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*Jérôme Henry**
*Jens Weidmann***

**The French-German Interest Rate
Differential since German Unification:
The Impact of the 1992 and 1993 EMS Crises**

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* Banque de France, SEMEF

** University of Bonn

Institut für Internationale Wirtschaftspolitik

Lennéstr. 37

D-53113 Bonn

Phone: +49 228 73-9234, Fax: +49 228 73-7953

E-Mail: weidmann@unitas.or.uni-bonn.de

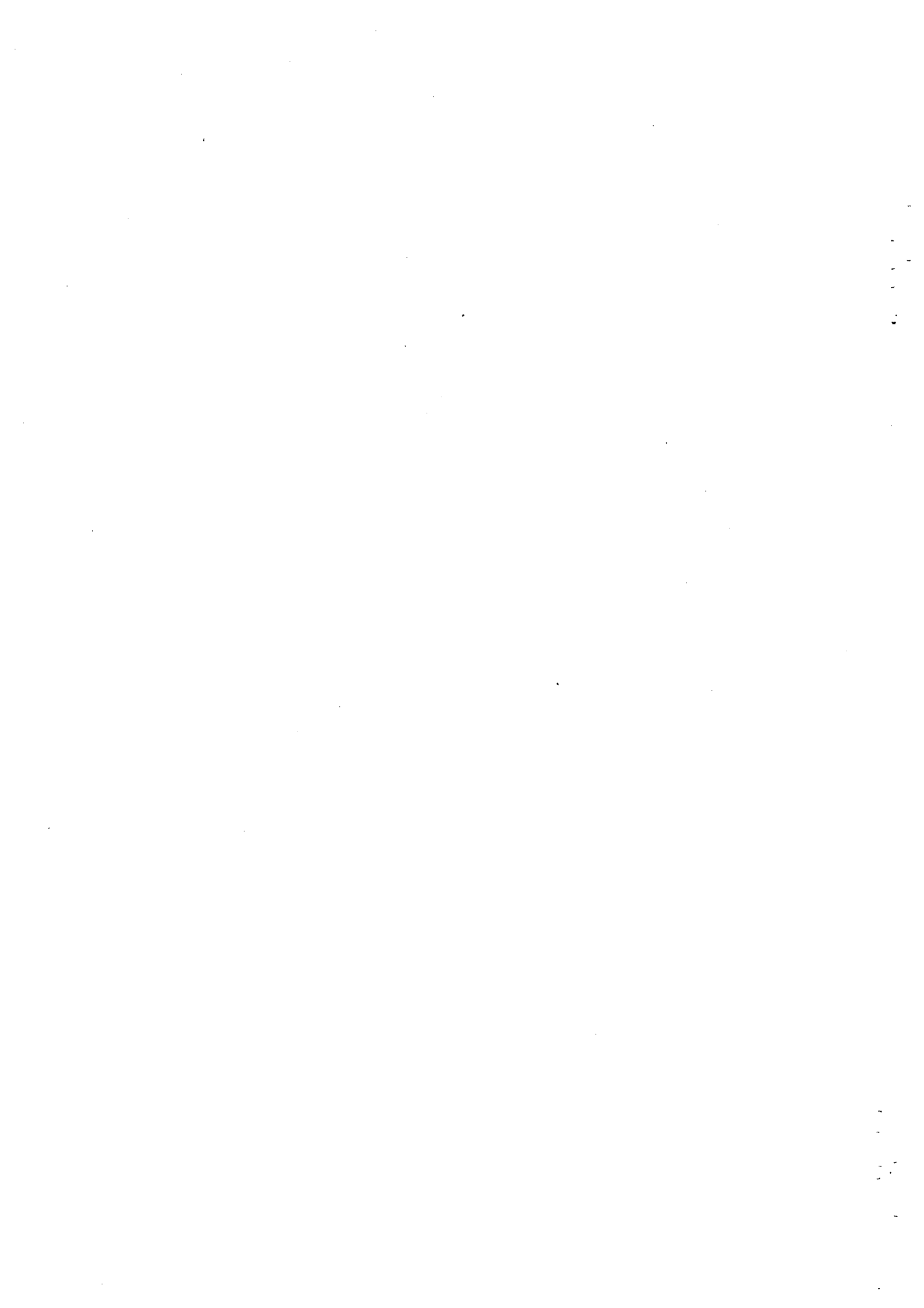
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Abstract: The EMS crises of 1992-1993, which resulted in the widening of the exchange rate bands, may have had some impact on the long-run structure of the system consisting of daily 1-month-Eurorates on German Mark, US-Dollar and French Franc. First, we find that both the US Eurorate and the German-French Eurorate differential are stationary over the December 1990 to December 1993 period, within a Gaussian VAR. Second, using GARCH-models to account for heteroskedasticity it is demonstrated that Gaussian models can induce a misleading interpretation of the linkages, namely about the effects of the American rate on the French one. In spite of the crises and the changes in the ERM, the volatility parameters for the German-French interest rate differential are quite stable over the sample. This can be related to the observation that the July 1993 crisis is not linked to a specifically high volatility.

Key words: Interest rates, Cointegration, Heteroskedasticity, GARCH, EMS, Asymmetry in the ERM

JEL classification: F33, E43, C32



Introduction

The severe pressures on the European Monetary System (EMS) which started in September 1992 finally led to the suspension of the narrow bands ($\pm 2.25\%$) within which the exchange rates of the member countries were allowed to fluctuate. By now there is a burgeoning literature on the reasons for the Exchange Rate Mechanism (ERM) crisis,¹ whereas there is virtually no study focusing on the question whether the widening of the ERM fluctuation band – often referred to as the *de facto* collapse of the EMS – actually induced a change in the system's structure. This expected structural change could have involved both the volatility and the linkages within the EMS. In terms of conditional volatility of the exchange rates, Ayuso *et al.* (1994) showed that only a few currencies have experienced a rise in volatility after the change in the ERM. We will ask whether the same kind of conclusion can be drawn from the analysis of the long-run linkages between the levels of interest rates.

More particularly, we will evaluate the extent to which the last ERM crises did affect the trivariate system that consists of daily observations of German, French and US one-month off-shore interest rates. As Henry and Weidmann (1994) – henceforth denoted HW – showed, applying non-stationary multivariate econometrics, there exists a structural break around German unification in the trivariate system: since the end of 1990 the long-term dynamics of the system is subject to more restrictions than before, with only one common stochastic trend left in the system.² As HW's sample period ends in December 1992, one could ask whether these results stand up when the period is extended to include the turbulences that hit the European Monetary System during 1993.

¹Examples are De Grauwe (1994) or Vlaar (1994). Most of these studies rely on the assumption that Germany independently determines the nominal interest rates in the EMS and that these interest rates were not 'optimal' for the non-German EMS countries on grounds of short-run considerations related to the recession.

²The authors apply the Johansen (1988, 1991) cointegration techniques. Both Katsimbris and Miller (1993) and Kirchgässner and Wolters (1993) have shown the superiority of an approach with more than two variables over a bivariate one, used e.g. by Karfakis and Moschos (1990). Yet, all these author's results are based on the 1979-1988 period.

Furthermore, we want to check whether the stronger evidence in favor of an asymmetric functioning of the EMS – prevailing from 1990 on, emphasized by both Gardner and Perraudin (1993) and HW – vanishes once observations posterior to July 1993 are included in the sample. In our context, asymmetry is defined as the presence of causality from the German rate to the French one, without there being any influence the other way around. There are two competing hypotheses about this issue. On the one hand, one could suppose that larger fluctuation bands partly resolve the ‘N-1 problem’³ leaving more room to the French central bank to pursue its own monetary policy and therefore reduce asymmetry. On the other hand, one could follow Herz (1993), who states that “the anchor role of the Deutschmark seems to be independent of the EMS exchange rate mechanism: it is the result of the autonomous decision by many European countries to pursue more restrictive monetary policies [...] via an exchange rate and not a monetary target.” Thus, the widening of the fluctuation bands should not have mattered in this respect and the asymmetry should have remained. We also want to check whether the hypothesis of German dominance still holds.⁴ German dominance is defined as the joint property consisting of the just mentioned asymmetry and the absence of influence of non-European rates, i.e. the American rate, on the French rate. This latter feature confers a so-to-say ‘monopolistic’ position to the German rate in the EMS.

The paper is organized as follows: In the first section, it will be checked whether the results of HW concerning the long-run linkages and the asymmetry in the

³ The ‘N-1 problem’ states that when N countries come together to fix their exchange rates, only one degree of freedom remains for the determination of monetary policy. This degree of freedom can be managed jointly by all participating countries (co-operative solution of the ‘N-1 problem’) or conferred to one country exercising a leadership role in the system. In the latter case an enlargement of the fluctuation bands gives some scope for the ‘dominated’ countries not to follow the monetary policy of the country holding the dominant position.

⁴ See Fratianni and von Hagen (1990) or HW for a more detailed exposition of this concept. HW also carry out a causality and long-run neutrality analysis according to which the German rate is the only weak exogenous variable in the system. Moreover the only non-neutral (see Stock and Watson 1989) stochastic trend in the system is that of the innovation in the German rate process. The German interest rate consequently appears to behave quite independently since 1990, and dominates the French Eurorate too.

trivariate system are modified by the inclusion of the observations for 1993, thus asking if one can already observe a significant break since the last ERM crisis.⁵ In the second section, an analysis of the time series behaviour of the French-German interest differential is presented (see Caramazza 1993 for a closer examination of the link between exchange rate expectations and interest-rate differentials at different maturities). If the uncovered interest parity (UIP) is valid, this differential is equal to the expected change of the exchange rate plus a risk premium. Special attention will be paid to the heteroskedastic nature of the data. Finally, some conclusions are given in section three.

1. Up-dating the long-run linkages in the Gaussian framework

The interest rates under review are one-month off-shore rates collected daily on the Euromarkets by the Banque de France at 11.30 a.m. Although intra-daily data were available, the daily frequency seemed to suffice as the rates hardly vary within the space of one day. The one month maturity was chosen – in common with other empirical studies on this topic, e.g. Gardner and Perraudin (1993) – on the basis of this market's good liquidity. The use of Eurorates lends the observations some homogeneity, which is on the contrary not ensured with the on-shore rates, due to a varying and non-zero subjective probability of re-adoption of capital movements controls. Indeed, the differential between off- and on-shore rates did increase during the recent EMS crises.

⁵The expression 'long-run linkages' is to be understood in a purely statistical sense, meaning the linear stationary combination of non-stationary variables, the so-called cointegration relations. Actually, in economic terms, the interpretation of these relations may be difficult, because of the short observation period we consider, no matter the large number of observations (see Hakkio and Rush 1991). Anyway the high frequency is absolutely necessary to causality analysis, otherwise bilateral causality is often found, e.g. Katsimbris and Miller (1993). Moreover, for statistical inference to be correct and dynamic specifications not to be misspecified, we do need to include those cointegration relations which are relevant to our system, contrary for instance to Gardner and Perraudin (1993) whose model is in first-differences of the interest rates. Such an analysis expressed in terms of 'long-run' dominance can also be found in Kirchgässner and Wolters (1993).

