Exchange Rate Volatility: Asymmetric Effects

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Abstract

This paper studies the evolution of the daily exchange rates volatilities of five european currencies against the US dollar. The aim of this paper is to perform whether there are common factors in the evolution of these exchange rates flexibles during stability and unstability periods. Several alternative models have been proposed in the literature o to the model time varying volatilities. In this paper, we fit two parametric models, GARCH and GJR-GARCH for the years 1992 to 1993 and 1995 to 1997. We will show how these models within-sample estimates of volatility can be captured asymmetric effects of news, specially in periods with high speculation. Summarizing, we can conclude that these results have the attractive over the exchange rate flexible markets, particularly in the risk premium exchange rate manage.

1.0 Introduction

It is a well-known fact that the evolution of exchange markets is a basic and important variable of any economy. In fact, the exchange rate represents one of the most important indicators of an economy, as its the variations reflect, to a reasonable extent, movements in the international market, not only those of goods and services, but those of capital flows as well, affecting all of the other macro-economic variables. Furthermore, it also influences the behaviour of companies considerably, particularly if they operate abroad, whether in import/export or in investments or financing in foreign currencies.

As such, this paper aims to find some sort of structure that would explain the volatility of the exchange rate, which is generally considered to be an accurate measurement for quantifying exchange risks. In first place, we study of the Exchange Market for the peseta against the American dollar, this being the currency of reference in most economic and financial transactions. We do so from a dual perspective, considering both stable and turbulent financial environments, and we finally extend the analysis to other European currencies like the Italian lira, the French franc, the pound Sterling and the German mark as comparative references.

With such an aim, and basing ourselves on the literature previously published on the evolution of exchange rates, among which we should highlight those on the Theory of the Parity of Purchasing Power (Casel, 1919), the flow models (Meade, 1951; Fleming, 1962), and the asset models (Frenkel, 1976; Dornbusch, 1976); Branson, 1976, 1977). The general consensus is that these models are not very efficient in explanation. This is probably due to the fact that they consider macro-economic variables. This topic has generated a great deal of literature on theoretical models that try to reflect the empirical regularities observed in the evolution of exchange rates over time. This approach is based on the concept that the exchange rate is the price of a financial asset being traded on market (Mussa, 1979). Among such models, we should mention those based on the existence of bubbles (Evans, 1986); Messe, 1986); those that assume the existence of different types of agents in the markets (i.e., chartists and fundamentalists), (Frankel and F root, 1988), and those that seek to detect behavioural patterns in exchange markets. Among the latter group we find ARCH models of Conditional Volatility (Engle, 1982), models based on the theory of Chaos (Frank and Stengos, 1988), and others based on “target zones” (Krugman, 1991; Bertola and Caballero 1992).¹

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To summarise then, a great deal of the theoretical arguments forwarded by the above-mentioned studies focus on the possibility of predicting the variations in exchange rates, supposing that they have a lineal behaviour. It is quite possible, however, that these variations might be lineally un-correlated and, at the same time, show a “non-linear dependence”. As a result, the most recent innovations in Financial Econometric offer new modelling tools, among which we find parametric ARCH models (Engle, 1982) and GARCH models (Bollerslev, 1986) to facilitate the modelling and to predict the volatility of exchange rates when the non-linear predictable dependency in the variance is, in fact, linear. To do so, we carry out a two-fold study of the Spanish Peseta/US Dollar exchange market, comparing the period of monetary turbulence between 1992 and 1993 with one of relative stability, from 1995 to 1997, both of which proved to be of great importance to the exchange markets of the European Union. We also do a similar and parallel analysis in an extension to other European currencies as a comparative reference.

This article makes an important contribution to the data already presented in the above-mentioned studies. We do a novel form of analysis, considering the samples from two different perspectives, financial instability vs. stability. Secondly, the analysis is done employing models that reflect asymmetric effects of the price-volatility relationship, as well as the effect of such conditional volatilities on the mean returns.

The paper is structured as follows: In Section 2, we explain the data employed and present an analysis of the behaviour of the returns considered. Section 3 gives a detail of the GARCH(p,q) model developed by Bollerslev (1986) and the GJR-GARCH model presented by Glosten, Jaganathan y Runkle (1993), in an effort to verify whether behaviour of the distribution of the exchange rates considered is strictly different between the two samples. The results of this analysis are presented in Section 4. We verify the existence of a trade-off between profits and risk, following methodology of Engle, Lilien y Robins (1987), which allows us to detect certain “common” factors in their behavioural patterns. Finally, in Section 6, our conclusions are expressed.

