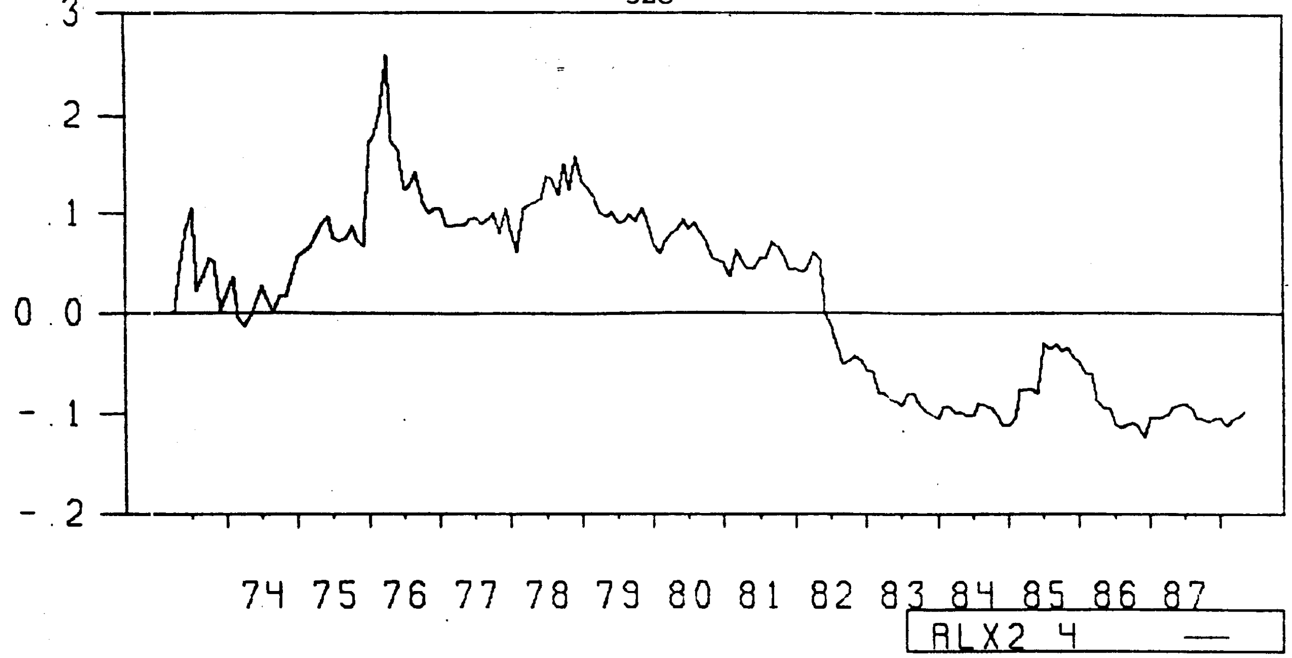
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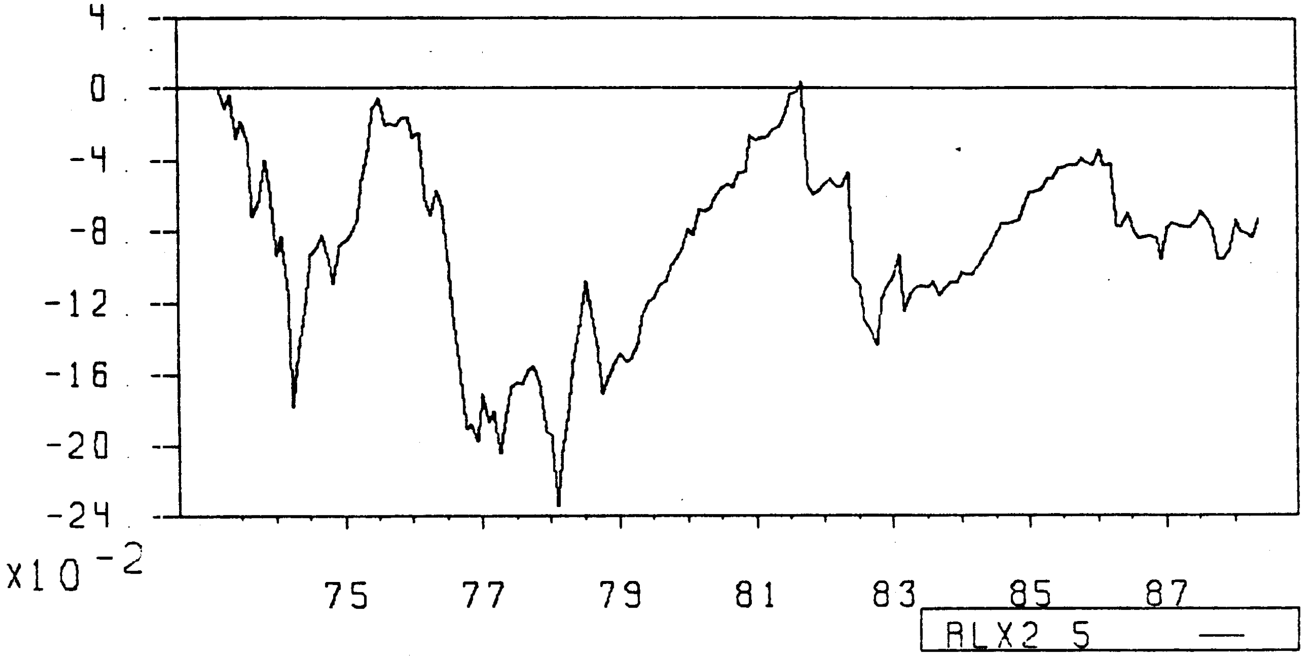
