The Exchange Rate Crisis of September 1992 and the Pricing of Italian Financial Futures

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The futures contracts on long term bonds issued by the Italian Treasury and the futures contracts on Eurolira deposits traded in the LIFFE have reacted in a specular way to the exchange rate crisis of September 1992. The pricing of the Italian Government Bonds (BTP) futures becomes more volatile and seems to be affected by a time-varying risk premium. The pricing of Eurolira futures becomes more efficient after the crisis and its volatility declines. These results indicate that the credibility crisis of the Italian Government bonds does not spill over to the off-shore Eurolira assets. © 1998 John Wiley & Sons, Inc. Jrl Fut Mark 18:827–849, 1998

INTRODUCTION

In the early 1990s the London International Financial Futures and Options Exchange introduced financial futures contracts in order to facilitate the trade of Italian bonds and of Eurolira deposits. These contracts are similar to contracts for assets of the main industrialized countries that

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are traded in London. The Italian Government Bond (BTP) futures contract—launched on 19 September 1991—involves a notional bond with a face value of 200 million lira and a 12% interest rate. The Eurolira futures contract, introduced on 12 May 1992, consists of an agreement to buy or sell a one billion lira notional three-month interbank deposit in Eurolira at a predetermined price and delivery date.

The pricing of these contracts has been strongly affected by the financial crisis of September 1992, a crisis that started as a standard European Monetary System (EMS) currency realignment and ended up by threatening the very existence of the EMS.[1]

The present article analyzes the behavior of these contracts taking into account: (i) the regime shift associated with the exchange rate crisis of September 1992; (ii) the differing behavior over time of futures prices of interest and the relevance (if any) of the onshore versus the offshore nature of the underlying asset; (iii) the economic rationale of the time-varying risk premium that seems to characterize the rate of return of some futures contracts from October 1992 onward.

The first section of this article analyzes the statistical properties of the rate of change of daily futures prices. The implications of the martingale hypothesis and the nature of the asymmetries characterizing the volatility of these financial asset prices are reported in the following section. The next two sections assess the relevance of the regime shift in the pricing of these contracts in the context of an asymmetric ARCH (GARCH) modeling framework. The residual non-normality of the conditional distributions is further investigated with the help of a Generalized Error Distribution.

Financial futures have reacted in a specular way to the exchange rate crisis of September 1992. The pricing of BTP futures becomes more volatile and seems to be affected by a time-varying risk premium. The pricing of Eurolira futures becomes more efficient after the crisis and its volatility persistence declines. These results are probably due to a decline in credibility of Italian Treasury bonds that does not spill over to offshore Eu-

[1]The no vote in the Danish referendum on the Maastricht Treaty on 2 June had weakened the scenario of a smooth process of monetary unification. It was generally believed, in the Summer of 1992, that the French referendum of 20 September on Maastricht Treaty ratification would have been followed by a realignment of the EMS exchange rate parities. The crisis started before the French referendum with the lira devaluation of 13 September and soon involved currencies such as sterling for which there was no need for a devaluation on a purchasing power parity basis. A possible interpretation is that speculators thought German interest rates to be too high and in contrast with the domestic economic policy objectives of Germany's European partners. For an analysis of the impact of German reunification on the 1992 EMS crisis see Congdon (1993), Temperton (1993) and Eichengreen and Wyploz (1993), among others.

rolira assets. Indeed, the latter seem to have become more attractive after the 1992 crisis in spite of the exchange rate devaluation.

THE 1992 CRISIS AND THE INTEREST RATE FUTURES TIME SERIES

Historical Background

The Summer of 1992 saw a change for the worse in the markets' perception of the Italian exchange rate policy credibility. This change was due essentially to the persistent failure by the Italian Government (and Parliament) to reduce the yearly budget deficit and to keep the huge public debt under control. A moderate real appreciation of the lira within the EMS had been a basic tenet of the Government's counterinflationary policy, because of the discipline it was supposed to impose on the pricing behavior of firms and on the wage demands of trade unions. A prerequisite of its success, however, was a consistently rigorous budget policy of the Government, which was missing. From January 1987 (the date of the last EMS currencies realignment) to June 1992 Italian consumer prices had risen by 35% whereas the corresponding German prices had risen by 15% only. The real appreciation of the lira was fast becoming a major source of potential EMS instability.

Progressively, from June 1992 onward, the Bank of Italy had to raise the official interest rates in order to withstand the increasing pressure on the exchange rate.[2] The bank had to intervene heavily in the foreign exchange market in June and at the end of August, making full use of the EMS (very) short-term stabilization agreements, which involved unlimited mutual lending between central banks. It could not cope, however, with the unprecedented size of the speculative attacks. At the beginning of September the Bundesbank arrived at the conclusion that the unlimited purchase of lira at the compulsory EMS intervention rate was incompatible with its objectives of domestic monetary policy and therefore withdrew its support. On 13 September a 7% relative devaluation of the lira

[2]Central bank currency became very expensive. On 4 June the rate on central bank repurchase agreements rose from 12% to 14.5%. On 5 July the rate of discount was raised from 12% to 13% and on 16 July to 13.75%. After a .5% decline on 4 August, it rose again and peaked at 15% on 4 September. Liquidity control by the Bank of Italy was very tight and short-term interest rates rose sharply. The three-month treasury bill rate rose from 14.42% in June 1992 to 18.05% in September and the three-month interbank deposit rate from 13.46% in June to 18.22%, on average, in September, well above the corresponding Eurolira rate of 16.98% (Banca d'Italia, 1993, pages 182–183). Obstfeld (1995) attributes this onshore interest premium to market fears of a possible reintroduction of capital controls. This differential almost disappeared after the crisis.

in the EMS was agreed upon.[3] On 15 September both sterling and again lira were heavily attacked by speculation. Central banks were unable to withstand the pressure: on 16 September the Bank of England, after having spent one half of its official reserves, announced its intention to suspend sterling from the European Exchange Rate Mechanism (ERM) and the Italian authorities had no choice but to follow suit.

Successful stabilization of the lira within the ERM bands would have required a further substantial increase in interest rates, which would have destabilized the government bond market. The secondary market for government bonds was already going through a serious crisis. There had been a sharp decline in ownership by households of long-term fixed-rate debt (it would fall from 34% to 21% of the total during the year) and the price of 10-year Treasury fixed rate bonds (BTP) had fallen markedly during the summer and would reach a minimum at the beginning of October. Indeed, the yield differential between Italian and German medium- to long-term government bonds had risen sharply from its 4.5% level in late spring and would peak at 7.5% in October. [For more details, see Uhlig (1997).] The statistical behavior of interest rate futures prices is strongly affected by these events.