The observation period starts on December, 3rd, 1990. This date coincides with the most obvious positive shock to the German Eurorate, which occurs somewhat after German monetary unification. Recursive OLS and FIML analysis in HW indicate that a structural break did occur between 1989 and 1990, rather corresponding to a smooth transition process⁶ than to a one-time sudden break. Our sample begins when this transition is completed. The last observation included is December, 22nd, 1993, which means that we add a whole year to HW's sample.

A priori, the inclusion of the year 1993 in our sample should lead to an increased volatility, though this may depend on the model from which the conditional variance is derived (see Ayuso *et al.* 1994). This simply stems from the fact that this period contains important speculative crises and that about one third of the observations were made during the "new" ERM allowing departures from the central parities of $\pm 15\%$, thus institutionally allowing a greater volatility. Therefore, we have to re-estimate the trivariate VAR-model presented in HW to see if the number of long-run relations and the German dominance properties of the system were affected by this extension of the sample. The econometric model (R) we refer to is the following VAR($m+1$)model:

$$\Delta X = A(1)L^{m+1}X + \sum_{l=1}^m \Phi_l L^l \Delta X + M + E \quad (R)$$

where X is the vector of the three interest rates ($r_{FRF}, r_{DEM}, r_{USD}$), E is the error term and M a vector of intercepts. M is restricted so that no drift appears in the system, in which case M enters only the level component and ΔX has zero-mean. By definition, the number of 'long-run' relations, denoted r^* , is equal to the rank of $A(1)$. Once r^* is known with non-standard likelihood ratio (or LR) tests, some linear restrictions on $A(1)$ can be tested for with LR tests under standard χ^2 distributions, as presented in Johansen (1991).

⁶ See Luukkonen *et al.* (1988).

1.1 Two stationary components in the trivariate system of interest rates: The German-French differential and the dollar Eurorate

The number of cointegration relations does not vary with the lag length m in (R), which can be estimated to lie somewhere between 5 and 25, according to the limits found respectively following the Schwarz (SIC) and Akaike (AIC) rule. We find that r^* is equal to 2, for both Johansen tests (see Table 1).

Table 1: Johansen tests for cointegration in the gaussian VAR(10)

<i>Statistic and H_0</i>	$p = 0$	$p = 1$	$p = 2$
$\lambda_{\max} (r^*=p)$	29.0***	17.7**	4.4
Trace ($r^*\leq p$)	51.1***	22.1**	4.4

Note: The presented results are derived from the model and Data Generating Process with a restricted constant and no trend. ** stands for a null (H_0) rejected at a risk of 5% and *** for the like at 1%.

It is worth noting that in this case, i.e. two cointegration relations and three variables, it is always possible to rewrite the model so that each relation includes only two of the three variables, the two long-run relations being orthogonal by construction. On the contrary, the exclusion of one variable of *both* cointegrating vectors does imply an effective restriction, which is rejected for all three variables in our case (with standard χ^2 tests under the null of two stationary components).

Of course, one can still impose less demanding restrictions on the two orthogonal cointegration relations, say force the one including the French and the German rate to be equal to the interest rate differential and require moreover the exclusion of the German rate in the second one (say including Dollar and Mark rates). This is tantamount to saying that the US rate and the German-French Eurorate differential are stationary or, in other words, mean-reverting. This particular overidentifying restriction is not rejected by our data, but leads nevertheless to a considerable loss in the likelihood (LR test-statistic=4.5, χ^2 distributed, p-value=11%).

Table 2: Estimations results with an identified normalised long-run model

Variable	ECM Coefficients (%)		Cointegration vectors			
	λ_1	λ_2	for ($r_{FRF}, r_{DEM}, r_{USD}, constant$)			
r_{FRF}	5.1	0.39	$\mu_1 = (1$	-1	0	0.55)
r_{DEM}	0.17	0.15	$\mu_2 = (0$	0	1	-3.77)
r_{USD}	0.33	-1.2				

Note: μ is of dimension 2×4 , because, along with the three rates, one has to take into account the restricted constant in the long-run parameters. $\lambda\mu$ is the ECM component of the VAR dynamics, i.e. $A(1)$ in equation (R). Loadings, i.e. λ components, are typed in bold if greater than 1%.

If only the stationarity of the US rate is tested, we obtain a higher p-value (27%), which means that the hypothesis of a stationary French-German interest rate differential is the most fragile of the two assumptions. However, in the bivariate context considering only the French and the German rates, the stationarity of the interest rate differential (henceforth *ID*) is comfortably accepted at the 45% level.⁷ According to the point estimates found for the restricted system (see Table 2, vectors μ_1 and μ_2), the German-French interest rate differential is on average set at 0.55 percentage points, whereas the US rate reverts towards a mean of 3.77 points.

1.2 The impact of deviations from equilibrium: The weak exogeneity of the German Eurorate

Let us now turn to the identification of the error-correcting dynamics, given the accepted restrictions on the cointegrating vectors. The estimation of the relevant *ECM*

⁷ In the bivariate model the unrestricted long-run relation can be written as:

$$r_{FRF} = 0.6r_{DEM} + const. \Leftrightarrow ID = r_{FRF} - r_{DEM} = -0.4r_{DEM} + const.$$

which means that we face a risk premium on French assets which decreases with the level of the German rate (a similar result and unrestricted long-run coefficient is found over 1991-1992, see HW). Strictly speaking, we cannot tell for sure that any of our series is definitely $I(0)$ or $I(1)$, due to observational equivalence considerations (see Campbell and Perron 1991). All we can say is that *during a particular interval in time* the series behaves like $I(1)$ series and, in this case, the most appropriate econometric technique is the non-stationary battery, allowing for cointegration (see Campbell and Perron 1991 again). This is especially the case for interest rates, where we feel quite uneasy about accepting the non-stationarity on theoretical grounds: the less numerous the stochastic trends, the more convincing the econometric model.

coefficients (i.e. the loadings of both long-run vectors) leads to the results in Table 2, in which λ and μ are the loadings and the cointegration relations respectively. Apparently the European rates are disconnected from the US Euro rate as only the French-German interest rate differential, corresponding to μ_1 , influences the French rate, whereas the German rate does not seem to be influenced by any of the two long-run relations. The US rate, on the contrary, depends only on its past, as far as levels are concerned. It is worth recalling however, that these findings do not exclude transitory effects from the American variable to the European rates (coming from the terms in first difference in the VAR). Having accepted the above overidentifying restrictions on the long-run components, we can therefore test for and confirm the weak-exogeneity property of the German rate already established in HW (p-value: 64% in the trivariate system and 58% without the Dollar rate). This property also holds with unrestricted cointegration relations.

Besides, in terms of common trends, the neutrality test in the trivariate ECM-VAR is reduced to a simpler test, that can be implemented in the bivariate system (exactly like in the Stock and Watson (1989) analysis of money-output causality) involving only the French and the German rates, as the US rate is $I(0)$ and consequently does not contain any stochastic trend at all.

In sum, the neutrality tests with the sample including 1993 are similar to the ones excluding this year. The German variable is not affected by the trend specific to the French rate (Wald-test, p-value: 86% for the German rate). Once the system is orthogonalized, so that the instant correlation between both European rates has been removed from the French innovation, the MA representation of the ECM-VAR yields a zero long-run response of the German rate to an impulse from the French rate. On the contrary, the French rate is indeed influenced by the orthogonal component of the German innovation (p-value=0.9%).

Therefore, the re-estimation of the cointegrated VAR of HW over the extended sample confirms the stability of the number of long-term relations in spite of the turmoil affecting the Euro franc in January and July 1993. In addition, the identification of these two cointegration relations indicates two stationary variables: the French-German interest differential and the American rate. Note that it is

impossible to confirm the stationarity of either of these two variables by analyzing only the data for 1990 and 1991. This means that econometrically, we are now able to accept those particular identifying restrictions which look intuitively appealing from an economic point of view, contrary to the relations proposed in HW where the sample stopped in December 1992. In spite of the high-frequency and short time period covered, the cointegration analysis does bear economic sense, which is all the better.

Moreover, the magnitude of the error-correction terms suggests that there exists a 'partition cointegration' (see Konishi *et al.* 1993 for a definition and application of this concept) in the system, according to which each of the two cointegrating vectors would solely enter the dynamics of its own long-run components. In other words, the US rate appears weakly exogenous with respect to the German-French interest rate differential, whereas both European rates would be weakly exogenous with respect to the Eurodollar rate. This corresponds to a kind of a long-run disconnection between the European and the American markets for short-term assets.

1.3 A closer look at the French-German Eurorate differential

Independently of the causality linkages among these three interest rates one could question the stability of the French-German interest rate differential, which appears stationary when the period 1991-1993 is considered, but non-stationary over the period 1991-1992. The first explanation for this somewhat puzzling result one could think of is that the observation period from 1991-1992 seems to be too short to fill the long-term relation with economic sense (see Hakkio and Rush 1991). The small time period covered by the sample, in spite of the large number of observations used, appears to be a serious problem for economic interpretation. The power problem, mentioned by Pearse and Snell (1995), is not very important here, since we have almost rejected all the non-stationarity in our data. Moreover, the error-correction coefficients are low so that only a small portion of deviations from equilibrium is corrected each day, which makes the adjustment towards equilibrium quite slow. For instance, the crises at the end of 1992 and at the summer of 1993 are so important that

it is only during the last four months of 1993 that a return to the assumed 'equilibrium value', which clearly shows up since 1991, could be observed.⁸

A second possible explanation might also be related to the daily observation frequency. This frequency appeared to us as a necessary condition to examine the causal linkages at work (see Gardner and Perraudin 1993 for a similar concern). Unfortunately, high frequency time series depart quite often – at least in finance – from the usual hypotheses made in both, stationary and non-stationary econometrics. The daily Eurorates we analyze are certainly no exception to that rule. During the speculative crises occurring in our sample period the volatility of the time series considered, especially that of the French Eurorate, increases considerably, whereas shortly after the end of the crisis the volatility dampens again to return to its pre-crisis values. Apparently, this is at odds with the representation of the time series as an integrated process (as for instance in the previous section or in Katsimbris and Miller 1993), which has either a trending or an infinite variance. Nevertheless, it is generally suggested, see for example Diebold *et al.* (1993), that in spite of its rough description of reality the use of cointegration techniques is useful to capture the long-run dynamics of the system under study.

