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The Impact of News on the Exchange Rate of the Lira and Long-Term Interest Rates

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SINTESI

Il contenuto di questo lavoro esprime solamente le opinioni degli autori; pertanto non rappresenta la posizione ufficiale della Banca d'Italia.

Il lavoro fornisce un contributo alla letteratura economica empirica sui prezzi delle attività finanziarie nel filone della "finanza comportamentale" (*behavioural finance*) con uno studio dal carattere innovativo sugli effetti del "rumore" politico e del rilascio di dati economici sulle variazioni del tasso di cambio e del tasso di interesse a lungo termine. La "finanza comportamentale" analizza l'interazione tra variabili finanziarie e azioni degli agenti economici dal punto di vista microeconomico; in questo ambito, la diffusione di notizie inattese dal mercato (cosiddette politico-economiche) o la pubblicazione di valori di variabili macroeconomiche diversi da quelli attesi provocano una variazione dell'insieme di informazioni disponibili agli agenti economici determinando un mutamento del processo di generazione delle variabili monetarie e finanziarie.

Il lavoro analizza l'impatto delle notizie di carattere politico ed economico su alcune variabili del mercato valutario e finanziario italiano (tasso di cambio della lira rispetto al dollaro, tasso di rendimento del Btp decennale e prezzo del corrispondente contratto *futures*), con particolare attenzione all'effetto sulla variabilità. Il periodo in esame, comprendente episodi di elevata turbolenza finanziaria e politica, è caratterizzato dal succedersi di tre governi. Esso copre il periodo che va dal marzo 1994, che coincide con la fine di una serie di governi "tecnici", al novembre 1996, mese in cui si pose fine alla sospensione della partecipazione della lira agli Accordi Europei di Cambio.

Le notizie politiche ed economiche (raccolte dai maggiori quotidiani nazionali e internazionali nonché dalla Reuters) sono classificate secondo la prevedibilità del loro rilascio: <u>inattesa</u> (perlopiù notizie politiche, come dichiarazioni di fonti governative e dirigenti politici che forniscono informazioni rilevanti riguardo allo scenario economico e politico) e <u>attesa</u> (rilascio periodico a date prefissate di dati macroeconomici ritenuti variabili fondamentali dagli agenti economici, come il tasso di inflazione, i prezzi della produzione industriale, i tassi all'emissione dei titoli pubblici ecc.). Prima del novembre 1996 in Italia l'influenza delle notizie politiche sui prezzi delle attività finanziarie è stata più marcata di quella registrata negli altri paesi industriali presumibilmente a causa della profonda trasformazione che si stava verificando nell'assetto politico italiano. Ugualmente, il deterioramento del disavanzo pubblico, che ha coinciso con la crisi in

Messico del marzo 1995, ha reso i mercati finanziari italiani particolarmente sensibili alla diffusione di dati relativi al deficit pubblico, al tasso di emissione dei titoli del Tesoro a breve termine, al tasso delle operazioni pronti contro termine condotte dalla Banca centrale, al tasso di inflazione osservato preliminarmente nelle città campione.

L’analisi è divisa in due parti. La prima esamina l’impatto di ogni singola notizia politica ed economica sulle variazioni percentuali di ogni variabile finanziaria e sulla loro variabilità condizionale. Nella seconda fase, le notizie che sono risultate significative sono aggregate in alcune variabili *dummy* composte e sono utilizzate in un modello multivariato a varianza condizionale autoregressiva generalizzata (*Generalized Autoregressive Conditional Heteroskedastic model, Garch*). Si esamina inoltre la persistenza di una notizia negativa sulla variazione del prezzo delle attività finanziarie tramite una funzione di risposta a un impulso.

I risultati mostrano che le notizie influenzano sia la media sia la variabilità condizionale delle variazioni giornaliere del tasso di cambio della lira rispetto al dollaro, del rendimento del Btp decennale e del corrispondente contratto *futures*. Inoltre, si evidenzia che esiste un significativo cambiamento di regime nella variabilità non condizionale delle variabili finanziarie durante i tre governi che si sono succeduti nel periodo in esame. Infine, diversamente dall’opinione diffusa, l’impatto delle notizie politiche e dei dati macroeconomici è stato più pronunciato sulla variabilità condizionale del tasso di cambio della lira rispetto al dollaro che su quella del tasso di interesse sul Btp decennale.