Statistical Properties of the Futures Times Series

The sample period extends from 19 September 1991 to 7 April 1994 for the BTP futures price time series and from 12 May 1992 to 7 April 1994 for the Eurolira futures price series. The data are sampled daily and are provided by the LIFFE. Three BTP futures contracts and four Eurolira futures contracts (six from 15 March 1994 onward) are traded every day. Closing settlement prices of the futures contracts nearest delivery are used until two weeks before the delivery date. Then closing prices for the next delivery date are used, in order to avoid expiration effects.

Day-to-day changes, Δf_t, of the logarithm of the prices obtained with this rollover strategy are reported in Figure 1 and Figure 2. A shift in their behavior associated with the exchange rate crisis can be easily identified: BTP futures prices become more volatile after the crisis, whereas Eurolira futures prices become somewhat less volatile. These shifts, however, are not synchronous. Data examination suggests that the lira devaluation of 13 September is immediately reflected in a sharp shift in the price of Eurolira futures contracts as the antispeculative short-term interest rate policy of the Bank of Italy becomes less stringent. It seems to affect the

[3] The lira was devalued by 3.5% and all the other ERM currencies were revalued by 3.5%. For more details on the exchange rate crisis, see Santini (1993).

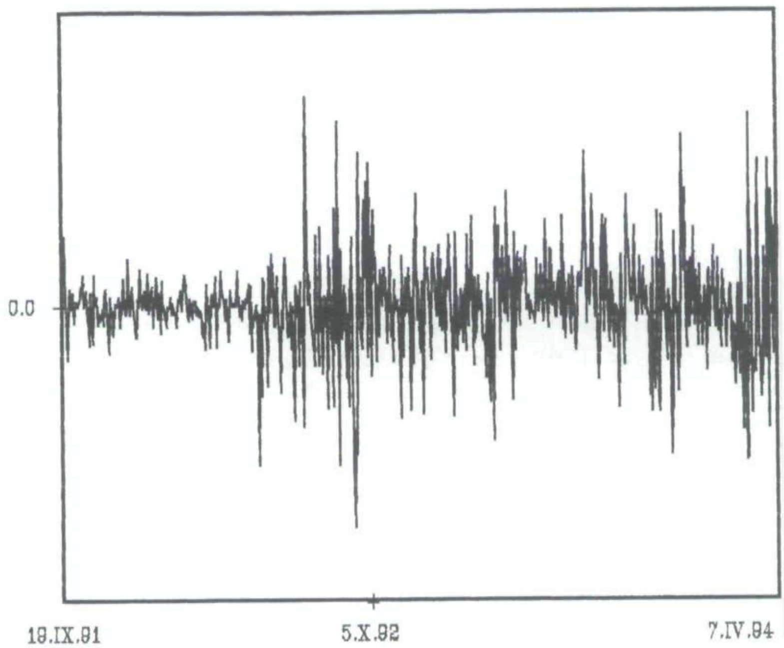

FIGURE 1
BTP futures price rate of change.

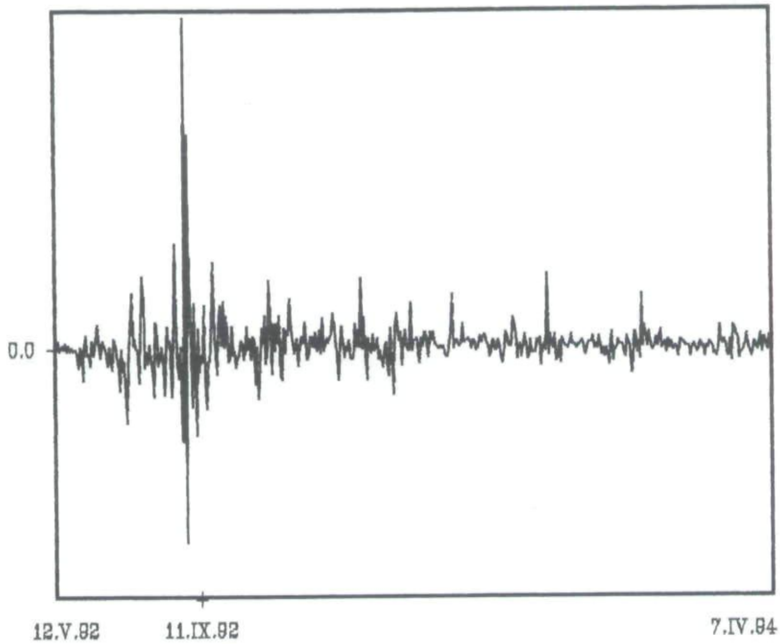

FIGURE 2
Eurolira futures price rate of change.

pricing of BTP futures contracts from 6 October 1992 onward. The behavior of these contracts is probably influenced by the crisis of the secondary Treasury bond market mentioned in the previous section, which lags the exchange rate collapse by three weeks. (It is well known that the prices of Treasury bond futures contracts and of the underlying CTD bonds are linked by cash and carry arbitrage operations.) The dating of the sample breakpoints is corroborated by Wald tests for parameter stability—with unknown breakpoint under the alternative—applied to the OLS estimates of a first order autoregression of the changes in the logarithm of the BTP futures prices and of a second order autoregression of the changes in the logarithm of the Eurolira futures prices.[4]

The principal features of the distributions of these prices in the relevant subperiods (12 May 1992–11 September 1992 and 14 September 1992–7 April 1994 for the Eurolira futures and 19 September 1991–5 October 1992 and 6 October 1992–7 April 1994 for the BTP futures) are set out in Table I.

The characteristics of the two distributions have changed considerably in the two subperiods, justifying the hypothesis of a regime shift. The standard deviations corroborate the hypothesis that BTP futures price (unconditional) volatility has increased after the September crisis and that Eurolira futures price volatility has decreased. At the same time the normality test statistics suggest that non-normality is more severe (because of a high degree of kurtosis) in the first subsample than in the second in the case of BTP futures prices whereas the opposite is true for Eurolira futures prices. The tests for ARCH are significant and support the hypothesis of a time-varying variance.[5] Also the properties of the distribution of $\Delta_\tau f_t$, for $\tau = 5, 10$, have been investigated, following Pesaran and Robinson (1993). As the differencing interval is increased, the distributions become more nearly normal, indicating that they are not stable Paretian.[6]

[4]As will be seen in the ARCH analyses below, these specifications seem to provide the best fit of the respective conditional means. $\mathrm{Sup}_{\pi\in\Pi}\, W_t(\Pi)$ has been estimated—following an approach set out by Andrews (1993)—where W_t is a Wald statistic for known structural break, t is the change point, T is the sample size, $\pi = t/T$ and $\Pi = [\pi_1, \pi_2]$ is some prespecified subset of [0, 1]. If we postulate that the change point may occur in the five-month interval 29 June 1992–27 November 1992, $\Pi =$ [0.305, 0.472] for the BTP time series and [0.070, 0.234] for the Eurolira time series. The estimated $\mathrm{Sup}_{\pi\in\Pi}\, W_t(\Pi)$ statistics are, respectively, 9.049 and 13.650 at the breakpoints mentioned in the text and exceed the corresponding 5% critical values (8.77 and 10.19) obtained from the table set out by Andrews (1993, page 840) for various degrees of freedom and alternative values of the parameter $\lambda = \pi_2(1 - \pi_1)/\pi_1(1 - \pi_2)$.