2.0 The Data

The database we employ is compiled from daily data on the spot exchange rate for the Spanish Peseta/US Dollar (PTE/USD), the Italian Lira/US Dollar (LIT/USD), the French Franc/US Dollar (FRF/USD), the Pound Sterling/US Dollar (GBP/USD) and the German Mark/US Dollar (DM/USD). The sample periods run from 2 January 1992 to 31 December 1993 (483 observations), and from 2 January 1995 to 30 December 1997 (730 observations). The data is taken from the Bank of Spain’s database (Servicio de Estadística y Central de Balances del Banco de España).²

Following a common practice in modelling financial series, the analysis we present considers that the data represented are the returns calculated as the first differences in the original series:

$$Y_t = (I - L)S_t \quad (1)$$

where S_t is the value of the spot exchange rate at moment t (expressed in logarithms) and Y_t is the increment on day t .³

2.1 Behaviour of Returns

In Charts 1 and 2 of the appendix, the main statistics on the yields series for the sub-sample 1992-93 and 1995-97 are presented separately, all expressed in logarithms. From the descriptive statistics on the yields we observe, as is usually the case in most financial series, that the mean is approximately equal to zero. The LIT/USD exchange rate showed the highest increase in yield (depreciations) during the sub-sample 1992-93 (7.1%), in contrast to 5.4% for the PTE/USD and DM/USD exchange rates, and 3.2% and 2.4% for the FRF/USD and GBP/USD rates respectively. In the sub-sample 1995-97 the Italian lira again suffered the greatest depreciation against the dollar, although considerably less than it had during the period of turbulence, between 1992 and 1993.

However, the falls in the daily yield rates (appreciations) are more significant during the 1992-93 period for the DM/USD exchange rate at 5.8%, followed by the GBP/USD rate, at 4.6%, in contrast to 3.8% for the PTE/USD,

3.2% for the LIT/USD and 2.5 for the FRF/USD exchange rate. Likewise, during the period 1995-97 it was again the German mark/US dollar rate that had the greatest percentage appreciation.

Our analyses show, therefore, that the DM/USD exchange rate had the greatest typical deviation in yield, (followed by LIT/USD), which clearly reflects its higher volatility, particularly during the period of financial uncertainty (1992-93). In any event, although the typical deviation is one of the measurements of variability most commonly used in the literature, this is merely a sample measure of the dispersion of a random variable, with no special structure being revealed for the evolution of the series of either the yields or the volatility.

From the analysis of the distributions of the yield we can see that the Jarque-Bera statistic clearly rejects the hypothesis of normality in the distribution, in all of the cases. Furthermore, the distributions are all leptocurtic, particularly in the case of the more volatile rates, (i.e., the Italian Lira and the German Mark against the Dollar). The same sort of analysis was done for the period 1995-97, and although the results vary in certain details, the German mark, however, continues to show a positive bias. The Ljung-Box statistic show that the auto-correlation coefficients of the retards of the tenth order are not significant, except in the case of the Spanish peseta and the German mark in the sub-sample of the period 1992-93, and the Italian lira in the sub-sample of the period 1995-97. On the contrary, the non-correlation hypothesis is consistently rejected for the series of the squares, which shows the existence of auto-regressive components. Furthermore, while the series analyzed do not seem to have any changes in their mean levels over time, the series of the squares (i.e., the volatilities) show correlations that are significantly different from zero for even high retards.

In summary, the results of the analysis of the statistics of the series, with leptokurtosis in the logarithmic differences of the exchange rates considered, as well as high auto-correlations in their volatilities, indicates that there is heteroskedasticity in the dynamics of series. As such, if the form of this non-linearity in the variance is known (linear or non-linear), the ARCH-type processes are an appropriate means for modelling it, as we show in the following section.

3.0 Volatility Models

In this section, we analyze the behaviour of the volatility of the daily yield of the exchange rate of the peseta and other currencies of the European Union against the American dollar. We do so by modelling moments of the second order - Empirical Models (Conditional Variance Models) – which are based on the assumption that the uncertainty associated with the prices, measured by the variances and covariances, changes over time. As such, we present an approach to the non-linearity of the variations in the exchange rates considered here (non-linearity in variance) from a parametric focus. To be more specific, we use the ARCH-type model (Autoregressive Conditional Heteroskedasticity) introduced by Engle (1982).