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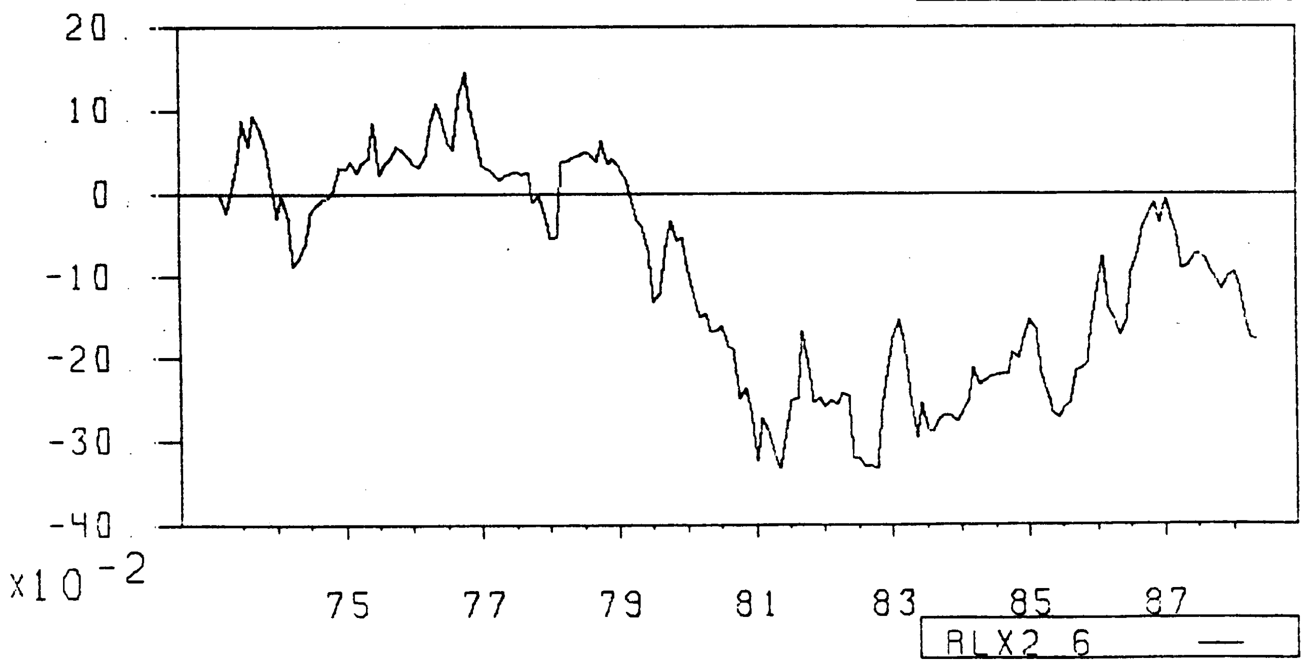
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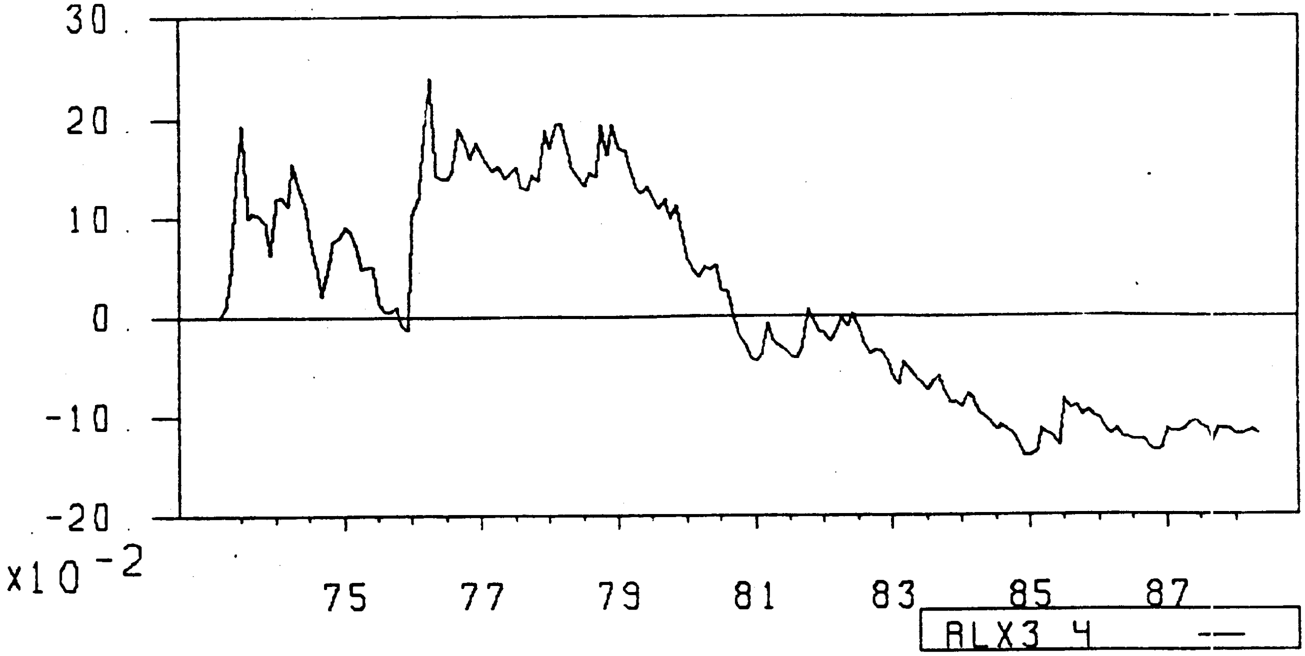