[5]The ARCH structure that shall be investigated below emerges as the outcome of a process of identification of the nonlinearities of the time series that follows the strategy set out by Hsieh (1989a). It is reported in an extended version of the article.

[6]A characteristic of these distributions is that they are stable under temporal aggregation.

TABLE I

Summary Statistics of the Distribution of $\Delta_\tau f_t$, $\tau = 1, 5, 10$

	Eurolira Futures Prices						*BTP Futures Prices*					
	12.V.92–11.IX.92			*14.IX.92–7.IV.94*			*19.IX.91–5.X.92*			*6.X.92–7.IV.94*		
	$\Delta_1 f_t$	$\Delta_5 f_t$	$\Delta_{10} f_t$	$\Delta_1 f_t$	$\Delta_5 f_t$	$\Delta_{10} f_t$	$\Delta_1 f_t$	$\Delta_5 f_t$	$\Delta_{10} f_t$	$\Delta_1 f_t$	$\Delta_5 f_t$	$\Delta_{10} f_t$
Mean[a]	−0.073	−0.275	−0.560	0.017	0.110	0.215	−0.032	−0.137	−0.238	0.064	0.279	0.513
Max[a]	1.234	1.399	2.248	2.529	5.014	3.787	2.171	3.276	2.587	1.983	4.435	6.150
Min[a]	−1.143	−2.921	−2.793	−2.378	−1.442	−2.503	−2.270	−3.092	−4.231	−1.607	−6.194	−7.352
S.D.[a]	0.340	0.761	0.949	0.268	0.577	0.707	0.423	0.814	1.120	0.570	1.427	2.177
Sk.	0.043	−0.302	0.273	0.427	2.585	0.609	−0.525	−0.293	−0.966	0.010	−0.521	−0.318
Kt.	6.774	4.135	3.615	38.717	21.315	6.366	11.581	5.990	4.161	3.786	5.268	3.472
J.B.	50.482[b]	5.650	2.166	21008.1[b]	5960.9[b]	210.9[b]	822.0[b]	100.9[b]	54.16[b]	9.79[b]	98.65[b]	9.93[b]
n	85	82	77	396	396	396	264	261	256	380	380	380
ARCH(1)	6.074[b]			257.033[b]			15.654[b]			29.418[b]		
ARCH(5)	19.328[b]			361.152[b]			30.814[b]			50.759[b]		

S.D. = standard deviation. Sk. = skewness. Kt. = kurtosis. J.B. = Jarque-Bera normality test statistic. n = number of observations. ARCH(k) = LM test statistic for k^{th} order ARCH.

[a]the entries of this row are multiplied by 100.

[b]significant at the 5% level.

TABLE II

Sign and Size Bias Tests

	Eurolira Futures Price Changes		*BTP Futures Price Changes*	
Tests	*12.V.92–11.IX.92*	*14.IX.92–7.IV.94*	*19.IX.91–5.X.92*	*6.X.92–7.IV.94*
Sign Bias	−0.1647	−0.0282	1.1606	1.4429
Negative Size Bias	−2.0279[a]	−11.3246[a]	−4.1661[a]	−4.0856[a]
Positive Size Bias	0.8714	11.4229[a]	1.8346	0.8679
Joint	8.0632[a]	225.7893[a]	26.9779[a]	24.7100[a]

Notes: The sign, negative size and positive size bias tests involve (t-statistics) tests of the null hypothesis that $b = 0$, $b' = 0$ and $b'' = 0$ in the following OLS regressions: (1) $v_t^2 = a + bS_{t-1}^- + e_t$; (2) $v_t^2 = a' + b'S_{t-1}^-\varepsilon_{t-1} + e_t$; (3) $v_t^2 = a'' + b''S_{t-1}^+\varepsilon_{t-1} + e_t$ $\varepsilon_t = y_t - \mu$, $v_t = \varepsilon_t \sigma$, where y_t is the logarithm of the first difference of the futures prices and μ and σ are the corresponding unconditional mean and standard deviation. S_{t-1}^- is a dummy variable that takes the value of one when ε_{t-1} is negative and zero otherwise. $S_{t-1}^+ = 1 - S_{t-1}^-$. The joint bias test is a Wald test of the null hypothesis that $b''' = c''' = d''' = 0$ in the regression $v_t^2 = a''' + b'''S_{t-1}^- + c'''S_{t-1}^-\varepsilon_{t-1} + d'''S_{t-1}^+\varepsilon_{t-1} + e_t$.
[a]significant at the 5% level.

It is usually believed that the volatility of stock returns is asymmetrically affected by shocks, negative innovations having a larger impact than positive ones. [Black (1976) and Campbell and Hentschel (1992) provide two alternative theoretical interpretations of this phenomenon.] The relevance of this hypothesis for the futures price changes of interest can be assessed with the help of the sign bias, negative size bias, positive size bias and joint bias tests set out by Engle and Ng (1993). Results are reported in Table II.

BTP futures price volatility seems to be influenced by large negative innovations in both subperiods and (possibly) by large positive innovations in the first. Eurolira futures price volatility is strongly affected by large positive and negative innovations in the second subperiod and by large negative innovations only in the first. These findings suggest that symmetric ARCH (GARCH) approaches may be inadequate to model the impact of news on futures price volatility.

The well-known test strategy set out by Perron (1988) is used to assess the stochastic structure of the futures prices time series under investigation. The hypothesis of a unit root does not seem to be rejected.[7]

[7]The DF and ADF statistics suggest that the logarithms of futures prices are I(1) in the time periods of interest and that the corresponding first differences are I(0). The latter follow a stationary process but are not iid, because of the presence of first and second moment serial dependence.