The ARCH(q) model proposed by Engle (1982) supposes that the variance in the process of the interference will vary over time, so that the conditional variance is a linear function of past errors squared, and as such, if the errors of the past are large, the variance will also be. The ARCH is certainly the model

that has been most frequently used in modelling the volatility of financial series and has been the centre of many debates over the last decade, from both the theoretical and the econometric point of view.⁴

The basic formula for this sort of modelling is to consider a y_t series (from $t=1$ to $t=T$), as a sequence of independent and identically distributed random variables, ε_t , with a unitary variance, multiplied by a factor σ_t , denominated *volatility*⁵. In other words:

$$y_t = \varepsilon_t \sigma_t; \quad \varepsilon_t \approx IIN(0, 1) \quad (2)$$

Over the last few years, there have been many alternative specifications proposed for measuring volatility, most of which are based on the Engle (1982) model, such as the GARCH, the EGARCH, the logGARCH, the GARCH-N, and the T-GARCH models, among others.⁶ In this section, we focus on the GARCH(p,q) model proposed by Bollerslev (1986), as it is one of the most frequently used models in the empirical literature, and on the GJR model developed by Glosten, Jaganathan y Runkle (1993), as they allow us to include the leverage effect, which appears in most financial series.

The GARCH model developed by Bollerslev (1986) considers that the conditional variance, σ_t^2 , depends not only on the previous innovations but also on its own previous values, as reflected in the following generic expression:

$$\begin{aligned}\sigma_t^2 &= \omega + \sum_{i=1}^q \alpha_i y_{t-i}^2 + \sum_{j=1}^p \beta_j \sigma_{t-j}^2 \\ &= \omega + A(L) y_t^2 + B(L) \sigma_t^2\end{aligned}\quad (3)$$

where $p \geq 0$, $q \geq 0$, and $A(L)$ and $B(L)$ are their corresponding retard operators.

In the GJR-GARCH model (Glosten, Jaganathan y Runkle, 1993), the conditional variance is given by the following equation:

$$\sigma_t^2 = \omega + \sum_{i=1}^q \alpha_i y_{t-i}^2 + \sum_{j=1}^p \beta_j \sigma_{t-j}^2 + \sum_{i=1}^q \gamma_i y_{t-i}^2 S_{t-i}^- \quad (4)$$

where S_{t-i}^- is a dummy variable that has a value of 1 when the yield is negative and 0 otherwise. The impact of the news, or shocks to the yield on the volatility depends on the sign of the parameter estimated by this dummy variable. In any case, this model, which allows for a different responses of the volatility to positive or negative shocks, supposes that the minimum volatility is observed when there is no news. Before estimating the models proposed, we have tested ARCH effects with the Lagrange test (LM), which the null hypothesis is the absence of heteroskedasticity and the test used is of the LM (ARCH) type.⁷ The results of the Engle test for the sample periods 1992-1993 and 1995-1997 are shown in Tables 3 y 4 of the Appendix, respectively, and they confirm the existence of ARCH effects.

4.0 Results

The results of the estimate for maximum likelihood of the GARCH(1,1) and of the GJR(1,1) models, in general, are quite similar. For this reason, and to facilitate the presentation and interpretation of the results, we have decided to relegate to Charts 3 and 4 of the appendix the estimates of the GJR modelling if the parameter of asymmetry is statistically significant and the logarithm of the function of likelihood is either greater or very near to that of the GARCH. Otherwise, the estimates derived from the GARCH modelling of the variance are presented. These charts also include the results of the logarithm of the function of likelihood, the Schwarz statistic, the Mean Standard error and the Wald test of combined significance of the parameters in the conditional variance estimated.

In view of the results of the estimate, the following considerations should be highlighted. First, that the results obtained from the models proposed for the estimated conditional variance of the daily returns of the PTE/USD exchange rate, as well as for all the other exchange rates considered in this analysis, show remarkable precision in estimating the auto-regressive effect, which indicates strong inertia in the behaviour of the volatility. While this is not too surprising, as we are working with daily data, it is nevertheless noteworthy that, for the case of the peseta, this volatility estimate is 0.733 for the 1992-1993 sample and reduces considerably during the 1995-1997 period of stability, to just 0.600. This greater inertia in the volatility for the period of considerable turbulence in the exchange markets of the European Union could be due to a greater influence of speculative factors during this period, and others of a more structural type during the 1995-1997 period, with an absence of any outstanding financial turbulence.