THE IMPACT OF NEWS ON THE EXCHANGE RATE OF THE LIRA AND LONG-TERM INTEREST RATES

by Fabio Fornari*, Carlo Monticelli*, Marcello Pericoli* and Massimo Tivegna**

Abstract

This paper analyzes the impact of news on several Italian financial variables, paying particular attention to the effect on the conditional volatility of these variables. The analysis spans a period of great financial and political turbulence in Italy, including the rapid succession of three governments. News releases (articles on political and economic events collected daily from both the Italian and international economic press) are classified as unscheduled (mostly political) and scheduled (i.e. economic and monetary statistics whose announcement is expected by market participants). The analysis is divided into two phases: first, we estimate the impact of each single political and economic news item on asset price changes and their conditional variance; second, those items that are identified as significant in the first stage are then aggregated into six dummies according to their nature and origin and employed as exogenous variables in a trivariate Garch scheme. Results show that i) news affects both the first and the second moment of the daily changes in the analyzed variables; ii) there is a significant regime shift of the unconditional variance of the analyzed variables across the three different governments; iii) the conditional variances display a significant — albeit rather small — seasonal dayweek pattern; iv) contrary to the conventional view, the impact of news on the conditional variance is more pronounced for exchange rates than for Italian long-term interest rates.

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1. Introduction[1]

This paper contributes to the *behavioural finance*[2] literature on asset pricing with an original novel study on the impact of political noise and economic data releases on exchange and interest rates. It provides a quantitative assessment of the impact of news, both expected and unexpected, on the Italian financial markets between 1994 — just after political elections were held under a new bi-partisan institutional setup — and 1996, when the Italian lira re-entered the European Exchange Rate Mechanism (ERM).

The Italian economic and political environment providing the background to our analysis is quite extraordinary in relation to experiences typically investigated by the literature in this field. Before the Italian lira re-entered the European exchange rate mechanism (ERM), the influence of political news on asset prices in Italy was much stronger than in the recent experience of any other developed country, mainly due to the severity of the various episodes of political crisis and to the resulting deep transformations in Italian politics (see the Appendix). Throughout this paper, the influence of unscheduled and scheduled news are kept separate, since the effects on the market can be expected to be very different. Political news is, by its very nature, unscheduled; inflation news, on the contrary, is generally scheduled in terms of official announcements, though we also considered news items relevant to price developments that are unscheduled, such as nation-wide wage settlements or news on commodity prices. Following this approach, interest rate news related to T-Bill auctions is scheduled, while the rate on repo operations by the Bank of Italy is not strictly scheduled, even though its timing can be generally anticipated with accuracy. News on public finance, so important for the outlook on Italian economic conditions, is generally unscheduled. For scheduled news, we used *surprise* variables, defined by the difference between actual and expected values, a quite common practice in the literature (e.g. Balduzzi, Elton and Green, 1998). In this paper expected values are proxied by the preceding value of the relevant series,

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[1] The authors wish to thank all the participants at the Link Fall meeting held in Rio de Janeiro, 14-18 September 1998, especially Jorge Braga de Macedo and Cesar Fuentes, as well as the participants at the QMF98 conference held in Sydney, 14-17 December 1998. Our gratitude goes also to the referee and to Gwyneth Schaefer.

[2] The literature on *behavioral finance* highlights, among other things, the role of information arrival in the mechanism of asset pricing, among other things; for an extensive introduction see Thaler (1993).

which amounts to assuming a static mechanism for the process according to which expectations are formed; this is also consistent with the hypothesis of unit root in the generating process of macroeconomic variables. The impact of news on exchange and interest rates is measured within a multivariate Var-Garch scheme (i.e. a vector autoregressive model with conditionally heteroskedastic errors), a novel approach in this literature.