Unit Root Tests

Eurolira Futures Prices				*BTP Futures Prices*			
12.V.92–11.IX.92		*14.IX.92–7.IV.94*		*19.IX.91–5.X.92*		*6.X.92–7.IV.94*	
f_t	Δf_t	f_t	Δf_t	f_t	Δf_t	f_t	Δf_t

Subsequently, only the first differences of the relevant time series will be considered in the empirical analysis.

ARCH MODELING OF FUTURES PRICES CHANGES

The Martingale Hypothesis and the Conditional Mean

Samuelson (1965) and Fama (1965), among others, have pointed out that the martingale behavior of an asset price is consistent with the hypothesis of (weak) efficiency of the corresponding market.

Let f_t be the natural logarithm of the price of a futures contract and let Z_{t-1} be a vector of N elements that enter the information set at time $t - 1$. The behavior of the conditional mean of the futures price rate of change can be modeled as

$$\Delta f_t = a_0 + \sum_{k=1}^{N} a_k Z_{k,t-1} + u_t \tag{1}$$

where the N dimensional vector Z_{t-1} may include dummy variables, such as day of the week dummies, and lagged rates of change of the futures prices. The martingale hypothesis requires that Δf_t be a fair game, i.e., that $a_k = 0$, $k = 0, 1, \ldots, N$ and that u_t be serially uncorrelated with mean zero. It does not require, however, that the error term be homoskedastic. Any systematic link between an element of the information set and Δf_t would provide evidence against the joint hypothesis of the martingale property of futures prices and of market efficiency. Under the maintained hypothesis of market efficiency, rejection of the martingale behavior could well be due to the presence of a risk premium.[8]

Φ_3	2.3981	29.2335[a]	4.4022	136.1310[a]	1.5193	151.7430[a]	1.4781	150.8870[a]
Φ_2	2.2956	19.4890[a]	3.4783	90.8861[a]	1.5813	101.1850[a]	2.2303	100.6170[a]
Φ_1	1.6708	29.4365[a]	2.0486	135.9680[a]	1.1058	149.7910[a]	3.0388	150.2130[a]
τ_τ	−2.1431	−7.6464[a]	−2.9094	−16.4628[a]	−1.6048	−17.5366[a]	−1.3101	−17.3698[a]
τ_μ	−1.1488	−7.6729[a]	−1.5794	−16.4784[a]	−0.1265	−17.4174[a]	−1.6280	−17.3305[a]
n	2	1	0	1	0	0	1	0

Notes: n = order of the OLS autoregression used to perform the ADF tests. The Φ_3, Φ_2, Φ_1, τ_τ and τ_μ distributions have been tabulated by Dickey and Fuller (1981) and Fuller (1976).
[a]significant at the 5% level.

[8]Hodrick and Srivastava (1987) have analyzed the relevance of a time-varying risk premium in the pricing of futures contracts and have derived the following relationship, based on the capital asset pricing approach of Lucas (1982)

$$F_{t-1,k} = E(F_{t,k-1}|\Psi_{t-1}) + C_{t-1}(F_{t,k-1}; Q_{t,1}R_{t-1,1})$$

Domowitz and Hakkio (1985), Morgan and Morgan (1987) and McCurdy and Morgan (1988), among others, suggest that the relevance of a time-varying risk premium be assessed with the help of the ARCH-M and GARCH-M models of Engle, Lilien and Robins (1987), and Bollerslev, Engle, and Wooldridge (1988). Equation (1) can then be rewritten as follows

$$\Delta f_t = a_0 + \sum_{k=1}^{N} a_k Z_{k,t-1} + d\sqrt{h_t} + u_t \tag{2}$$

where h_t is the conditional variance. The martingale hypothesis requires that u_t be serially uncorrelated, with mean zero, and that $a_k = 0$, $k = 0, 1, \ldots, N$ and $d = 0$. A coefficient d in eq. (2) that is significantly different from zero suggests the existence of a time-varying risk premium.

Relevance of Asymmetry in the Conditional Variance

A major criticism for the standard ARCH (GARCH) parameterization of conditional variance is related to the assumption that past unanticipated changes in the conditional mean have the same impact on volatility regardless of their sign.

Two specifications of the asymmetric impact of news on volatility are used in this paper. They are due, respectively, to Engle (1990) and Sentana (1995) and to Glosten, Jagannathan, and Runkle (1993). As pointed out by Engle and Ng (1993) and more recently by Hentschel (1995), they affect in a different way the "news impact curve," which relates revisions in conditional variance to unanticipated changes in the conditional mean. Engle and Sentana shift the standard symmetric GARCH impact curve to the right, whereas Glosten, Jagannathan, and Runkle (GJR) rotate it clockwise, obtaining a steeper slope on the negative side than on the positive one. These two asymmetries are intrinsically different. Engle and Sentana set up a threshold (on the news) which reduces the impact of positive shocks and raises the impact of equally large but negative ones. Positive price shocks at first reduce volatility, since they are associated

where $F_{t-1,k}$ is the time $t-1$ price of a futures contract with maturity date $t-1+k$, $E(F_{t,k-1}|\Psi_{t-1})$ is the one period ahead expected value of the futures price conditional on the information set at time $t-1$, $Q_{t,1}$ is the intertemporal marginal rate of substitution of money between $t-1$ and t, $R_{t-1,1}$ is one plus the riskless rate of interest between $t-1$ and t and $C_{t-1}(.)$ is the conditional covariance operator. The martingale hypothesis is contradicted if the covariance term is nonzero. The latter can be seen as a time-varying risk premium for a one period change in futures prices in a contract maturing at date $t-1+k$.

with an increase in the attractiveness of the asset. A behavior of this kind may characterize the pricing of assets that are affected by shifts in credibility, such as long-term government bonds, and the pricing of the corresponding futures contracts. Glosten, Jagannathan, and Runkle too assume that positive shocks have a smaller effect on volatility than negative ones, a small absolute value, however, being associated with a small impact on volatility. (The news impact curve is centered at $u_{t-1} = 0$.) Gagnon and Lypny (1995) use this approach to model the volatility of the price of three-month Canadian bankers acceptances and of the corresponding BAX futures contracts traded at the Montreal Exchange.[9]

DAILY PRICING OF BTP FUTURES

The properties of the conditional first and second moments of the distribution of BTP futures price changes are investigated at first over the whole sample period, using dummies to capture the structural shift associated with the exchange rate crisis, and then, in more detail, in each of the two relevant subsamples. An asymmetric volatility model is called for, in line with the outcome of the Engle & Ng asymmetry tests reported in Table II above. Some preliminary estimates suggest, moreover, that the quadratic GARCH conditional variance parameterization systematically outperforms the GJR GARCH parameterization. The conditional variance has thus been modeled as a QGARCH (1,1) process in the three time periods of interest.[10] For computational reasons futures prices rates of change are multiplied by 100.[11]

[9]They provide an explanation of asymmetry in terms of the quoting conventions of short term interest rate futures prices (100-the implied yearly interest rate) and of the observed behavior of interest rate volatility, which is proportionally lower when rates are low and higher when rates are high. It may well apply to Eurolira futures prices.