We should also point out that in all of the exchange markets analyzed, the shocks to the volatility are persistent, since the sum of the estimated coefficients α and β is close to one, indicating a very slow response from the volatility to an unexpected event in those markets. These results are even stronger for the 1992-1993 sample (except for the LIT), which is in keeping with the previously expressed line of reasoning. The greater persistence of the volatility of the German mark versus the rest of the exchange rates during the 1992-1993 period is noteworthy, and its relative drop during the 1995-1997 period is probably justified by the leading role of the German mark within the European Union.

Regarding the parameter that represents the GJR model, it only proves to be statistically significant for the period of turbulences between 1992 and 1993, (existence of a leverage effect), and thus, shows an asymmetric response of the volatility to both good and bad news. Is it interesting, however, that for this sample, although the impact of the news on the volatility is positive, independently of its sign, in our analyses the negative sign observed for γ in the PTE and the DMK rates, indicate that the negative values from the unexpected element (an appreciation of the PTE/USD and DM/USD rates) cause a smaller increase in the volatility than the positive values (a depreciation in the PTE/USD and the DM/USD rates) of the same size do. This negative sign seems to indicate, therefore, that the risk increases more when the yield on these exchange rates are lower than expected than when they are higher than expected. The reverse interpretation is then true for the positive signs of the FRF/USD and the GBP/USD exchange rates. The negative values (an appreciation) from an unexpected event generate greater increases in the volatility than positive values (a depreciation) of the same size do.

To summarise this section then, it should be remembered that our aim was to analyze the volatility seen in the yields of the exchange rates for a number of EU currencies against the US dollar, the Spanish peseta included. To do so, we followed their evolution through two clearly different periods: one of financial uncertainty in the European markets, between 1992 and 1993, and another of considerable stability, between 1995 and 1997. We have subsequently noted a couple of important characteristics about the volatility of the market. First, its great variability over time, in absolute terms, being particularly high for the 1992-1993 sample, and especially so for the DM/USD and the LIT/USD rates. Secondly, the auto-regressiveness of the volatility on the previous day which shows its remarkable persistence, particularly in the 1992 sample, which is probably attributable to peculiar characteristics of a market in which speculation is considerably high. Finally, the fact that evidence was found of a leverage effect that is only statistically significant for the 1992-1993 sample.

From the evidence obtained, we can conclude that the behaviour of the volatility of the exchange rates examined, and for the sample periods considered here, is variable over time and can be accurately modelled with ARCH models. Furthermore, the GJR-GARCH models seem to be comparatively better at explaining and predicting high volatilities (the 1992-93 period in our case) since the asymmetric effects of the unforeseen events are statistically significant. For the 1995-1997 sample, on the other hand, the GARCH model proves to be more appropriate.

Therefore, considering that the volatility is predictable in both samples to a certain extent, we do a complementary analysis in the following section, in an effort to obtain some sound information about the extent to which the volatility, measured by the conditional variance, can explain the rate of the yields of the exchange rates considered here, and particularly that of the peseta/dollar exchange rate. To do so, we now present the ARCH-M model, which proves to be a useful econometric tool for analyzing the influence that the variable risk which is variable over time exerts on the risk premiums requested by the agents who operate in these markets.

5.0 Returns-Risk Analysis

Based on the work of Engle, Lilien and Robins (1987), the empirical literature has tried, on many occasions, to verify the existence of a trade-off between returns and risk, with rather contradictory results. As such, we shall now seek some significant relationship between conditional variance and the conditional expectations of returns for the set of exchange rates considered here for the sample periods 1992-1993 and 1995-1997.

To do so, we use the auto-regressive mean conditional volatility model in mean (Glosten, Jaganathan and Runkle, 1993), to capture the asymmetric effects of the shocks on the volatility, and on the yield of the 1992-93 sample, while for the 1995-97 sample we use the GARCH(1,1)-M as these values are not in this case.