Several studies have already investigated the role of economic news in the determination of financial asset prices — particularly with reference to the North American and Japanese financial markets. News items typically considered in the literature concerns economic activity (unemployment, industrial production, GDP growth, retail sales, business climate), inflation (CPI, PPI, wage developments), balance of payments (trade and current account) and changes in official interest rates; in studies which focus more specifically on the stock market, news items regarding individual firms are also collected (e.g. Mitchell and Mulherin, 1994). These analyses find that excess returns and return variability are directly related to the diffusion of news. For example, Ederington and Lee (1993) evidence that a systematic pattern can be found in the conditional volatility of financial prices, i.e. a significant increase (decrease) the day before which (the very day in which) scheduled news is received by market participants; they also find that the announcement of unscheduled news increases price variability. In a recent work, Balduzzi, Elton and Green (1998) use expectations data to calculate *surprise* economic announcements, thus differentiating their analysis from previous studies which focused on price volatility only and used dummy variables as regressors. They find that several economic *surprises* have a significant impact on US bond prices, with Nonfarm Payroll being the most prominent; they also find that price adjustment is extremely quick, completed on average within a few minutes. Almeida, Goodhart and Paine (1998) perform an intradaily analysis of the Deutsche Mark-US dollar reaction to publicly announced macroeconomic information emanating from the two countries; they find that most announcements affect the movements of the exchange rate for fifteen minutes and that this effect decreases as the interval over which it is measured increases. The literature on the relation between news and exchange rates follows the monetary model scheme of exchange rate determination. Edwards (1983a, 1983b) is the first author to propose a news-driven scheme, deriving it from a structural monetary model. In this framework news is made up by the error terms of a reduced form monetary model. His approach has been replicated in various papers, either within purely autoregressive models aimed at generating the expected values of the relevant macroeconomic variables (Copeland, 1984; Edwards, 1983a; Bomhoff and Korteweg, 1983) or for the same purpose within structural VAR models of the economy (Branson, 1983; Edwards, 1983a; MacDonald, 1983; Fiorentini, 1994). A different kind of literature — closer in its approach to this paper — uses the differences between market expectations (polled and published by various organizations) and the actual values of the relevant variables within a reduced-form model of the exchange rate determination. In these studies, the data frequency utilized is either daily (Deravi, Gregorowicz and Hegi, 1988; Hardouvelis, 1988; Irwin, 1989; Hogan, Melvin and Roberts, 1991; Doukas and Lifeland, 1994) or intradaily (Ito and Roley, 1988; Ederington and Lee, 1996; Baestaens and Van den Bergh, 1996). Other less directly related works are Agmont and Findlay (1982), Diamonte, Liew and Stevens (1996), Erb, Harvey and Viskanta (1986), which study the relation between political risk and stock prices, and Allvine and O'Neil (1980), Hobbs and Riley (1984) which study the impact of presidential elections on stock markets. Finally, Tivegna (1996a, 1996b) does some first explanatory work in the investigation of the relationship between Italian political news and the Lira exchange rate. More recently, the deep financial crises in a number of Southern American and Far Eastern Asian economies have led to papers which assess the domestic and *cross-border* impact of economic and political news, thus adopting a methodology similar in spirit to the one in this paper (Calvo, 1996; Baig and Goldfajn, 1998).

Since most recent studies generally use intradaily frequencies, the reader may be puzzled by the fact that our paper uses daily observations, especially since the effects of scheduled news are reportedly short-lived, in some cases lasting only a few hours, and the aggregation imposed by the use of daily observations might therefore seriously hide the impact of such news on price changes. Balduzzi, Elton and Green (1998) are the only authors who compare the two situations, namely the effect of news in intradaily and daily settings. In their paper, despite what one might assume, five out of the ten news announcements found to be significant at an intradaily level remain so with daily data.[3] Utilizing daily observations is strongly motivated by the frequency with which political news is made available and by the

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[3] Consumer prices, initial jobless claims, housing starts, producer prices and M2 lose their significant impact on note and bond price changes.

importance of considering such information. In fact, despite the abundance of analyses, very few works focus on the impact of political events or, more generally, the impact of political noise on financial asset prices; the collection of papers on *noise* in financial markets edited by Thaler (1993) includes only one paper (Cutler, Poterba and Summers, 1993) explaining the effect of political news on the US stock market.