[10]The standard QGARCH(1,1) conditional variance can be formulated as

$$h_t = \alpha_0 + \alpha_1(u_{t-1} - b)^2 + \beta h_{t-1}$$

where b is a shift coefficient used to model asymmetry (Engle (1990), Campbell and Hentschel (1992)). I have followed Sentana (1995) and have estimated the QGARCH(1,1) conditional variance

$$h_t = \theta + \alpha u_{t-1}^2 + \beta h_{t-1} + \psi u_{t-1}$$

in the following restricted version, which ensures nonnegativity of h_t,

$$h_t = c_0^2 + c_1^2(u_{t-1} - b)^2 + c_2^2 h_{t-1}$$

where $\theta = c_0^2 + c_1^2 b^2$, $\psi = -2c_1^2 b$, $\beta = c_2^2$ and $\alpha = c_1^2$. The standard errors are computed using the delta method. The corresponding robust standard errors are obtained replacing—in the delta method estimation procedure—the variance matrix of the coefficients with the robust asymptotic variance matrix set out by Bollerslev and Wooldridge (1992, Theorem 2.1, page 148).

[11]The corresponding GARCH(1,1)-GJR model has been estimated in the three relevant time periods and the relative goodness of fit of the two asymmetry specifications has been assessed with the help

TABLE III

QGARCH(1,1) Model Estimates

19.IX.1991–7.IV.1994

$$\Delta f_t = a_0 + a_1 \Delta f_{t-1} + a_3 Md_t + a_4 Dum_t + u_t \tag{3}$$

$$h_t = \theta + \alpha u_{t-1}^2 + \beta h_{t-1} + \psi u_{t-1} + \delta Md_t + \gamma Dum_t \tag{4}$$

a_0	a_1	a_3	a_4	LLF
0.003	0.133	−0.051	0.032	−360.114
(0.131)	(2.527)	(−1.328)	(0.918)	
[0.189]	[3.086]	[−1.461]	[1.148]	

θ	α	β	ψ	δ	γ
0.017	0.229	0.631	−0.125	0.014	0.045
(2.628)	(10.800)	(10.386)	(−5.009)	(1.012)	(3.294)
[2.819]	[3.686]	[8.865]	[−4.084]	[0.891]	[2.009]

Sk.	*Kt.*	*J.B.*	*ARCH(4)*	*L.B.(4)*	*S.B.*	*N.S.B.*	*P.S.B.*	*Jt.*
−0.272	5.148	129.685[a]	0.132	3.369	−0.250	0.451	0.004	0.232

Notes: Sk. = skewness. Kt. = kurtosis. J.B. = Jarque-Bera normality test statistic. ARCH(k) = LM test statistic for k^{th} order ARCH. L.B.(k) = Ljung Box Q-statistic for k^{th} order serial correlation. S.B. = sign bias test. N.S.B. = negative size bias test. P.S.B. = positive size bias test. Jt. = joint bias test. *t*-statistics are in round brackets. Robust *t*-statistics are in square brackets.
[a]significant at the 5% level.

The following relationships have been estimated over the time period 19 September 1991–7 April 1994

$$\Delta f_t = a_0 + a_1 \Delta f_{t-1} + a_3 Md_t + a_4 Dum_t + u_t \tag{3}$$

$$h_t = \theta + \alpha u_{t-1}^2 + \beta h_{t-1} + \psi u_{t-1} + \delta Md_t + \gamma Dum_t \tag{4}$$

Md_t is a dummy variable that takes the value of one if t is a Monday and zero otherwise. Dum_t is a dummy variable that takes the value of zero for the period 19 September 1991–5 October 1992 and the value of one for the period 6 October 1992–7 April 1994, which follows the exchange rate crisis.[12]

of the Akaike Information Criterion (AIC).

	19.IX.91–7.IV.94	*19.IX.91–5.X.92*	*6.X.92–7.IV.94*
QGARCH(1,1)	740.228	121.268	587.443
GARCH(1,1)-GJR	742.379	142.462	596.308

Lower values are associated with the QGARCH(1,1) model.

[12]I am following here the suggestion of a referee. For an analogous use of regime shift dummies see, among others, Koutmos, Negakis, and Theodossiou (1993).

The severe kurtosis and the subsequent non-normality of the standardized residuals evidenced by the Jarque-Bera test statistic suggest that the conventional standard errors are biased. Accordingly, in Table III are reported also the t-ratios based on the robust Quasi Maximum Likelihood Estimation (QMLE) inference procedure of Bollerslev and Wooldridge (1992). The martingale hypothesis is rejected, as lagged futures price changes systematically affect current changes. There is evidence of a structural change in the variance and, to a lesser extent, in the mean of the distribution. (Note, however, that robust t-statistic tests tend to be conservative.) The regime shift dummies have positive coefficients and corroborate the preliminary findings of Table I of an increase—in the second subperiod—in the average futures price rate of change and in its volatility.

The estimates over the two subsamples provide additional information on the effects of the exchange rate crisis. In the year preceding the crisis the authorities were still able to issue large quantities of long-term fixed-rate bonds (BTP) with no significant strain in the financial markets. The yield differential between Italian and German medium to long-term government bonds was stable and the Milan secondary market for Treasury assets was still growing steadily.

Futures price behavior reflects this state of affairs. It can be modeled by the following relationships

$$\Delta f_t = a_3 Md_t + u_t + a_5 u_{t-1} \tag{3'}$$

$$h_t = \theta + \alpha u_{t-1}^2 + \beta h_{t-1} + \psi u_{t-1} \tag{4'}$$

The estimates are set out in the first column of Table IV. The market seems to process information in an efficient way, even if the martingale hypothesis is rejected. The change in the logarithm of BTP futures prices is affected by a Monday effect only and the residuals of the conditional mean relationship follow, as is often the case with financial data, a MA(1) process.[13] Here too I have added robust t-ratios since the standardized residuals have a non-normal distribution. Table V presents robust Lagrange multiplier tests based on the outer product of the gradient (OPG-LM) for variables omitted from equations (3′) and (4′) due to Wooldridge

[13]Lo and MacKinlay (1990), Baillie and Bollerslev (1990) and Susmel and Engle (1994), among others, attribute a (spurious) negative serial correlation in the returns to the averaging of the sample prices used to build the relevant time series and to the non-synchroneity of these prices. In the present paper, however, the serial correlation is positive and may require an alternative explanation, such as the presence of a time-varying risk premium which is too elusive to come out in the conditional mean specification. In this case the martingale hypothesis would be violated.