Formally, the model proposed is based on the supposition that there is a process $\{Y_t\}$ which can be expressed in the following way:

$$Y_t = \mu_t + \varepsilon_t \quad (5)$$

where μ_t is the conditional mean of the series:

$$\mu_t = E_{t-1}(Y_t) \quad (6)$$

and which also supposes that the conditional mean of the series is an increasing function of conditional variance σ_t of the disturbance ε_t which we can express as follows:

$$\mu_t = \zeta + \Psi \sigma_t \quad \Psi > 0 \quad (7)$$

σ_t having been previously modelled with a GARCH or GJR-type model.

Following the example of Peña (1993), the parameter Ψ is interpreted as the coefficient of a market agent's aversion to risk. Consequently, the excessive yield on a given stock, as expression [7] states, can be considered as a premium on risk that is variable over time.

Chart 5, in the Appendix, shows the results from the different models estimated for the exchange rates examined here. From that analysis we have found evidence of significant effects being exerted on the conditional expectations by the conditional variance for the LIT/USD and the PTE/USD exchange and particularly so during the 1992-93 period financial instability. In this sub-sample we also observe that the effect of the conditional volatility of the above-mentioned exchange rates have on the mean return is positive. This means that the increases in the conditional volatility increase the variations in the mean yield on the long term and, as such, the risk of these currencies depreciating with respect to the Dollar. These results are, in general, in keeping with others previously published.⁸ Nevertheless, the results we obtain for the entire set of exchange rates for the 1995-1997 sample, the Peseta included, do not show any significant effects on the conditional media from the conditional variance.

Finally, with the aim of concluding this study with a useful contribution of relevant information regarding the reliability of predictions based on the regularities discovered here, as well as to provide some additional empirical value to the analyses, we have tested our method's capacity to predict with Theil's coefficient, which compares the predicting power of the GARCH/GJR-M models, with an ingenuous random walk model. As can be seen in Chart 5 in the Appendix, the value of the Theil coefficient is less than one, which indicates the satisfactory level of predicting power in the models proposed here, even if the results are not entirely satisfactory as the values are all close to one.

We believe that the results presented here should be of great interest to institutions and/or companies who invest or trade in the international currency-exchange markets where speculation and maximum protection against risk are paramount. On this point, we should like to refer readers to the work of Glen y Jorion (1993), from which it is deduced that such coverage to reduce the risk is only profitable if it is not accompanied by a trade-off in reduced returns.

6.0 Conclusions

Throughout this study, we have endeavoured to analyze the predictable behaviour of the volatility through the use of ARCH-type processes. The idea behind the study was to try and analyze both the common and the peculiar features of the volatility, using two contrasting sub-samples, a period of financial turbulence within the European Union,

from 1992 to 1993, and another sample period of considerable financial stability between 1995 and 1997. Our results should be of great interest to any firm or institution that trades or invests in international markets with flexible currency exchange rates, in which the management of risk is paramount.

With the estimating of GARCH(1,1) and GJR-GARCH(1,1)-type models, we have obtained some noteworthy results. In the first place, that the volatility of exchange rates depends on its past record and is therefore, to a certain extent, predictable. Our analyses also show that the GJR-GARCH-type model is especially appropriate for the 1992-93 period, in which there is clear evidence of asymmetric values in the relationship between shocks and volatility, indicating that the model captures the non-linearity of the data quite accurately. For the 1995-97 period, however, the GARCH proves to be more appropriate. This seems to confirm the belief that the ARCH-type model is comparatively better at explaining and predicting high levels of volatility, as in the 1992-93 sample period in our analysis.

The accuracy with which the auto-regressive effect is captured in both samples is also remarkable, and particularly so in the period of great turbulence, reflected in the high t-ratio of the parameter associated with the conditional variance of the previous period. The persistence of the volatility also proves to be much stronger in the 1992-1993 period, and is particularly so for the German Mark/US Dollar exchange rate during the 1992-1993 period. Furthermore, it is precisely this exchange rate that shows the least volatility during the 1995-1997 sample period, which is probably logical if we consider the leading role of the German Mark among the European Union currencies.

With the advent of the Euro as the common currency among the member countries of the European Union, and particularly in the current setting of the steady globalization of financial activities and the consolidation of a single market for financial services in Europe, we feel that an interesting analysis for future research would be the regularities that exist in the behaviour of the Euro, during periods of relative stability and in others of relatively high speculation, in comparison with the two strongest currencies, the US Dollar and the Japanese yen, which would also contribute to strengthening the results of our analyses reported here.