Indeed, the study of the residuals from the restricted VAR estimated in HW clearly reveals that the variances of these residuals vary over time (see Henry 1993). The most appropriate model to cope with this kind of problem seems to be the Generalized Auto Regressive Conditional Heteroskedastic GARCH(1,1) model presented in Bollerslev 1986 (see Bollerslev *et al.* 1992 or West *et al.* 1992 for surveys of the models to be used when dealing with interest rates).

⁸As can be seen, since 1983 the interest rate differential converges slowly to zero until 1991, when a side movement initiates. The mean-reversion property (according to the usual ADF tests or Johansen LR tests) is nevertheless crucially linked to the stages through which the series returns to its 'normal' mean after a crisis. Excluding the speculative periods does lead to even more (observed) non-stationarity, as found by Kunst (1992) who showed that the mean-reversion mechanism works only when the shocks are large enough. See Figure 1 for a plot of the French-German interest rate differential. The ADF analysis does not give clear-cut results as to the I(1)-ness of our data, contrary to the findings of Kirchgässner and Wolters (1993).

The conditional variance, σ^2 , that is the residual variance of a given model (X) linking series x and y , is defined by:

$$\begin{cases} \sigma_t^2 = \alpha\sigma_{t-1}^2 + \beta\varepsilon_{t-1}^2 + \gamma + \phi_t & \text{(I)} \\ y_t = Ax_t + \varepsilon_t & \text{(X)} \end{cases}$$

where ϕ_t represents a white noise error term.

Thus, the conditional variance depends on its past value and the last squared shock. There are obviously some similarities with the ARMA(p,q) class of models. In the case of a zero β , one gets the so-called ARCH(1) model, where the conditional variance is simply autoregressive and follows an AR(1) process. Moreover, like in the ARMA-case, the time series properties of the variance are closely linked to those of coefficients α and β : both α and β have to be non-negative and their sum must be less than one. If either of the coefficients is negative one could obtain negative values of σ_t^2 which is impossible for a variance. If the sum of α and β is equal or greater than unity the corresponding time series of ε_t would not be covariance stationary which means that its variance σ_t^2 does not converge and could even show an explosive behaviour. Our residuals from the VAR are actually integrated in variance (the sum of α and β approaches unity), denoted IGARCH(1,1). An alternative representation of the crises could be the inclusion of discrete jumps (e.g. Poisson processes), Vlaar and Palm (1993), for example, modelled exchange rates in this way.

2. Modelling the interest rate differential's mean and volatility

In spite of the heteroskedasticity which affects our data, the long-run linkages among the three variables are admittedly quite robust to the erroneous specification of the variance process underlying the FIML estimation of the VAR. In fact, Gonzalo (1994) showed that, when the process exhibits ARCH-effects, the Johansen (1988, 1991) procedure is the most robust method to detect long-run interactions. Even the results of the Dickey-Fuller test seem not to be seriously altered by special cases of heteroskedasticity, i.e. GARCH (1,1) effects, as suggested by Kim and Schmidt (1993) or Hecq and Urbain (1993). Similarly, the Wald tests (asymptotically equivalent to likelihood ratio tests) for causality are not affected, though the lag length up to which causality occurs would appear longer than it actually is, according to

Hecq (1993). In our case, the MA-component is unfortunately much larger than the AR-component, which suggests the robustness of standard Gaussian tests is more affected than it would be with a highly autoregressive conditional variance.⁹ As a consequence, we have to investigate further the time series properties of the French-German interest rate differential, since this variable appeared stationary for standard (Johansen's LR and DF) tests during the period of interest (1991-1993), but without explicitly considering the heteroskedasticity in the data. This will be done with various sets of conditioning variables, i.e. various specifications for (X), ranging from a constant up to a full ECM. Moreover, the latter case will allow for a test of the asymmetry and German dominance hypothesis. We will also establish a systematic comparison between the estimations over two samples, 1991-1992 and 1991-1993, to track the break in the underlying parameters, if any.

2.1 Autoregressive and moving average components of the conditional variance

A natural starting point to account for heteroskedasticity is the estimation of an ARCH(1)-model, i.e. a GARCH with β set to zero in (I). We also need a model to derive the conditional variance we are going to model. Given that the focus in financial time series analysis is rather on the volatility and covariances (like in asset price models), the models for the mean are often simplistic. Accordingly we will begin with such a coarse behaviour for the mean of the French-German interest rate differential, particularly with the following system:

$$\begin{cases} ID_t = \delta_0 + \varepsilon_t & \text{(Xm)} \\ \sigma_t^2 = \alpha\sigma_{t-1}^2 + \gamma + \phi_t & \text{(II)} \end{cases}$$

The estimation indicates strong ARCH-effects which are – without doubt – strengthened by the absence of any autoregressive dynamics in the conditional model for ID , which is simply demeaned through (Xm). Obviously, the characteristics of the residual series may depend upon the specification of the conditioning model (X).¹⁰

⁹For instance, the coefficient α in the case of the Eurofranc is too important (> 0.25).

¹⁰The model is estimated by maximum likelihood (ML). More specifically, the GARCH process might be estimated by three different procedures:

The AR-coefficient in the conditional variance equation is 1.003 (for the 1991-1993 period), thus being almost equal to unity. In addition the constant is different from zero consequently leading to a permanently growing variance, as it appears on Figure 2, where SE denotes the estimated standard error, i.e. σ_t .

The same type of rather brute conditional model for *ID*, this time with a GARCH(1,1) disturbance, i.e. using (I) instead of (II) along with (Xm), reproduces the results of the ARCH(1) model about the integrated variance but differs by the fact that the integration is due to an MA-coefficient β around unity, whereas the AR-coefficient α comes close to zero. This result stands up whatever period is used, hence is robust to the inclusion of the 1993 data.

It is nevertheless obvious that the model generating the residuals, (X) has to be a bit more intricate than the above-mentioned one. For instance King *et al.* (1994) used a multivariate factor model. In our case, a more 'realistic' model for *ID* can be derived from the analyses usually performed with less heteroskedastic data, e.g. Eurorates at lower frequencies. First we will consider autoregressive components and

(i). Simultaneous one-step joint estimation of α , β and the parameters of the model for *ID*. The log-likelihood function depends on α , β and the parameters of the model for *ID* and can be written as: $-1/2 \log|\sigma_t| - 1/2(\varepsilon_t \sigma_t^{-1} \varepsilon_t)$. This is the 'ideal solution' according to King *et al.* (1994).

(ii). Two-step estimation by separating the long run of the model from the short run. First, the model for *ID* is estimated by assuming normal error terms. Then α , β and the long-run parameters of the model for *ID* are computed by the simultaneous ML method described under point (i). Restricting the long-run parameters to be equal to their first step estimates means assuming that the cointegration parameters are estimated without bias when the GARCH effects are neglected (which can be expected to be true, cf. Gonzalo 1994).

(iii). Estimation by a two step procedure where the parameters of the model for *ID* and α , β are estimated separately. Initially, the *ID* model is estimated by Ordinary Least Squares (OLS); the residuals of this estimation are then used to estimate α , β by maximum likelihood. Then the model for *ID* is re-estimated by Generalized Least Squares (GLS) taking into account the residual variances estimated during the last step. The residuals of this regression are then used to re-estimate α , β by ML and so forth upon achievement of convergence.

Here, (ii) is used, the results from Henry (1993) derive from (i). Both methods are superior to (iii) in order to reach an asymptotically consistent covariance matrix, similar to the one used in McCurdy and Morgan (1991). All calculations have been performed with RATS under the BHHH algorithm.

second we will include error-correction dynamics, which is here tantamount to including some effects of the level of the interest rate differential and of the lagged differenced values of the French and German interest rates (allowing for the absence of any common factor between the latter two). In addition, a third model could include some influence of the American Eurorate on the French one in a context where the EMS functions asymmetrically, the German rate causing the French one, without there being German dominance. This leaves us with three types of (X) models and the same number of conditional variances to be estimated.

2.2. Variance of the interest rate differential adjusted for autoregressive and ECM parts

The first dynamic model we estimate explains the interest rate differential by its own lagged values, this is the so-called AR(10)-GARCH(1,1) formulation:¹¹

$$\begin{cases} ID_t = d + \sum_{i=1}^{10} a_i ID_{t-i} + \varepsilon_t & \text{(XAR)} \\ \sigma_t^2 = \gamma + \alpha \sigma_{t-1}^2 + \beta \varepsilon_{t-1}^2 + \phi_t & \text{(I)} \end{cases}$$

It is obvious from the estimates given in Table 3 that the inclusion of autoregressive term in the model for ID does not change the non-stationarity property of the model's variance. This result holds true even for the error correction model with the interest rate differential as long-term attractor which has the following representation:

$$\begin{cases} \Delta ID_t = d + \sum_{i=1}^{10} a_i \Delta r_{FRF_{t-i}} + \sum_{i=1}^{10} b_i \Delta r_{DEM_{t-i}} - c ID_{t-1} + \varepsilon_t & \text{(XEEM)} \\ \sigma_t^2 = \gamma + \alpha \sigma_{t-1}^2 + \beta \varepsilon_{t-1}^2 + \phi_t & \text{(I)} \end{cases}$$

Table 3: AR-GARCH process for the French-German Eurorate differential

	γ (%)	α	β	Likelihood
1990-1992	0.04 (2.8)	0.64 (16.5)	0.76 (8.2)	750
1990-1993	0.07 (3.4)	0.66 (24)	0.64 (10.2)	1005

Note: *t*-statistics in brackets

¹¹For the LM or F tests only the first ten lags are significant.

This model postulates that the French rate is tied to the German rate in the long run, but in the short run the two rates are allowed to differ. To sum up, the estimations shown in table 4 and the previous ones, with (XAR), lead to very similar parameters and conditional variances (see Graphs 3 and 3a).

Table 4: ECM-GARCH process for the French-German Eurorate differential

<i>Period</i>	γ (%)	α	β	<i>Likelihood</i>	c (%)
1990-1992	0.05 (2.5)	0.55 (11.4)	0.82 (6.2)	793	1,1 (0,6)
1990-1993	0.06 (3.3)	0.65 (22.5)	0.59 (8.5)	1033	1,9 (1,6)

Note: t -statistics in brackets.

Besides, there is some difference between the two periods examined, which lies in the significance of the error-correction coefficient. The coefficient c becomes significant – even though at a low level – only when the year 1993 is included in the sample period. Indeed, this finding is consistent with the multivariate FIML analysis where we could accept the joint stationarity of both ID and the Eurodollar rate only when the 1993 observations were included in the sample. It is worth noting that the OLS-estimates do not corroborate these results. OLS-estimates of (XECM) yield $c = 5.5\%$ (t -stat = 4.5) over 1991-1992 and $c = 8\%$ (t -stat = 4.3) over 1991-1993. This contradiction illustrates the relevance of FIML (as compared to OLS) for the analysis of long-run linkages in the presence of heteroskedasticity (*cf.* Gonzalo 1994).