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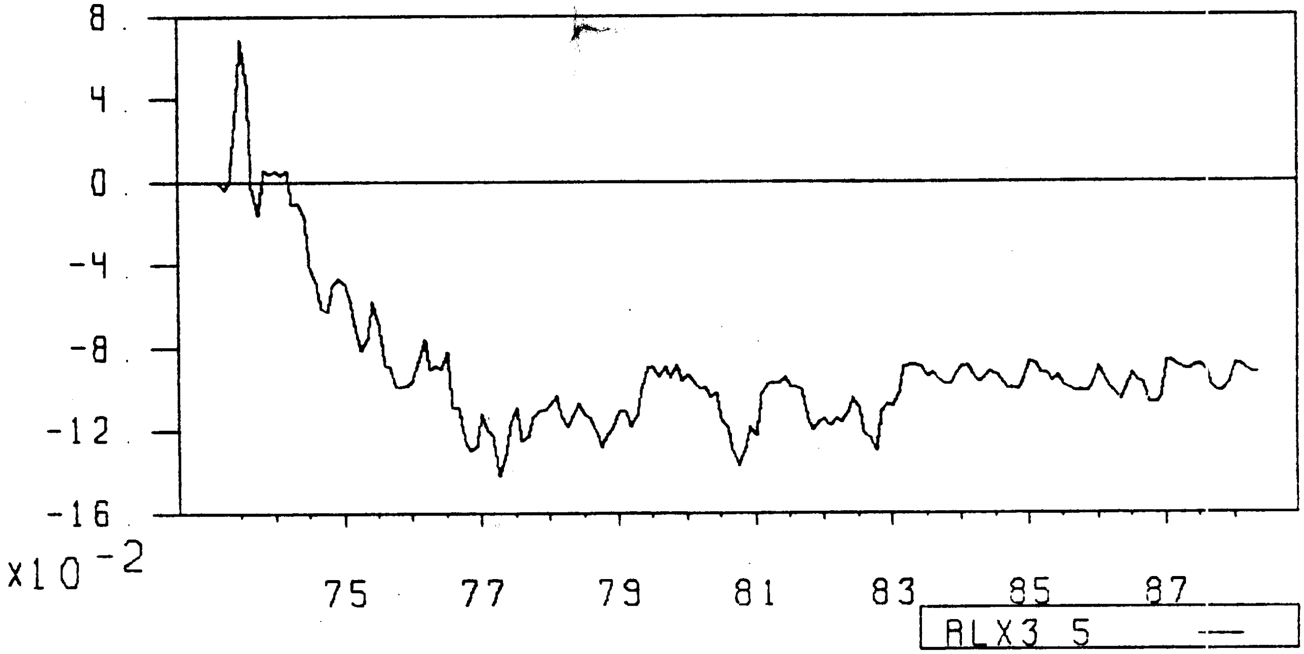
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