Without a doubt, the results obtained from our study demonstrate the relevance of the variable of volatility, and specifically, with the use of the ARCH family of models, in identifying and explaining highly speculative processes. This affords a certain degree of predictability and anticipation of crises in currency exchange markets, and facilitates a better management of the exchange risks to which economies in general and companies in particular are exposed.

Finally, we could hardly close this paper without indicating the possible approach that future studies along these lines, such as an attempt to improve on the model presented here for the prediction of the volatility of exchange rates, bearing in mind that the foreign exchange market is a totally international market and, thus, interferences in one market often spill-over into another. As such, it would be interesting to do a similar analysis to the presented here, considering that variations in exchange rates not merely oscillations in nominal terms, but in real terms as well, in other words, after corrections have been made for inflation.

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Footnotes

¹ For more details on the existing literature, see, Bajo and Sosvilla (1993), and Frankel and Rose (1995), among others.

² The variables studied here have not been corrected by intervention analysis.

³ On this point, remembering that stationarity is a necessary condition for analysing and/or modelling time series, the Dickey-Fuller unitary root test (ADF) and Phillips-Perron (P-P) test were previously applied to the exchange rates

considered here. The results obtained shown that the random walk model in approaching the evolution of the series, as it has shown that they have an integrated term, while their first differences are, indeed, stationary.

⁴ In considering the arguments in favour of this type of modelling, it is important to remember that ignoring the presence of conditional heteroskedasticity leads to a loss of efficiency in the estimation of the parameters of the model and, furthermore, if the model includes lags of the endogenous variable as explanatory variables, it can even mean that the estimates of the standard errors of the parameters estimated may not be consistent. For a more detailed theoretical analysis of this sort of modelling see the original article by Engle (1982) and Bollerslev, Engle and Nelson (1994) also provide a comprehensive theoretical analysis.

⁵ The volatility is a positive function that is variable over time and is measurable with respect to $\Psi_{t-1} = \{y_1, \dots, y_{t-1}\}$, the set of information available at instant $t-1$.

⁶ A detailed analysis of these models can be found in Bera y Higgins (1993) op. cit., and Shepard (1996), among others, where, apart from the descriptive analysis, they also supply specific references for each case.

⁷ This test is asymptotically equivalent to using TR^2 , where R^2 is the determining coefficient from the auxiliary regression of the ordinary least squares residuals squared ($\hat{\varepsilon}_t^2$) on a constant and p lags of itself, all obtained under the null hypothesis of homokedasticity. In other words, it is based on estimating the following regression:

$$\hat{\varepsilon}_t^2 = \alpha_0 + \alpha_1 \hat{\varepsilon}_{t-1}^2 + \alpha_2 \hat{\varepsilon}_{t-2}^2 + \alpha_3 \hat{\varepsilon}_{t-3}^2 + \dots + \alpha_q \hat{\varepsilon}_{t-q}^2$$

and if there were no ARCH or GARCH effects, the estimations of the coefficients $\alpha_1, \alpha_2, \dots, \alpha_q$ would not be significant. It must be remembered that, under the null hypothesis of the errors not following an ARCH or a GARCH model, it can be verified that the statistic TR^2 converges to a χ_q^2 , which allows us to carry out a normal parametric test. The validity of this test is quite independent of the functional form of the conditional variance (σ_t) whenever it is a function of the lagged disturbances squared.

⁸ See Marco (1998), among others.

Appendix

**TABLE 1: Empirical Regularities On The Returns
(1992-1993)**

	PTE	LIT	FRF	GBP	DM
Mean	0.007	0.0007	0.0002	-0.0004	0.0003
Maximum	0.054	0.071	0.032	0.024	0.054
Minimum	-0.038	-0.032	-0.025	-0.046	-0.058
Std. Dev.	0.0089	0.0092	0.0079	0.0085	0.0093
Skewness	0.7959	1.3479	0.5751	-0.7465	-0.2263
Kurtosis	7.9252	11.4192	4.9028	5.7090	10.5985
J-B	538.09	1169.55	99.29	192.16	1163.69
(p-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Q(10)	19.152	7.889	8.323	10.653	17.442
(p-value)	(0.038)	(0.640)	(0.597)	(0.385)	(0.065)
Q(20)	34.765	20.322	16.159	20.971	26.379
(p-value)	(0.021)	(0.438)	(0.707)	(0.399)	(0.154)
Q₂(10)	69.719	47.585	87.516	37.975	23.299
(p-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Q₂(20)	84.016	53.403	114.31	58.686	28.540
(p-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.097)