Moreover, as can be seen in Figures 4a to 4c the distribution of the standardized errors, i.e. the ratio ε_t/σ_t , approaches a homoskedastic distribution. This observation can be confirmed by a test for ARCH, which is to regress the squared residuals series on its own lags and then test for the joint significance of the lagged terms. As we expected, considering the first ten lagged values of ε_t^2 , the LM-test and F-test both lead us to accept the hypotheses that the corresponding coefficients are all zero at the 20% significance level. Nevertheless, the estimated variance σ_t^2 is still characterized by a divergent behaviour.

2.3. Making allowance for the Dollar's influence

Since we found a second long-term relation in the Gaussian Johansen's VAR, namely the one corresponding to a stationary Eurodollar rate, we can also estimate a model taking care of the effect of the Dollar rate:

$$\begin{cases} \Delta ID_t = d + \sum_{i=1}^{10} a_i \Delta r_{FRE}_{t-i} + \sum_{i=1}^{10} b_i \Delta r_{DEM}_{t-i} + \sum_{i=1}^{10} b_{1i} \Delta r_{USD}_{t-i} - c_1 ID_{t-1} - c_2 r_{USD}_{t-1} + \varepsilon_t & \text{(XECM\$)} \\ \sigma_t^2 = \gamma + \alpha \sigma_{t-1}^2 + \beta \varepsilon_{t-1}^2 + \phi_t \end{cases} \quad (I)$$

When (XECM\$) is OLS-estimated, the two error-correction coefficients for the interest rate differential and the American rate are equal to 5.8% (t -stat=4.5) and 0.6% (t -stat=0.5) respectively, thus indicating at first glance the absence of any long-run influence of the US rate on the French rate. However, the estimation allowing for heteroskedasticity, (XECM\$) along with (I), leads to somewhat different results and consequently to a different interpretation of the interactions in the trivariate system (see Table 5).

Table 5: ECM-GARCH process for ID (Eurodollar rate included)

<i>Period</i>	γ (%)	α	β	<i>Likelihood</i>	c_1 (%)	c_2 (%)
1990-1992	0.0 (0.3)	0.56 (11.4)	0.92 (6.2)	837	3.6 (2.5)	-1.3 (4.3)
1990-1993	0.02 (1.9)	0.61 (27.9)	0.71 (8.5)	1082	2.6 (2.8)	-1.4 (4.8)

Note: t -statistics in brackets

By accounting for time-varying variance the coefficient of the Dollar rate becomes significantly positive. Hence, the interest rate on the Eurofranc is not only influenced by the evolution of the German interest rate, but depends also on the American one, though with a much lower coefficient. Roughly speaking, the effect on the French rate of a given shock on the Deutschmark rate has an effect twice as large as a shock of the same size on the American rate.

It is also a bit risky to put too much weight on the result as we are still in the presence of an explosive conditional variance (see Figure 5 for the sample including 1993). Even in the present specification the sum of α and β is largely greater than

unity. This leads us finally to the estimation of an integrated GARCH (IGARCH), where the sum of the aforementioned coefficients is forced to equal unity.

The estimated specification resembles closely to the previous system, but only one parameter has to be estimated in the GARCH dynamic (III):

$$\left\{ \begin{array}{l} \text{(XECM\$)} \\ \sigma_t^2 = \gamma + (1 - \beta)\sigma_{t-1}^2 + \beta\varepsilon_{t-1}^2 + \phi_t \end{array} \right. \quad \text{(III)}$$

As before, both of the error-correction coefficients (see Table 6) are of the same size and the estimated variance corresponds closely (see Figure 6), yet two facts seem disturbing. First, the likelihood declines considerably. Second, the squared standardized residuals plotted in Figure 7 show an autocorrelation significant at the 3% level (for the F- and LM-tests).

Table 6: ECM-IGARCH process for ID (Eurodollar rate included)

<i>Period</i>	γ (%)	β	<i>Likelihood</i>	c_1 (%)	c_2 (%)
1990-1992	0.01 (4.4)	0.46 (11.4)	813	4.5 (2.5)	-1.9 (5.7)
1990-1993	0.01 (5.4)	0.36 (13.8)	1052	3.8 (4.6)	-1.6 (5.1)

Note: *t*-statistics in brackets

The residual non-normality may also be connected to some asymmetry in the distribution, with large positive residuals occurring more frequently than negative ones, due to the speculative waves against the French franc. In this case the GARCH models presented above would not be sufficient and had to be replaced by asymmetric GARCH models, i.e. the following ECM-E-GARCH model:

$$\left\{ \begin{array}{l} \text{(XECM\$)} \\ \sigma_t = \alpha\sigma_{t-1} + \beta_1\varepsilon_{t-1}^2 + \beta_2\varepsilon_{t-1}^2\xi_{t-1} + \gamma + \phi_t, \quad \text{(IV)} \\ \text{where } \xi_t = \begin{cases} 0 & \text{for } \varepsilon_t \leq 0 \\ 1 & \text{for } \varepsilon_t > 0 \end{cases} \end{array} \right.$$

Moreover, we can see that the scaled residuals of July 1993 are not larger than, for instance, the values of September 1992. This could indicate that the crisis that led to the collapse of the ERM was not so crucial in terms of the shock's relative magnitude. The same kind of conclusion is derived from Figure 6b, where can be seen that the

time-varying standard-error SE (or σ_t) has also been much higher in September 1992 than in July 1993. The former ERM has, apparently, been much more damaged by the occurrence of many crises in a row than by the mere size of the July 1993 shock.

Another possible solution to the unsatisfactory fit of the IGARCH model could be to link the level of the interest differential to its volatility. The resulting model is the so-called GARCH in mean or shortly GARCH-M model. The underlying economic rationale is that the risk premium is positively related to the uncertainty in the market, which can be approximated by its volatility. Henry (1993) found some evidence that there is a correlation between the estimated volatilities of the VAR-GARCH and the level of the interest rates, thus indicating a GARCH-M model. However, the correlations found are quite unstable shading some doubts on the reliability of the results, even only over a few years.

The conditioning model might also have been unstable because of variations in the expected exchange rate changes (see Ayuso *et al.* 1994). Therefore, we have re-estimated the system (XECM\$) + (II) with the inclusion of a break in mean in April 1993. We had the intuition that the integrated variance, as well as the integrated levels, might as well come from a misspecification of the deterministic components of *ID*. This date corresponds to end of a period of continuous speculative attacks against the French franc, initiated in September 1992. The speculative period ended temporarily after the general elections of March 1993, when devaluation expectations proved to have been wrong (see Figure 1), as the new government did not allow a devaluation of the Franc, usual device in the French political tool-kit prior to 1988. Though this break is indeed significant, the conditional variance remains integrated and, moreover, the ECM term properties are not changed by the inclusion of the dummy variable.

Regardless of the integrated variance and of the asymmetric distribution of our standardized residuals, we can ascertain a remarkable temporal stability of the parameters in the dynamic of the variance and of the error-correction coefficients. This leads us to conclude that, if any, the structural break in the system must have modified principally the short-run dynamics, i.e. by altering the influence of the current and lagged values of the differenced Eurorates on the French rate, rather than

the conditional variance process and the long-run linkages. Moreover, the effects of German interest rate shocks differ from those of the American rate, since the latter, contrary to the former, has mean-reversion properties on the sample and thus possesses no stochastic trend.

Up to now, our analysis is still confined to reduced form models, though the VAR innovations of the French and German Eurorates on the one hand, and of the German and the American rates on the other hand, appear to be correlated. In a structural form, one would add a term $\Delta r_{DEM,t}$ to (XECM\$) and re-test for the significance of $r_{USD,t-1}$. By doing so, the t -stat for the latter variable may fall down if the current changes in the German rate were correlated with the past values of the American rate. Actually this is the case according to the OLS estimates in a Gaussian framework or even with the estimates of a GARCH model.

On the contrary the ML estimate of E-GARCH disturbances around a structural model shows again evidence of a significant link between past levels of the American Eurorate and changes in the French rate and, moreover, greatly improves the likelihood. For all reduced form specifications, the instant correlation between the French and German rate as well as the ECM term c_1 are always significant, contrary to the April 1993 dummy that becomes insignificant (see Table 8 for the structural model with EGARCH).

Table 8: Structural ECM-EGARCH process for ID (Eurodollar rate included)

<i>Period</i>	γ (%)	α	β_1	β_2	<i>Likelihood</i>	c_1 (%)	c_2 (%)
1990-1993	0.03 (3.7)	0.61 (18.8)	0.25 (4.9)	0.77 (7.0)	1264	2.1 (2.3)	-0.9 (4.1)

Note: t -statistics in brackets. The likelihood for the reduced form ECM-EGARCH would be 1116.

The asymmetry in the response to shocks is quite obvious, the positive shocks on the French rate being three times more persistent on the variance than negative ones. The integrated variance does arise from this specific response to innovations corresponding to increases in the French rate. It is also interesting to find that the AR component of the conditional variance is not affected at all by the asymmetry in the MA component.

3. Conclusions

Besides some modifications in the transitory dynamics of our trivariate system, the year 1993 and the so-called collapse of the EMS do not seem to have affected the long-run linkages of the system (i.e. the linkages between the levels of interest rates) when analysed in a Gaussian cointegrated VAR framework.

The conclusions concerning the asymmetry of the relation between the French and German interest rates are not jeopardized: the French rate adjusts to departures from equilibrium of the interest rate differential between the two ERM member currencies whereas the German rate is not influenced by this error-correction term. Moreover, both the weak exogeneity and long-run neutrality tests (presuming Gaussian errors) accept the neutrality of the interest differential with respect to the German rate. The German interest rate also appears quite independent from any other interest rate, including the American one, since the end of 1990.

The German dominance, which in our context amounts to adding the absence of American influence on the French rate to asymmetry, is also verified when only the stochastic trends (i.e. the long-run linkages) are considered. This finding simply stems from the different time series properties of the series: it is clear that the non-stationary component of the French Franc cannot be due to the American Eurorate as the latter is stationary.

Regarding the volatility of the French-German Eurorate differential, it is worth noting that some of the specifications finally retained (GARCH errors with a reduced form or E-GARCH errors with a structural form) lead to question the German dominance with respect to the stationary variables as, contrary to the homoskedastic model, the American rate exerts a significant influence on the French rate in these specifications. Anyway, the amount of information contained in the German rate is much larger than that in the American rate since the latter is stationary whereas the former appears non-stationary over the period under review. This means that the prevailing influence remains, by far, that of the German rate, meaning that German dominance is not a very unrealistic assumption.