This paper is organized as follows. Section 2 discusses the distinction between scheduled and unscheduled news as well as the methodology adopted for the definition of the event variables which are key to our quantitative analysis (also see the Appendix for this classification). In Section 3 the significance of these variables is individually tested in an univariate Garch setting for the exchange rate of the lira vis-à-vis the Deutsche mark. In Section 4 the event dummies are aggregated into six variables that are then employed in a trivariate Garch model of the mark/dollar exchange rate, the lira/mark exchange rate and the interest rate of the Italian ten-year Government bonds (BTP); in another specification of the same model, this last variable is replaced by the price of its futures contract. In this setup, regime shifts of both the unconditional and the conditional variance as well as dayweek seasonality are explicitly taken into account. Next, the multivariate model is employed to simulate — by means of an impulse response framework — the effect of shocks hitting the mean and the variance of financial asset prices. Section 5 concludes.

2. News

This section classifies news according to its different sources. During a typical workday, investors, dealers and traders are hit by a massive amount of information: news in general, releases of macroeconomic data, rumours concerning political and economic events. Following Ederington and Lee (1996) news is classified into two categories: news that is not expected or whose time of release is uncertain, henceforth *unscheduled*, and news whose release is expected by the market, henceforth *scheduled*. To provide an example, a declaration made by the Treasury Minister in a TV-interview on the tax treatment of bonds is unexpected and will almost certainly affect the market, with an impact on both bond prices and their volatility. Conversely, there is news[4] concerning macroeconomic data or the results of financial auctions whose announcement is scheduled (typically with a fair degree of precision) and thus expected by market participants. For example, a sharp unexpected variation in the preliminary consumer price index is important insofar as it can change the stance of monetary policy.

2.1 Unscheduled news

Unscheduled news is an economic or institutional event, declaration or disclosure that is either totally unexpected or, if expected, then occurring at an unknown time. This news is therefore not likely to be fully embodied in observed prices. The acquisition and analysis of the latest available information is central to the daily activity of trading financial assets. Such keen attention to the latest available information stands in sharp contrast to standard theories of asset pricing that hinge on the relationship between prices and fundamental variables. However, as the emerging approach of *behavioural finance* has brought to the fore, market participants prefer to follow information coming from the latest news rather than forecasting future developments in fundamental economic variables, at least in a short-term perspective.

In order to define variables expressing the release of unscheduled news, reference is made to headlines from the *Financial Times* and *Il Sole-24 Ore*, the leading Italian financial newspaper, as well as the Reuters terminal. A detailed description of the methodology used for the specification of the news variables (henceforth also referred to as *dummies*) is provided in the Appendix.

2.2 Scheduled news

Scheduled news encorporates time series of Italian macroeconomic data that are described in Table 1.[5] Since repo operations are the instrument used by the central bank to fine tune liquidity conditions, their announcement provides a clear signal of the Bank of

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[4] The word *news* is slightly inappropriate in this context since scheduled news in our classification is a public release of economic data or of the results of a scheduled T-bill auction or of the interest rate resulting from a semi-scheduled (but widely expected by the market) repo operation by the central bank.

[5] Each series equals nil when there is no release or the value of the relevant variable when data are released.

Italy's short-term monetary intentions. The average rate observed at the auction of three-month T-bills is the *pivot* rate at this maturity. Preliminary data on CPI comes from a sample group of Italian towns and are commonly used as a leading indicator of inflation.[6] Similarly, final data on the PPI index are followed as a leading indicator of the CPI, which is the principle measure of inflation.

3. Univariate preliminary analysis

3.1 The model

The variables representing both scheduled and unscheduled news, discussed in the previous section and in the Appendix, are inserted in the specification of a standard univariate Garch model of the lira vis-à-vis the Deutsche mark exchange rate; here, the daily logarithmic changes of the exchange rate (*dmit*) are modelled as a first order autoregressive process with autoregressive conditional variance. In order to avoid having the same regressor in both the mean and the variance equations, the dummy variables are used as a predetermined regressor in the mean and in the variance equations in turn.[7] Denoting by x_t the logarithmic change in the exchange rate, these two Garch models are defined as:

the mean model

$$\ln(dmit_t / dmit_{t-1}) = x_t = \mu_0 + \phi x_{t-1} + \theta_1 \ln(BTdmit_{t-1} / dmit_{t-1}) + \theta_2 \ln(BTdmus_{t-1} / dmus_{t-1}) + \mu_1 News_t^{+} + \mu_2 News_t^{-} + \varepsilon_t$$

$$\varepsilon_t | I_{t-1} \sim N(0, \sigma_t^2)$$

$$\sigma_t^2 = \omega + \alpha\varepsilon_{t-1}^2 + \beta\sigma_{t-1}^2 ,$$

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6 In 1997 the number of cities included in the sample group was increased from 10 to 11. This preliminary index accounts for over 80 percent of the final CPI. Hence its figures are often confirmed in the final release.