Notes:

- (i) J-B is the Jarque-Bera test, null hypothesis of normality.
- (ii) Q(10) and Q(20) denotes the Ljung-Box statistic for the first k autocorrelations. $Q^2(10)$ and $Q^2(20)$ denotes the Ljung-Box statistics of (returns)² for the first k autocorrelations

TABLE 2: Empirical Regularities On The Returns
(1995-1997)

	PTE	LIT	FRF	GBP	DM
Mean	0.003	0.0001	0.0002	$8.3 \cdot 10^{-05}$	0.0002
Maximum	0.036	0.055	0.029	0.024	0.031
Minimum	-0.034	-0.024	-0.038	-0.018	-0.035
Std. Dev.	0.0064	0.0056	0.0058	0.0047	0.0064
Skewness	-0.0027	1.2013	-0.4832	0.0056	-0.4363
Kurtosis	7.1614	16.7777	7.5937	5.3566	7.1006
J-B (p-value)	526.01 (0.000)	5940.99 (0.000)	669.35 (0.000)	168.69 (0.000)	531.89 (0.000)
Q(10) (p-value)	10.284 (0.416)	21.382 (0.019)	14.072 (0.170)	8.641 (0.566)	10.508 (0.397)
Q(20) (p-value)	32.320 (0.040)	49.526 (0.000)	28.219 (0.104)	15.138 (0.768)	27.556 (0.120)
Q ₂ (10) (p-value)	69.739 (0.000)	61.989 (0.000)	94.631 (0.000)	48.927 (0.000)	72.037 (0.000)
Q ₂ (20) (p-value)	97.121 (0.000)	117.45 (0.000)	107.81 (0.000)	55.114 (0.000)	99.678 (0.000)

Notes:

- (i) J-B is the Jarque-Bera test, null hypothesis of normality.
(ii) Q(10) and Q(20) denotes the Ljung-Box statistic for the first k autocorrelations. Q²(10) and Q²(20) denotes the Ljung-Box statistics of (returns)² for the first k autocorrelations

TABLE 3: Garch (1,1)-Gjr (1,1) Models: Maximun Likelihood Estimation
(1992-1993)

$$Y_t = \varepsilon_t, \quad \varepsilon_t / \psi_{t-1} \sim N(0, \sigma_t^2), \quad \sigma_t^2 = \omega + \alpha \varepsilon_{t-1}^2 + \beta \sigma_{t-1}^2 + \gamma \varepsilon_{t-1}^2 S_{t-1}^-$$

	PTE	LIT	FRF	GBP	DM
ω	$1.02 \cdot 10^{-05}$ (3.8204)	$7.06 \cdot 10^{-05}$ (2.1530)	$2.45 \cdot 10^{-06}$ (1.4967)	$6.49 \cdot 10^{-06}$ (2.3462)	$9.86 \cdot 10^{-06}$ (4.4579)
α	0.21633 (4.3074)	0.1165 (4.1753)	0.0363 (1.7231)	0.0566 (1.4444)	0.3065 (4.9364)
β	0.73321 (12.0786)	0.7979 (12.2624)	0.8747 (20.0524)	0.7996 (15.3289)	0.6871 (14.3835)
γ	-0.20480 (-4.2709)		0.0975 (2.8875)	0.0904 (1.8619)	-0.2820 (-4.8795)
Log.Likelihood	1626.275	1606.308	1672.408	1633.309	1605.639
Schwarz	-9.3655	-9.3323	-9.6141	-9.4706	-9.2877
MSE	$4.1 \cdot 10^{-08}$	$3.8 \cdot 10^{-05}$	$3.9 \cdot 10^{-08}$	$3.6 \cdot 10^{-08}$	$4.3 \cdot 10^{-08}$
w ₁ (p-value)	440.372 (0.000)	21.816 (0.000)	2.253 (0.107)	3576.53 (0.000)	34.852 (0.000)
Engle (p-value)	8.394 (0.004)	3.426 (0.065)	12.338 (0.001)	3.037 (0.098)	10.354 (0.001)

Notes:

- (i) Maximun Likelihood Estimation of the parameters. t-statistics in parentheses.
(ii) Wald-Test: w₁ null hypothesis of $\alpha=\beta=\gamma=0$ on variance.
(iii) MSE is the estandard error.