Moreover, the parameters defining the underlying process for the time-varying variance seem quite stable, when the periods 1991-1992 and 1991-1993 are compared.

In this framework, the crisis of July 1993 is not associated with a significantly higher volatility, insofar as the whole period between September 1992 and August 1993 has been rather volatile.

Though we have shown that to derive the conditional variance the use of an ECM specification for the interest rate differential was superior to simpler dynamic conditioning models, we are nevertheless left with either some explosive variance or an asymmetric distribution of the residuals. These findings suggest that, similar to exchange rates, we might get a better fit by including discrete jumps or using asymmetric GARCH models when modelling interest rates.

Consequently, we face a dilemma with respect to the analysis of causal linkages. If we use monthly or quarterly instead of daily data, which are obviously much less flawed with heteroskedasticity, we run the risk of getting massive bilateral causality (i.e. we reject asymmetry too often). Conversely, if we stick to daily data we have to manage the increased complexity of the conditional variance prior to completing the analysis of causality at shorter time horizons, because the results are sensible to the conditional variance's model as was shown above.

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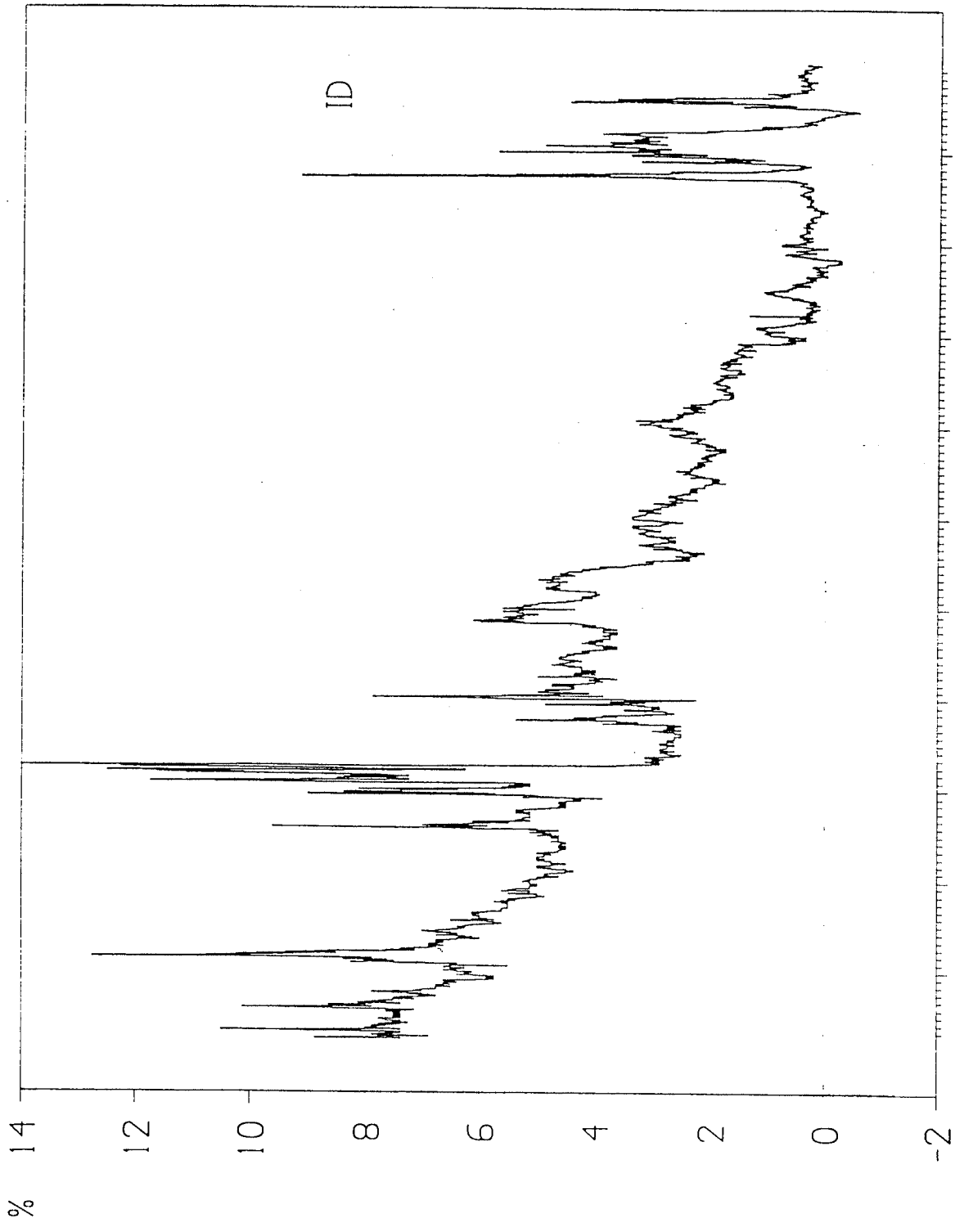