TABLE IV

QGARCH(1,1) Model Estimates over the Two Subsamples

19.IX.91–5.X.92

$$\Delta f_t = a_3 Md_t + u_t + a_5 u_{t-1} \tag{3'}$$

$$h_t = \theta + \alpha u_{t-1}^2 + \beta h_{t-1} + \psi u_{t-1} \tag{4'}$$

6.X.92–7.IV.94

$$\Delta f_t = a_1 \Delta f_{t-1} + a_2 \sqrt{h_t} + u_t \tag{3''}$$

$$h_t = \theta + \alpha u_{t-1}^2 + \beta h_{t-1} + \psi u_{t-1} + \delta Md_t \tag{4''}$$

	N	GED	N
υ	—	1.058	—
	—	(2.587)	—
a_1	—	—	0.109
	—	—	(2.170)
a_2	—	—	0.118
	—	—	(2.336)
a_3	−0.087	−0.069	—
	(−2.468)	(−6.340)	—
	[−2.080]		
a_5	0.160	0.067	—
	(2.023)	(2.462)	—
	[2.278]		
θ	0.007	0.006	0.020
	(1.128)	(1.457)	(2.429)
	[1.244]		
α	0.205	0.127	0.088
	(6.437)	(3.352)	(3.240)
	[1.883]		
β	0.747	0.795	0.838
	(10.236)	(2.406)	(26.730)
	[3.601]		
ψ	−0.075	−0.054	−0.075
	(−2.138)	(−2.339)	(−3.964)
	[−1.917]		
δ	—	—	0.101
	—	—	(3.742)
LLF	−54.634	−36.789	−286.721
Sk.	−0.474	−0.561	0.049
Kt.	6.652	7.293	3.439
J.B.	150.620[a]	206.015[a]	3.044
ARCH(4)	1.154	1.094	1.532
L.B.(4)	1.620	1.953	3.705
S.B.	0.884	0.966	0.208
N.S.B.	−0.522	−0.895	0.204
P.S.B.	−0.099	−0.085	0.387
Jt.	0.920	1.314	0.319

Notes: N = conditional normal distribution. GED = conditional generalized error distribution. *t*-statistics are in round brackets. Robust *t*-statistics are in square brackets.
[a]significant at the 5% level.

TABLE V

OPG Lagrange Multiplier Test Statistics for Omitted Variables

19.IX.91–5.X.92					
Equations (3′) and (4′)					
In Mean	Intercept	Δf_{t-1}	Δf_{t-2}	$h_t^{1/2}$	u_{t-2}
	0.5453	0.0646	0.0088	0.0011	0.1662
	[0.46]	[0.80]	[0.92]	[0.97]	[0.68]
In Variance	u_{t-2}^2	u_{t-2}	h_{t-2}	Md_t	
	0.0400	0.2328	0.0849	0.5214	
	[0.84]	[0.63]	[0.77]	[0.47]	
6.X.92–7.IV.94					
Equations (3″) and (4″)					
In Mean	Intercept	Δf_{t-2}	Δf_{t-3}	Md_t	u_{t-1}
	0.4987	1.1991	0.6339	0.4823	1.5415
	[0.48]	[0.27]	[0.43]	[0.49]	[0.21]
In Variance	u_{t-2}^2	u_{t-2}	h_{t-2}		
	1.2448	2.3344	0.3802		
	[0.26]	[0.13]	[0.54]		

Notes: The test statistics are asymptotically distributed as a chi-square with 1 degree of freedom. The statistics in the first subperiod are robust to non-normality. Probability values are in square brackets.

(1987, 1990). The selected model specification does not seem to be contradicted by the data, as the null is never rejected.

The use of a conditional normal distribution does not account for the degree of fat-tailedness of the unconditional distribution of price changes and alternative leptokurtic distributions are investigated, following Bollerslev (1987), Hsieh (1989b), Baillie and Bollerslev (1989), Baillie and DeGennaro (1990), and Pesaran and Robinson (1993), among others. Some preliminary experiments suggest that the Generalized Error Distribution (GED) provides the best fit.[14] The estimates of equations (3′) and (4′), associated with a GED specification for the conditional distribution of the residuals, are reported in the second column of Table IV. The values of the coefficients are similar to the previous ones and, on the whole, the quality of fit has improved. The estimated value of the tail

[14]The specification of the GED, normalized to have zero mean and unit variance, is the following:

$$g(v_t) = \frac{\upsilon \exp[-(1/2)|v_t/\lambda|^{\upsilon}]}{\lambda 2^{[1+1/\upsilon]}\Gamma(1/\upsilon)}$$

where $v_t\sqrt{h_t} = u_t$, $0 < \upsilon \leq \infty$, $\Gamma(.)$ is the gamma function and

$$\lambda = [2^{(-2/\upsilon)}\Gamma(1/\upsilon)/\Gamma(3/\upsilon)]^{1/2}$$

υ is a parameter governing the thickness of the tails. When $\upsilon < 2$ ($\upsilon > 2$) the distribution of v_t has thicker (thinner) tails than the normal and when $\upsilon = 2$ v_t has the standard normal distribution.

thickness parameter υ is 1.058 and is significantly different from zero at the 5% level. (The likelihood ratio test statistic for the hypothesis that $\upsilon = 2$ is 35.69 and strongly rejects the null that the conditional distribution is normal.) The standardized residuals are serially uncorrelated and show no evidence of residual ARCH. They are affected, however, by a high degree of kurtosis. With both conditional distributions the sign and size bias tests of Engle and Ng are not significant. The QGARCH (1,1) conditional variance specification seems to adequately model the impact of news on volatility.

In the second subperiod, from 6 October 1992 to 7 April 1994, market efficiency declines. The martingale hypothesis is clearly rejected as the conditional mean is affected by the lagged futures price rate of change and by the square root of the conditional variance. BTP futures price behavior can be modeled as follows

$$\Delta f_t = a_1 \Delta f_{t-1} + a_2 \sqrt{h_t} + u_t \tag{3''}$$

$$h_t = \theta + \alpha u_{t-1}^2 + \beta h_{t-1} + \psi u_{t-1} + \delta Md_t \tag{4''}$$

The estimates are reported in the third column of Table IV. Diagnostic statistics indicate no autocorrelation in the standardized residuals or squared residuals. The use of the normal conditional distribution, combined with the QGARCH (1,1) conditional variance, has been successful in explaining the original leptokurtosis in the unconditional distribution of futures price changes. Kurtosis drops to 3.439 and the Jarque-Bera statistic is not significant at the 5% level. OPG-LM tests for variables omitted from equations (3″) and (4″) are reported in Table V. None of the variables examined is significant and the selected model specification does not seem to be contradicted by the data. Here too the sign and size bias test statistics are no longer significant.