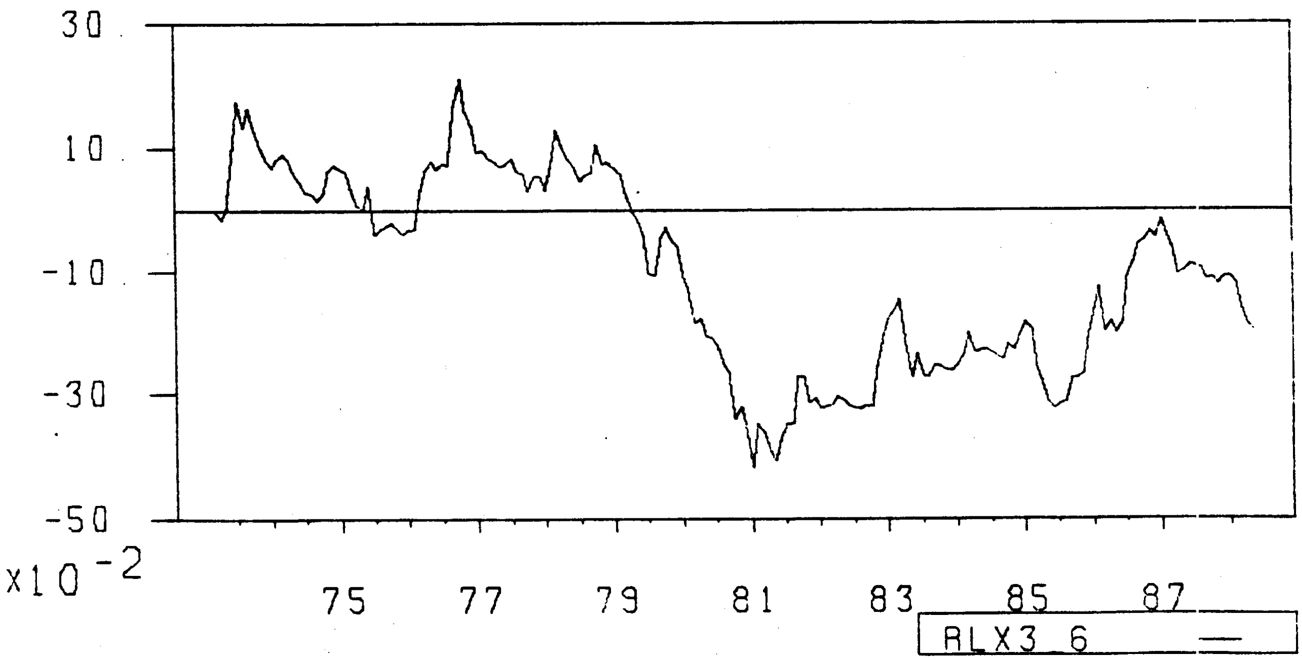
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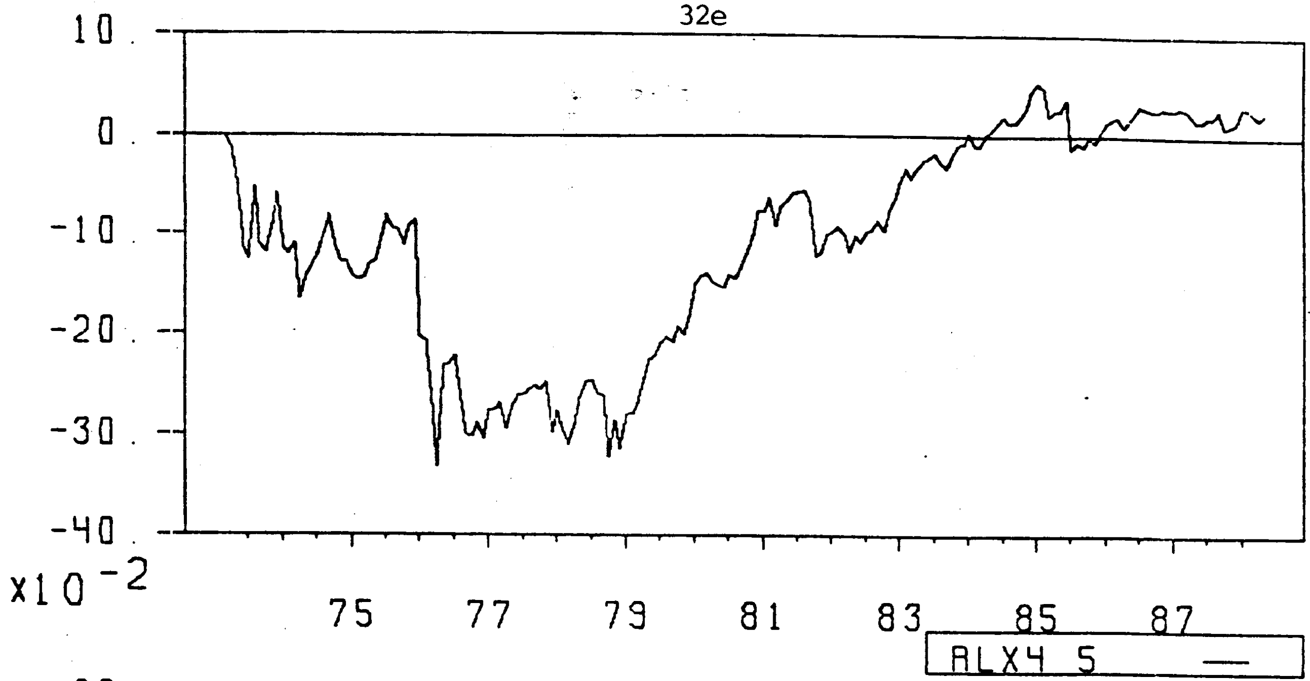