Fig. 1 1984 1987 1990 1993

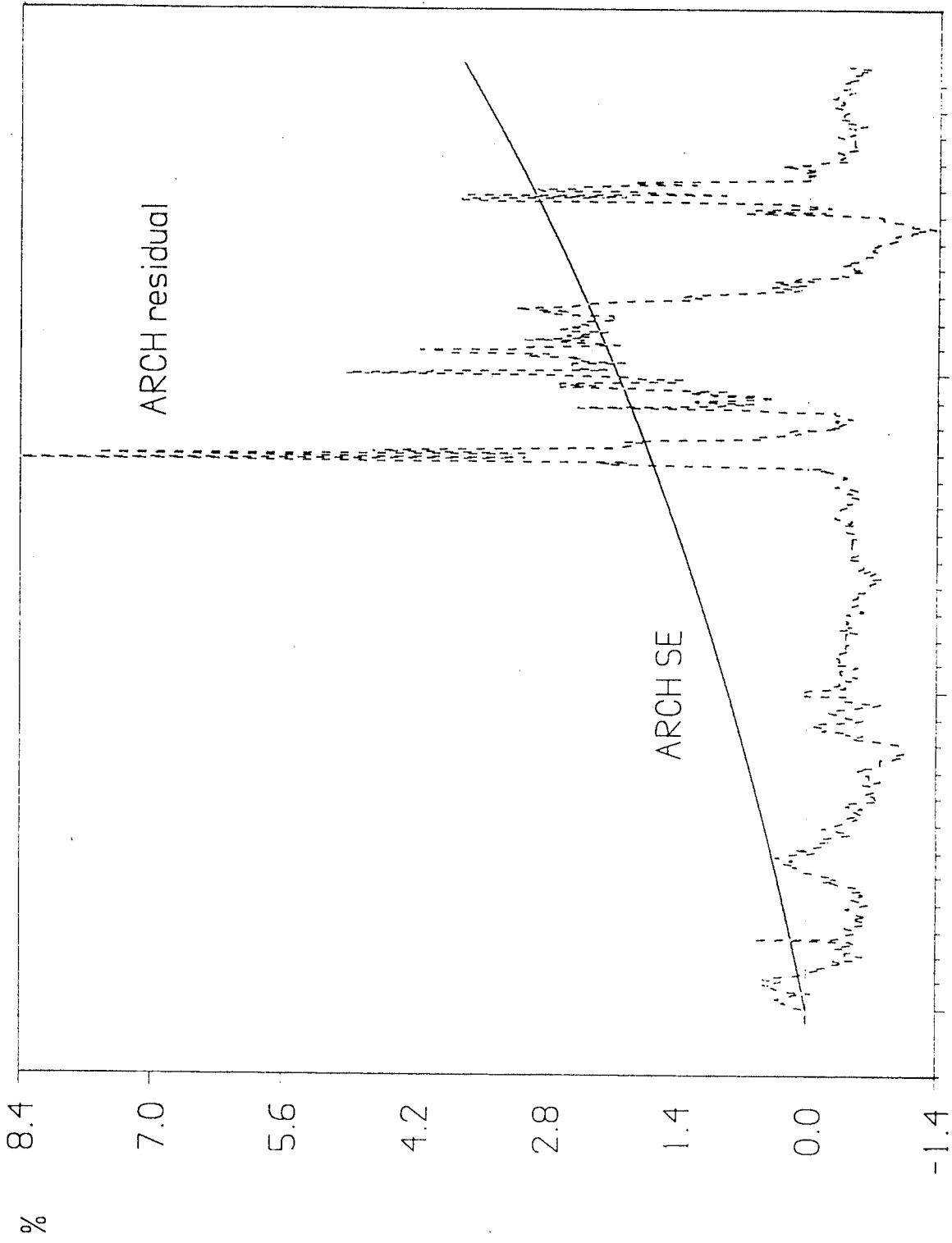


Fig. 2

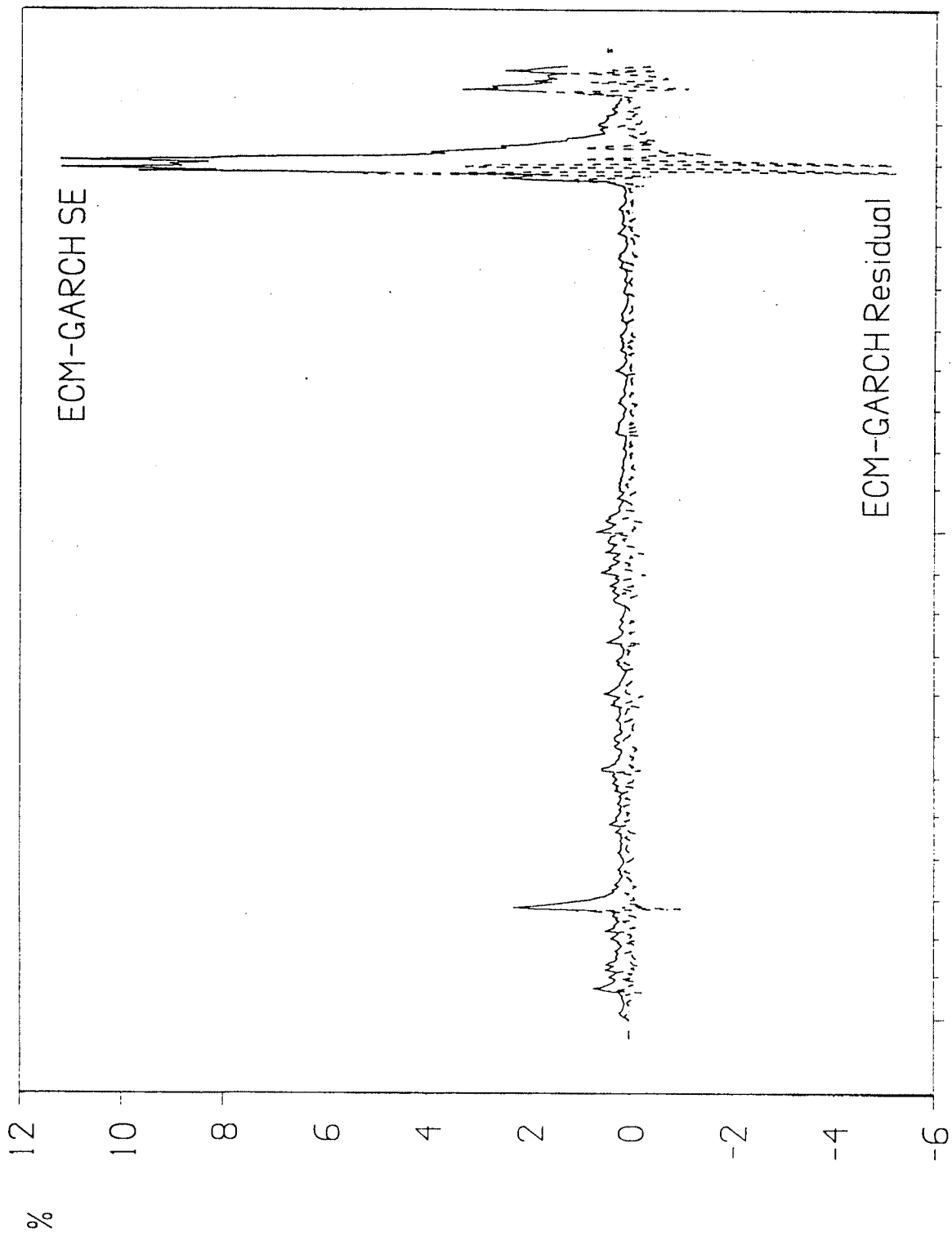


Fig. 3a

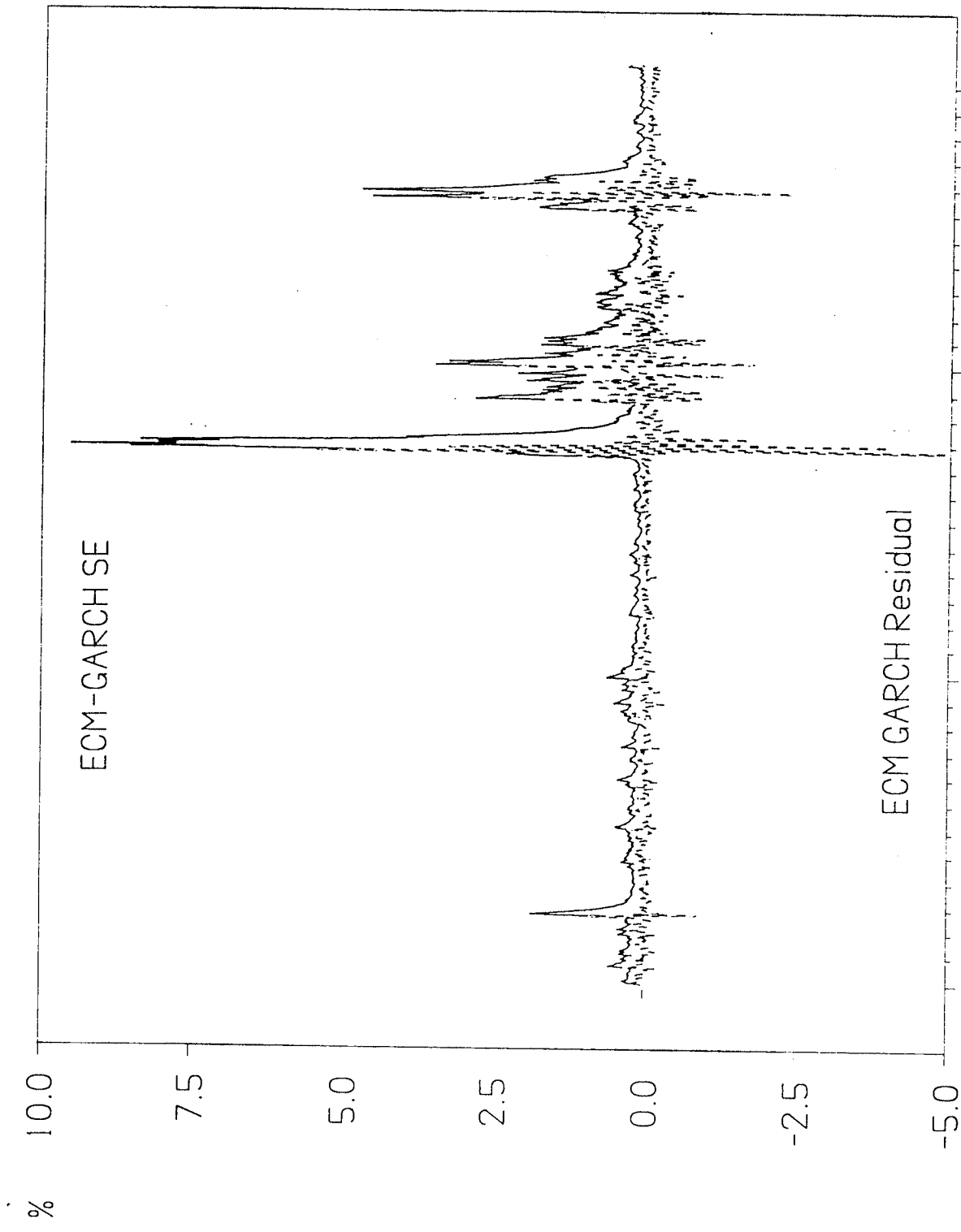


Fig. 3b

91

92

93