TABLE 4: Garch(1,1)-Gjr(1,1) Models: Maximun Likelihood Estimation (1995-1997)

$$Y_t = \varepsilon_t, \quad \varepsilon_t / \psi_{t-1} \sim N(0, \sigma_t^2), \quad \sigma_t^2 = \omega + \alpha \varepsilon_{t-1}^2 + \beta \sigma_{t-1}^2 + \gamma \varepsilon_{t-1}^2 S_{t-1}^-$$

	PTE	LIT	FRF	GBP	DM
ω	5.38.10 ⁻⁰⁶ (0.4577)	5.52.10 ⁻⁰⁷ (2.5281)	1.09.10 ⁻⁰⁷ (1.2168)	6.41.10 ⁻⁰⁷ (0.6775)	4.91.10 ⁻⁰⁷ (0.4178)
α	0.150033 (1.6348)	0.0973 (5.1825)	0.1499 (2.0868)	0.1499 (1.9882)	0.1500 (1.6354)
β	0.60000 (3.7285)	0.8874 (34.6565)	0.5999 (3.9848)	0.5999 (2.7845)	0.6000 (3.7245)
γ					
Log.Likelihood	2625.009	2847.117	2740.318	2884.132	2614.785
Schwarz	-10.0472	-10.31512	-10.2670	-10.6530	-10.0493
MSE	1.2.10 ⁻⁰⁵	4.5.10 ⁻⁰⁷	3.6.10 ⁻⁰⁷	3.5.10 ⁻⁰⁸	4.8.10 ⁻⁰⁶
w₁ (p-value)	40.378 (0.000)	31.886 (0.000)	7.253 (0.101)	76.553 (0.000)	84.852 (0.000)
Engle (p-value)	5.394 (0.003)	6.426 (0.035)	10.348 (0.002)	7.037 (0.048)	9.345 (0.002)

Notes:

- (i) Maximun Likelihood Estimation of the parameters. t-statistics in parentheses.
- (ii) Wald-Test: w_1 null hypothesis of $\alpha=\beta=\gamma=0$ on variance.
- (iii) MSE is the estandard error.

TABLAE 5: Gjr(1,1)-M Models Results

$$\Delta \log X_t = Y_t = \varepsilon_t = \kappa \sigma_t^2 + \nu_t$$

	PTE	LIT	FRF	GBP	DM
1992-1993					
κ	8.4464 (1.5255)	8.2757 (1.5076)	3.2675 (0.5364)	-5.7028 (-0.9930)	3.4515 (0.6795)
Log.Likelihood	1627.055	1660.793	1669.338	1633.058	1605.535
Schwarz	-9.3629	-9.3245	-9.6040	-9.4682	-9.2711
Q(10) (p-value)	10.689 (0.382)	3.1614 (0.977)	5.1663 (0.880)	6.7416 (0.798)	7.5002 (0.678)
Q(20) (p-value)	20.232 (0.414)	8.1045 (0.991)	8.9853 (0.983)	11.863 (0.921)	12.484 (0.898)
U Theil	0.69	0.65	0.88	0.71	0.73
1995-1997					
K	5.4548 (0.9564)	-9.6344 (-1.3844)	3.7106 (0.5441)	0.5064 (0.0633)	3.2069 (0.5349)
Log.Likelihood	2656.388	2805.591	2745.973	2885.464	2657.448
Schwarz	-10.0304	-10.2988	-10.2498	-10.6348	-10.0317
Q(10) (p-value)	4.4837 (0.923)	4.9799 (0.893)	6.9702 (0.728)	10.608 (0.389)	6.6239 (0.760)
Q(20) (p-value)	22.532 (0.312)	21.529 (0.367)	17.003 (0.653)	16.610 (0.678)	19.715 (0.5349)
U Theil	0.89	0.89	0.89	0.91	0.98

Notes:

- (i) Maximun Likelihood Estimation of the parameters. t-statistics in parentheses.
- (ii) Q(10) and Q(20) denotes the Ljung-Box statistic for the first k autocorrelations.

Notes