The exchange rate crisis of September 1992 and the subsequent turmoil in the secondary market for Treasury bonds seriously affect the pricing of BTP futures contracts as the counterinflationary credibility of the monetary authorities declines and doubts arise on the very solvency of the Italian Treasury, compounded by the perception of an endemic political instability. The own conditional variance, seen as a proxy for a time-varying risk premium, has a strong positive effect on the conditional mean. (In the literature, as pointed out by Campbell, Lo, and MacKinlay (1997, pages 496–498), such an effect has usually been found to be either nil or very small.) The shift in the conditional mean brings about the increase in volatility that has been detected in the statistical analysis of

TABLE VI

ARCH(3)-GJR Model Estimates

12.V.1992–7.IV.1994

$$\Delta f_t = a_0 + a_1 \Delta f_{t-1} + a_2 \Delta f_{t-2} + a_3 Md_t + a_4 Dum_t + u_t \quad (5)$$

$$h_t = \alpha_0 + \alpha_1 u_{t-1}^2 + \alpha_2 u_{t-2}^2 + \alpha_3 u_{t-3}^2 + \eta S_{t-1}^- u_{t-1}^2 + \gamma Dum_t \quad (6)$$

a_0	a_1	a_2	a_3	a_4			LLF	
−0.050	0.146	−0.071	−0.068	0.065			91.911	
(−2.055)	(3.451)	(−2.274)	(−4.714)	(2.678)				
[−1.481]	[2.425]	[−0.750]	[−2.169]	[1.879]				
α_0	α_1	α_2	α_3	η	γ			
0.033	0.124	0.658	0.040	0.348	−0.019			
(6.247)	(7.309)	(8.080)	(1.173)	(2.334)	(−2.607)			
[2.096]	[1.880]	[3.270]	[1.123]	[1.882]	[−1.214]			
Sk.	*Kt.*	*J.B.*	*ARCH(4)*	*L.B.(4)*	*S.B.*	*N.S.B.*	*P.S.B.*	*Jt.*
0.874	8.147	576.211[a]	0.810	2.550	0.844	0.527	0.271	1.108

[a]significant at the 5% level.

Table I. Volatility persistence does not change significantly over the two time periods of interest.

DAILY PRICING OF EUROLIRA FUTURES

The Engle and Ng asymmetry tests of Table II are compatible with an asymmetric parameterization of Eurolira futures price volatility. Preliminary investigation suggests that the GJR asymmetry specification provides the best fit.[15]

The following ARCH(3)-GJR model is thus estimated, at first over the whole sample period and then over the two relevant subperiods. Table VI presents preliminary quasi maximum likelihood estimates of the joint parameterization of the conditional mean and of the conditional variance of Eurolira futures price changes—eqs. (5) and (6)—from 12 May 1992 to 7 April 1994.

[15]The QARCH(3) conditional variance specification is outperformed by the ARCH(3)-GJR specification. Using the Akaike Information Criterion the following results are obtained

	12.V.92–7.IV.94	*12.V.92–9.IX.92*	*14.IX.92–7.IV.94*
QARCH(3)	−153.130	12.991	−186.024
ARCH(3)-GJR	−161.822	−2.294	−196.368

The AIC statistic is minimized by the ARCH(3)-GJR model.

$$\Delta f_t = a_0 + a_1 \Delta f_{t-1} + a_2 \Delta f_{t-2} + a_3 Md_t + a_4 Dum_t + u_t \quad (5)$$

$$h_t = \alpha_0 + \alpha_1 u_{t-1}^2 + \alpha_2 u_{t-2}^2 + \alpha_3 u_{t-3}^2 + \eta S_{t-1}^- u_{t-1}^2 + \gamma Dum_t \quad (6)$$

The dummy variable Dum_t takes the value of zero over the period 12 May 1992–11 September 1992 and the value of one over the period 14 September 1992—7 April 1994, Md_t, is a Monday effect dummy, which seems to affect the conditional mean only, and $S_t^- = 1$ if $u_t < 0$, $S_t^- = 0$ otherwise. The Jarque-Bera test statistic is highly significant and suggests that the standardized residuals exhibit non-normality because of leptokurtosis. I have also reported, therefore, t-ratios based on the robust inference procedure of Bollerslev and Wooldridge. The regime shift dummies have opposite signs and corroborate the findings of Table I of an increase in the mean value of the first difference of the logarithm of Eurolira futures prices after the crisis and of a decline in its volatility.

Here too the estimates over the two subsamples are highly informative. They suggest that the market becomes more efficient after the crisis.

In the period 12 May–9 September 1992 the martingale hypothesis is clearly rejected. (Two outliers are dropped following the introduction on 9 September 1992 of a 12.5% tax on Eurolira assets interest rates.) The pricing of Eurolira contracts is affected by the growing weakness of the currency and by the attempts by the Bank of Italy to ease the speculative pressure mentioned above. Equations (5′) and (6′) seem to provide an adequate description of Eurolira futures price behavior

$$\Delta f_t = a_0 + a_1 \Delta f_{t-1} + a_2 \Delta f_{t-2} + a_3 Md_t + u_t \quad (5')$$

$$h_t = \alpha_0 + \alpha_1 u_{t-1}^2 + \alpha_2 u_{t-2}^2 + \alpha_3 u_{t-3}^2 + \eta S_{t-1}^- u_{t-1}^2 + \delta Md_t \quad (6')$$

The estimates are reported in the first column of Table VII. The standardized residuals are serially uncorrelated and show no evidence of residual ARCH. The Jarque-Bera test statistic is not significant: the ARCH(3)-GJR parameterization, combined with the normal conditional distribution, has accounted for the leptokurtosis in the unconditional distribution of futures price changes. Here too, the OPG-LM test statistics set out in Table VIII suggest that the selected model specification is not contradicted by the data.