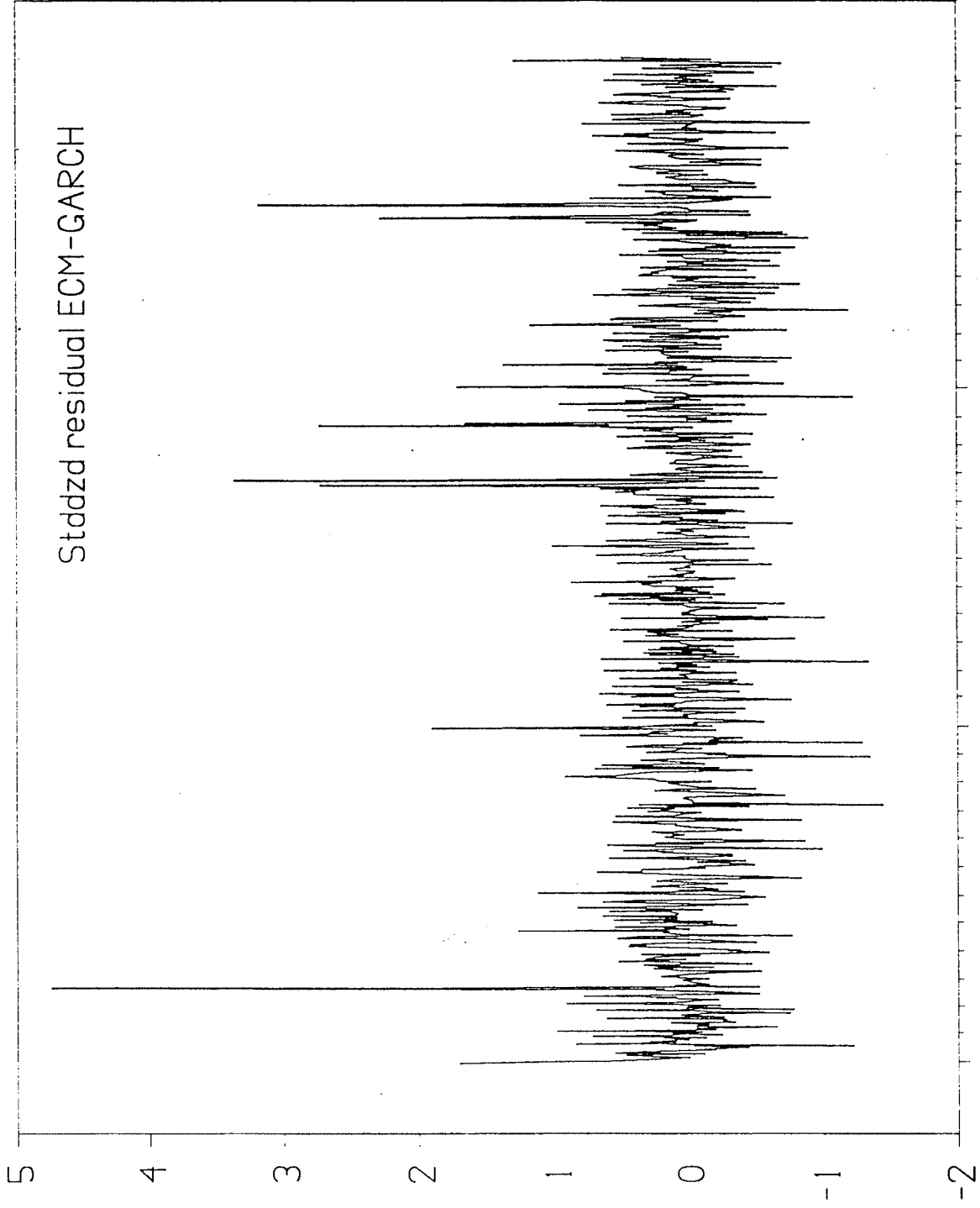


Fig. 4a 91 92 93

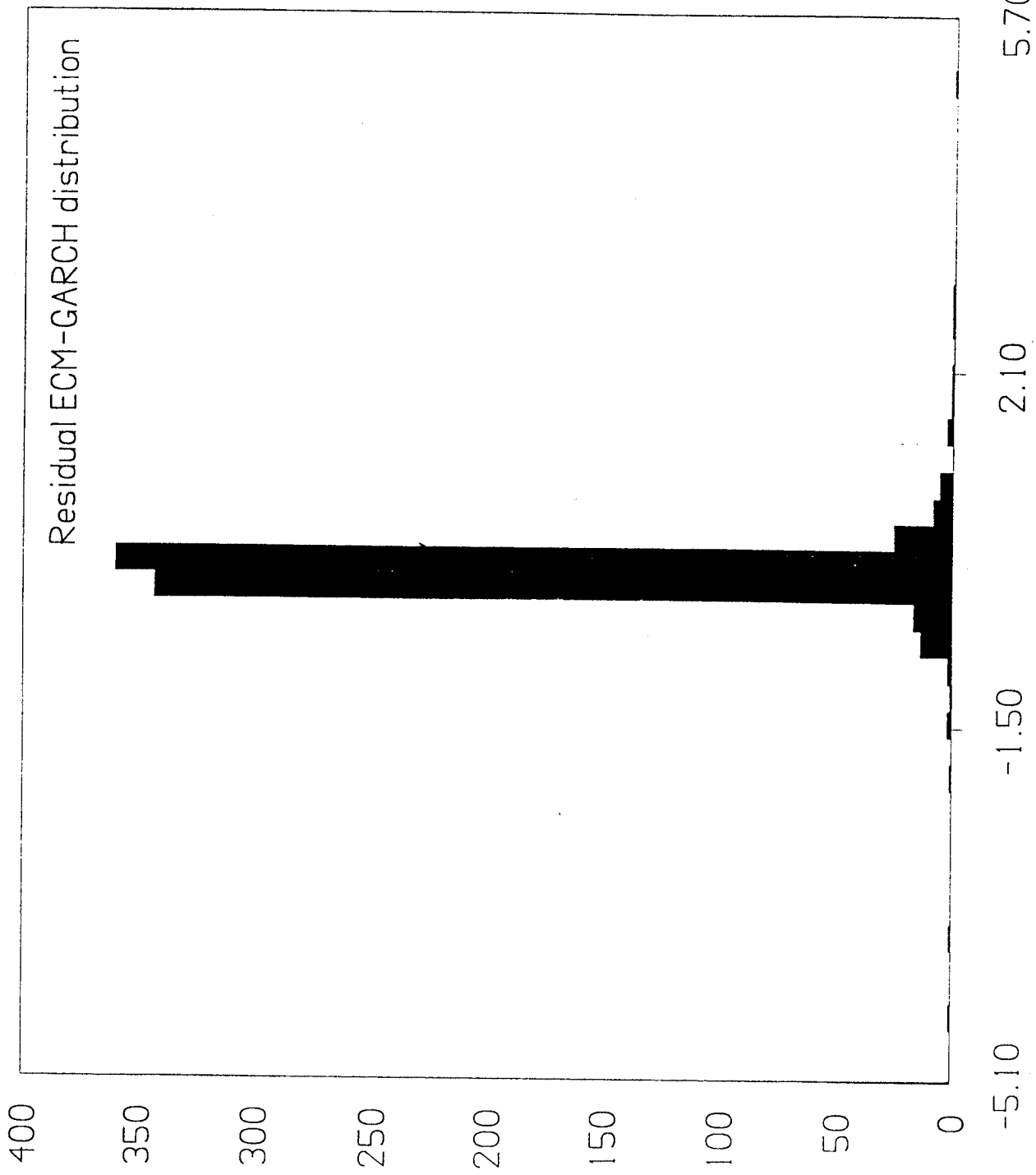


Fig. 4b

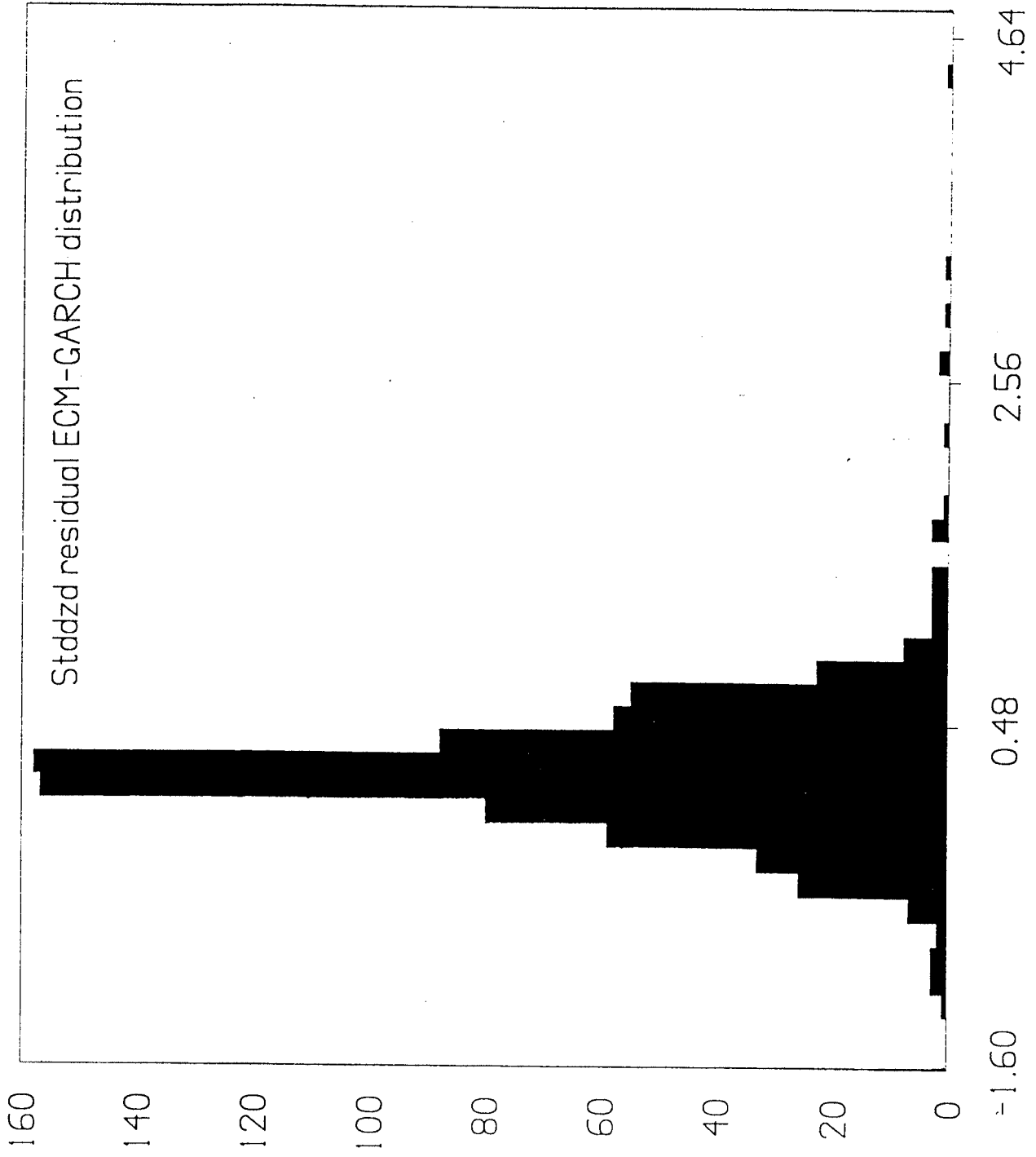


Fig. 4c

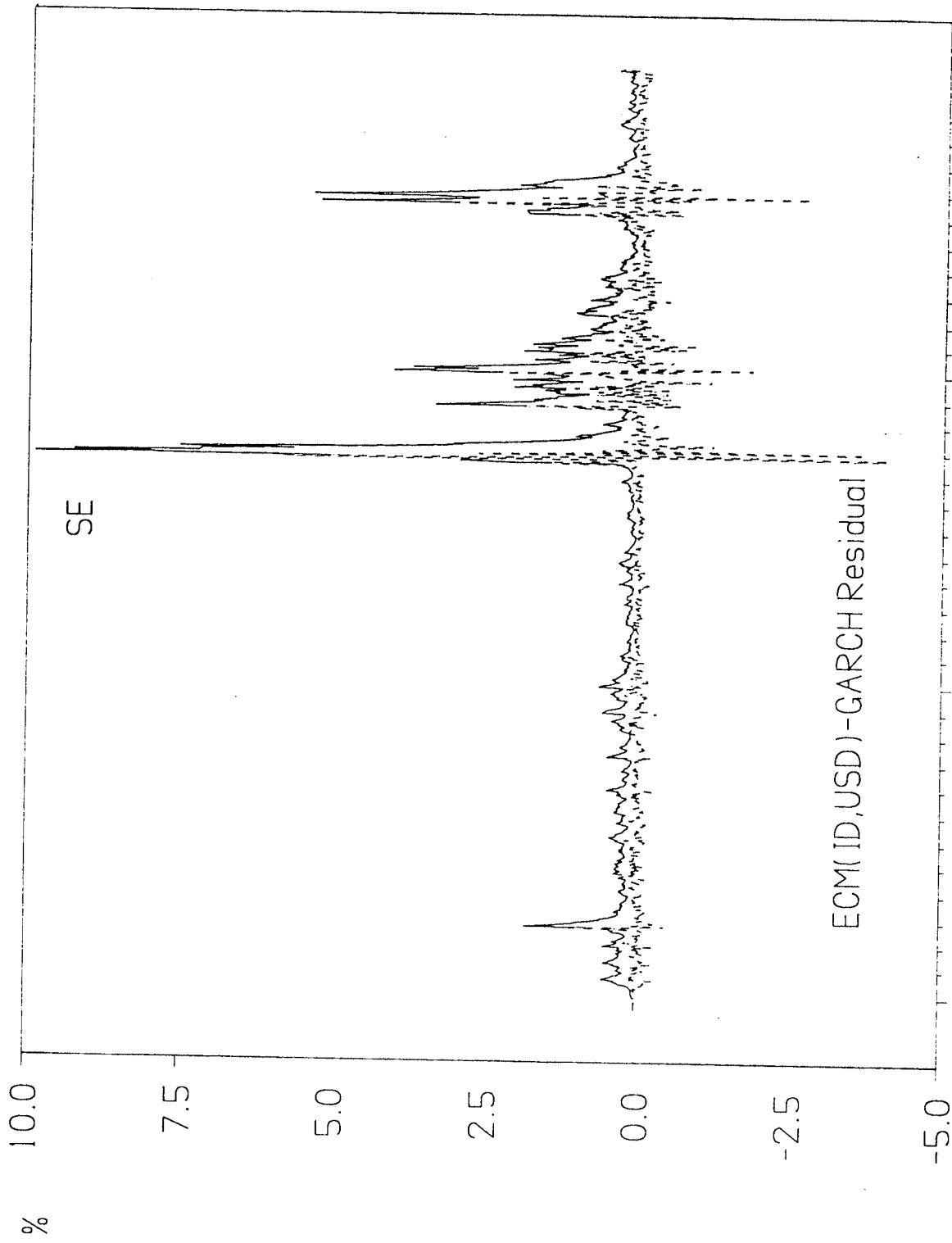