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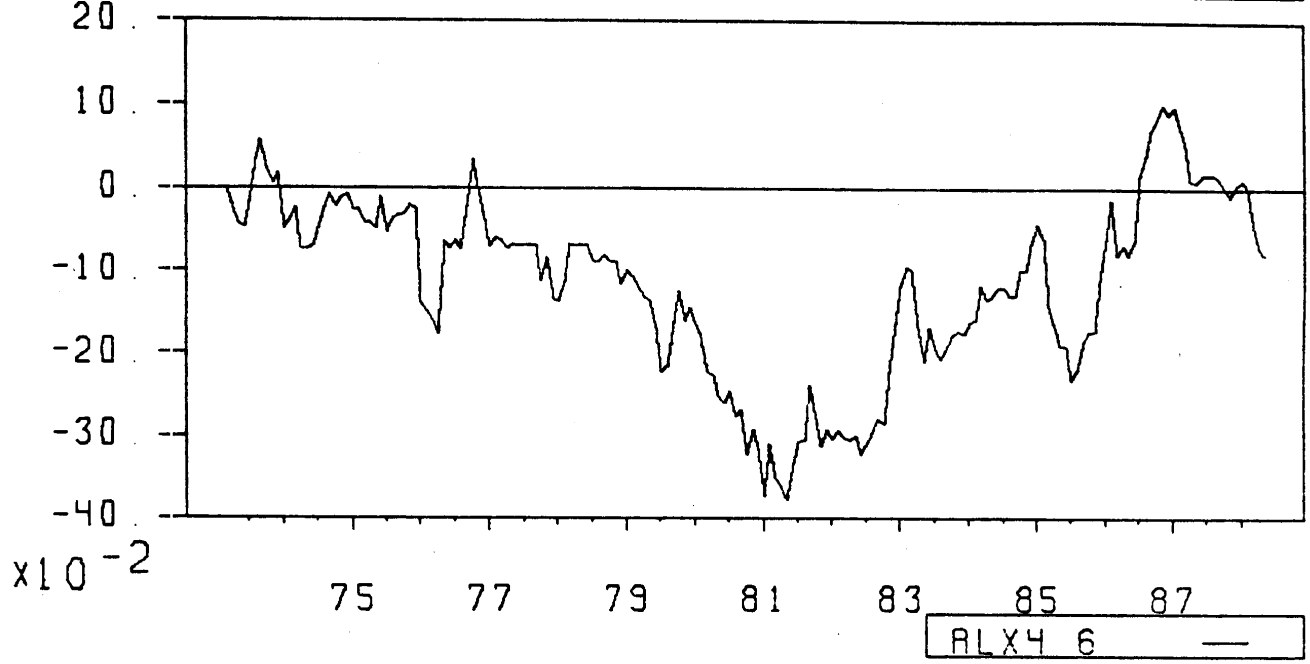
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