The lira floating outside the ERM, which follows the 1992 crisis, is associated with a reduction of tensions in the Italian short term financial markets. Eurolira futures price behavior reflects this new state of affairs as market efficiency improves and volatility declines. In the period 14

TABLE VII

ARCH(3)-GJR Model Estimates over the Two Subsamples

12.V.92–9.IX.92

$$\Delta f_t = a_0 + a_1\Delta f_{t-1} + a_2\Delta f_{t-2} + a_3Md_t + u_t \quad (5')$$

$$h_t = \alpha_0 + \alpha_1 u_{t-1}^2 + \alpha_2 u_{t-2}^2 + \alpha_3 u_{t-3}^2 + \eta S_{t-1}^- u_{t-1}^2 + \delta Md_t \quad (6')$$

14.IX.92–7.IV.94

$$\Delta f_t = a_0 + a_3Md_t + u_t + a_5u_{t-1} \quad (5'')$$

$$h_t = \alpha_0 + \alpha_1 u_{t-1}^2 + \alpha_2 u_{t-2}^2 + \alpha_3 u_{t-3}^2 + \eta S_{t-1}^- u_{t-1}^2 \quad (6'')$$

	N	*N*	*GED*
υ	—	—	0.901
	—	—	(12.592)
a_0	−0.031	0.014	0.007
	(−1.392)	(1.724)	(1.289)
		[1.284]	
a_1	0.290	—	—
	(2.440)	—	—
a_2	−0.306	—	—
	(−2.947)	—	—
a_3	−0.097	−0.057	−0.014
	(−2.151)	(−3.794)	(−1.243)
		[−1.846]	
a_5	—	0.110	0.079
	—	(1.824)	(1.973)
		[1.836]	
α_0	0.002	0.013	0.012
	(0.209)	(3.844)	(4.820)
		[4.102]	
α_1	0.001	0.171	0.177
	(0.020)	(3.442)	(5.235)
		[1.455]	
α_2	0.651	0.535	0.464
	(3.199)	(5.815)	(2.776)
		[3.161]	
α_3	0.191	0.036	0.085
	(2.009)	(0.971)	(0.937)
		[0.933]	
η	0.379	0.501	0.587
	(1.584)	(2.935)	(1.824)
		[1.816]	
δ	0.070	—	—
	(1.352)	—	—
LLF	11.147	106.184	154.099
Sk.	0.056	1.197	1.408
Kt.	2.772	8.957	10.032
J.B.	0.204	675.034[a]	941.866[a]
ARCH(4)	5.508	2.713	2.945
L.B.(4)	0.700	2.420	2.220
S.B.	−0.233	0.859	0.770
N.S.B.	0.312	0.387	0.651
P.S.B.	−0.250	−0.362	−0.343
Jt.	0.290	1.734	2.151

TABLE VIII

OPG Lagrange Multiplier Test Statistics for Omitted Variables

12.V.92–9.IX.92				
Equations (5′) and (6′)				
In Mean	Δf_{t-3}	$h_t^{1/2}$	u_{t-1}	
	1.8332	2.6241	2.9769	
	[0.18]	[0.11]	[0.08]	
In Variance	u_{t-4}^2	h_{t-1}		
	1.8165	2.0460		
	[0.18]	[0.15]		
14.IX.92–7.IV.94				
Equations (5″) and (6″)				
In Mean	Δf_{t-1}	Δf_{t-2}	$h_t^{1/2}$	u_{t-2}
	0.1459	1.0885	0.5021	0.4764
	[0.70]	[0.30]	[0.48]	[0.49]
In Variance	u_{t-4}^2	h_{t-1}	Md_t	
	2.9031	0.0461	0.9659	
	[0.09]	[0.83]	[0.33]	

Notes: The statistics in the second subperiod are robust to non-normality. Probability values are in square brackets.

September 1992–7 April 1994 the relevant price behavior can be modeled as follows

$$\Delta f_t = a_0 + a_3 Md_t + u_t + a_5 u_{t-1} \tag{5''}$$

$$h_t = \alpha_0 + \alpha_1 u_{t-1}^2 + \alpha_2 u_{t-2}^2 + \alpha_3 u_{t-3}^2 + \eta S_{t-1}^- u_{t-1}^2 \tag{6''}$$

The estimates are reported in the second column of Table VII. Futures price changes are affected by a constant and by a Monday effect only, and the residuals follow a MA(1) process. The persistence of the volatility shocks—measured by the sum of the coefficients of the lagged square residuals in equation (6″)—declines markedly. The robust OPG-LM tests reported in Table VIII are not significant and suggest that the selected model specification is appropriate.

An investigation of alternative leptokurtic distributions suggests that here too the Generalized Error Distribution provides the best fit. The estimates of equations (5″) and (6″), associated with a GED specification of the conditional distribution of the residuals, are reported in the third column of Table VII. The constant term and the coefficient of the Monday effect are not significantly different from zero. The market seems thus to operate in an efficient way, even if the martingale hypothesis has probably to be rejected because of the MA(1) structure of the conditional mean residuals. The likelihood ratio test statistic for the hypothesis that $\upsilon = 2$ is 95.83 and strongly rejects the null of a normal conditional distribution. Indeed, the estimated value of the tail thickness parameter is 0.901 and reflects the severe kurtosis of the conditional distribution. In both time periods the Engle and Ng sign and size tests are not significant.

CONCLUSION

The stochastic properties of the BTP and Eurolira daily futures price changes are explored. It is found that they are conditionally heteroskedastic.[16] The asymmetric impact of news on volatility is captured with the help of two different parameterizations and may be related to the properties of the underlying assets.

The September 1992 exchange rate crisis has strongly affected the pricing of these contracts. The behavior of BTP futures prices suggests that the exchange rate crisis has triggered a credibility crisis of Italian government bonds. A time varying risk premium affects the conditional mean of the daily BTP futures price rate of change from October 1992 onward. The analysis of Eurolira futures prices suggests that the exchange rate crisis has actually brought about an improvement in market efficiency and a decline in the persistence of the volatility of conditional variance. These results are conducive to the hypothesis that the exchange rate floating per se does not play a relevant role in determining day-to-day financial futures price behavior. The Italian risk premium seems to have financial origin, it affects the Italian government bond market and not the Eurolira assets market, a finding reflected in the pricing of the corresponding futures contracts.

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[16]The conditional variance parameterizations selected in this paper seem to be immune from the critique of Lamoureux and Lastrapes (1990). (Robust) OPG-LM tests for omitted variables have been performed, the null being that trade volume—i.e. the daily number of contracts—does not enter the conditional variance equation, null that is never rejected. Also the relevant models adding trade volume as a regressor in the conditional variance equation have been reestimated. The coefficients of the ARCH (GARCH) regressors do not become insignificant and the coefficient of volume is not significantly different from zero.

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