Fig. 5 91 92 93

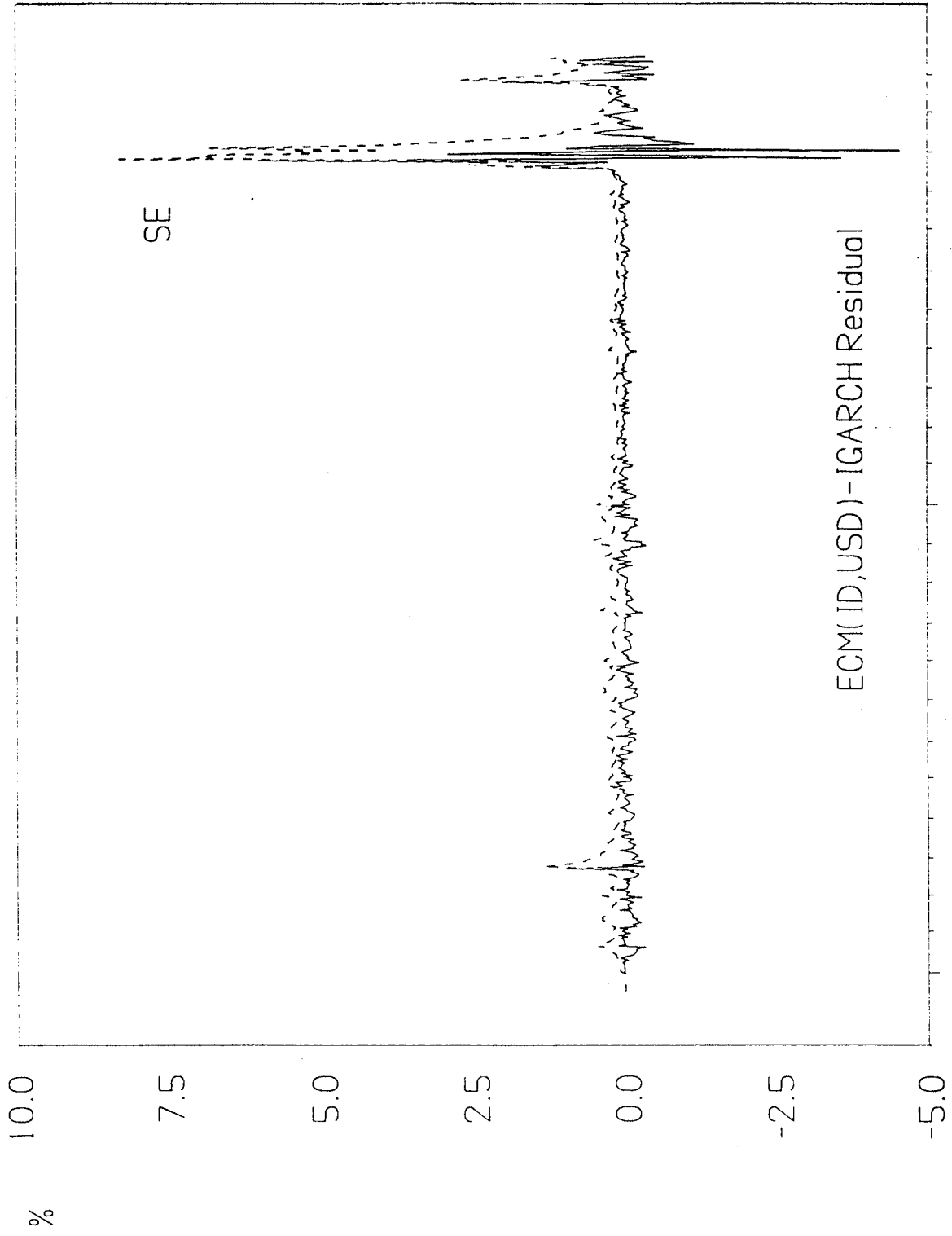


Fig. 6a

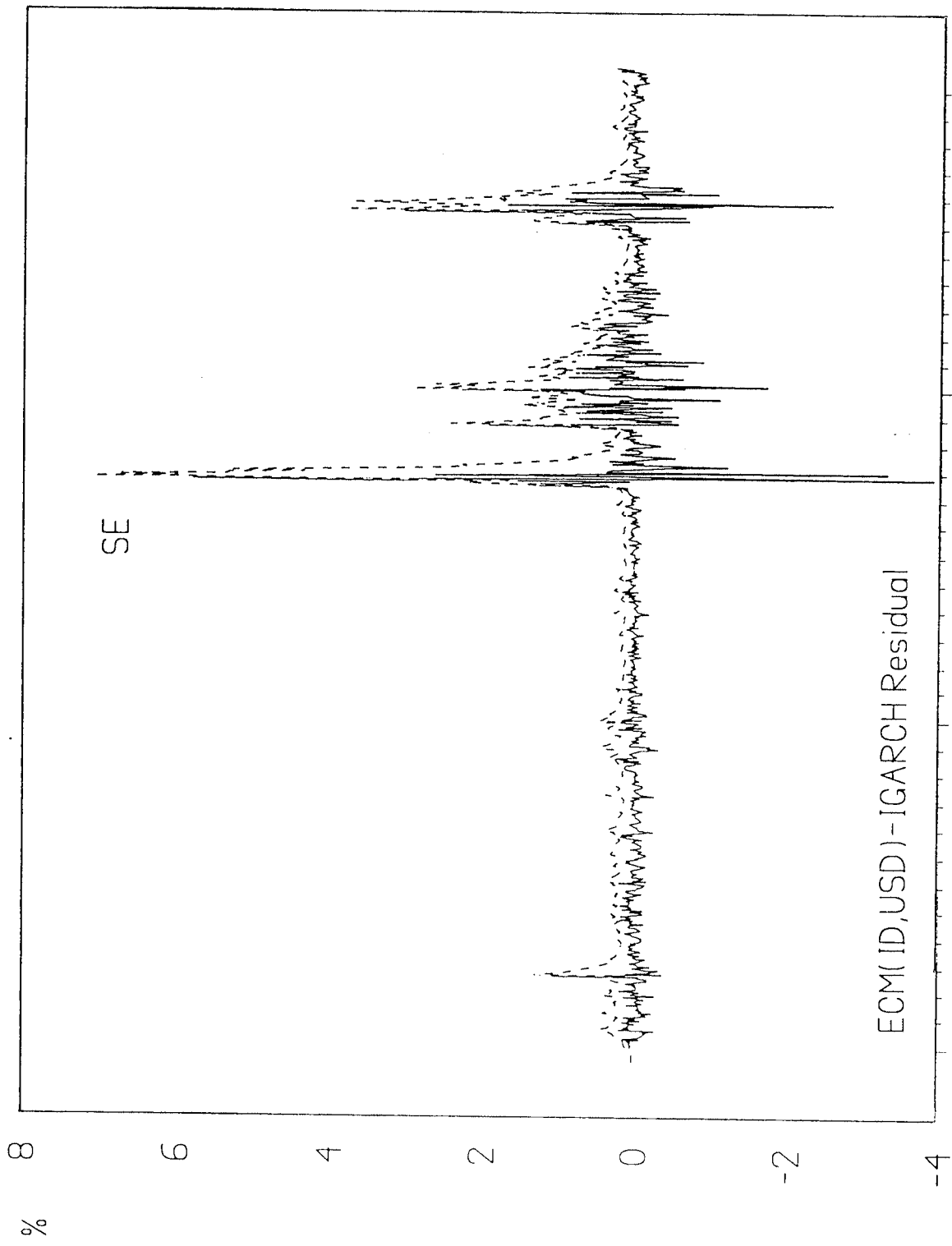


Fig. 6b 91 92 93

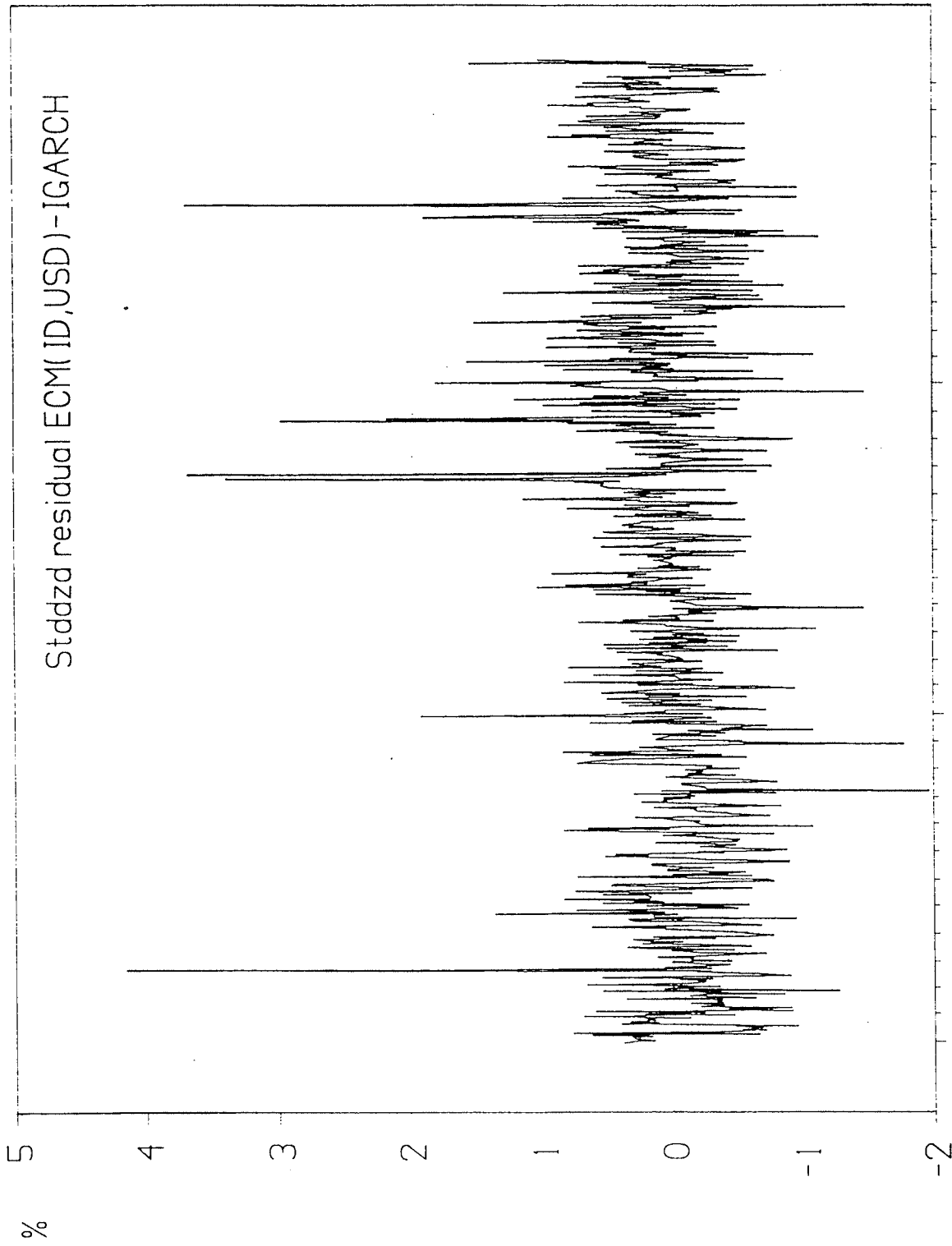


Fig. 7