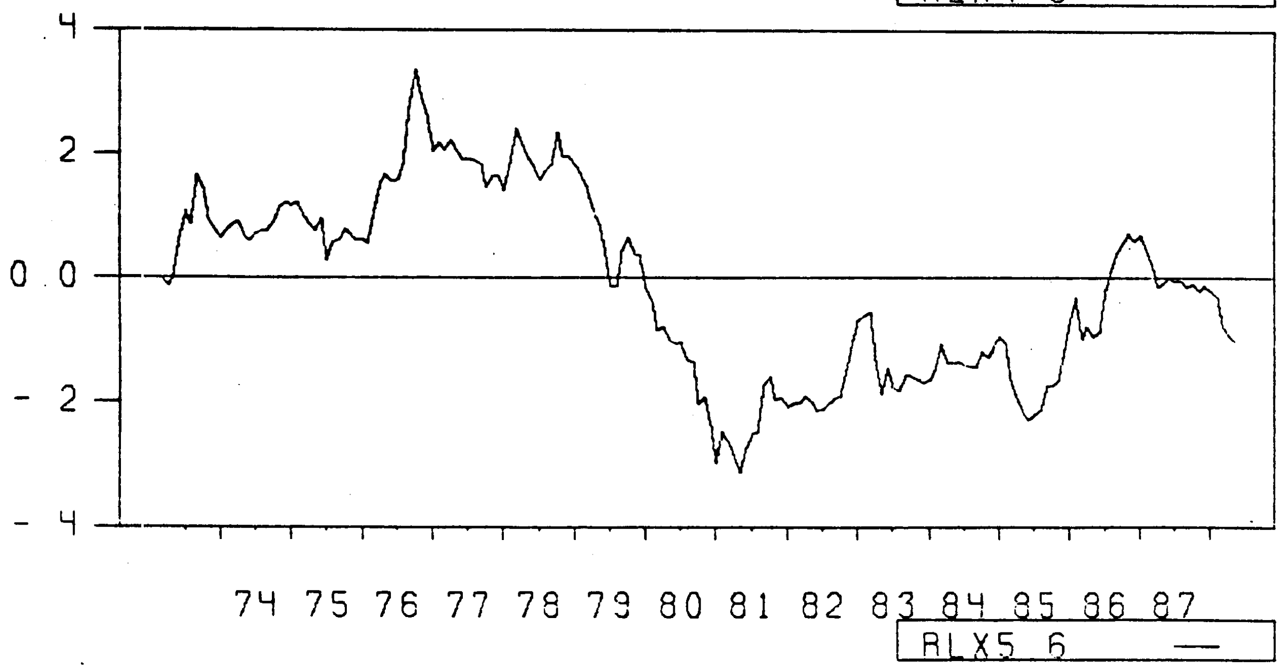
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RLX5 6

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