7 We report some evidence in favour of conditional heteroskedasticity. The daily rates of change of exchange rates, of the BTP (Buoni del Tesoro Poliennali) yield and the BTP futures price exhibit a remarkable clustering of anomalous observations. Their squares show significant autocorrelation up to the 20-th lag. The distributions of the rates of change of the four series is leptokurtic; analogous indications come from the coefficients of skewness and kurtosis and from the Jarque-Bera test, omitted for brevity.

or

the variance model

$$\ln(dmit_t / dmit_{t-1}) = x_t = \mu_0 + \phi x_{t-1} + \theta_1 \ln(BTdmit_{t-1} / dmit_{t-1}) + \theta_2 \ln(BTdmus_{t-1} / dmus_{t-1}) + \varepsilon_t$$

$$\varepsilon_t | I_{t-1} \sim N(0, \sigma_t^2)$$

$$\sigma_t^2 = \omega + \alpha\varepsilon_{t-1}^2 + \beta\sigma_{t-1}^2 + \lambda_1 News_t^+ + \lambda_2 News_t^-,$$

where *BTdmit* denotes the exchange rate of the lira vis-à-vis the Deutsche mark observed by Bankers Trust at 20.00 GMT in New York and *dmit* denotes the same rate as observed by Banca d’Italia at 13.15 GMT in Italy. Analogous notation holds for *BTdmus* and *dmus* which record the Deutsche mark-US dollar exchange rate observed by Bankers Trust and Banca d’Italia respectively. In the case of unscheduled news, $News^+$ is a dummy which equals 1 if the news is supposed to have had a negative impact on the exchange rates and 0 otherwise; $News^-$ is a dummy which equals -1 if the news is supposed to have had a positive effect and 0 otherwise. Regarding scheduled news, $News^+$ ($News^-$) is a dummy which equals the difference between the current and the previous value of the variable, the latter being an approximation, as previously discussed, of the best forecast made by rational agents, and is nil when no data are released; I_t is the information set available at time *t*.

The specification of the two models is best appreciated with reference to the timing of events described in Figure 1. The endogenous variable, namely $ln(dmit_t/dmit_{t-1})$, spans the segment AC; its lagged value the segment CD; the revisions of the exchange rate, namely $ln(BTdmit_{t-1}/dmit_{t-1})$, span the segment BC, overlapping with the endogenous variable. For this reason, exchange rate revisions are included as an additional regressor in order to *clean* the information set of what occurred the afternoon of the previous day in the US market. This permits a clearer interpretation of the relationship between news and the exchange rate over the morning of day *t*, insofar as it prevents attributing to news released on day *t* effects which instead originate from news released the previous afternoon and already incorporated in the quotations of New York. Similarly, the variable $ln(BTdmus_{t-1}/dmus_{t-1})$ *cleans* the information set from events concerning the mark-dollar exchange rate that occurred the previous afternoon in the US market.

3.2 Results

3.2.1 Unscheduled news

In the estimation of the mean model and the variance model each single dummy has been inserted, in turn, in the equation for changes in the lira-mark exchange rate and in the equation for the conditional variance of such changes.[8] The estimated coefficients are stable across equations and similar for the two models (Table 2a; results will be commented mostly qualitatively in the text). The constant of the mean equation is not statistically different from zero, indicating the absence of a positive drift in the daily returns; the lagged dependent variable is on average not significant. The lagged difference between the exchange rate observed in New York at 20.00 GMT of the previous day and in Rome at 13.15 GMT is very significant, with a coefficient slightly below 1; this follows naturally from the considerations put forward in the previous section (see Figure 1). The daily change in the lira-mark exchange rate is negatively correlated with the lagged difference between the quotation of the mark-US dollar exchange rate observed in New York at 20.00 GMT and in Rome at 13.15 GMT; this negative correlation is consistent with the empirical evidence of a relationship between the strengthening of the lira against the mark and the strengthening of the dollar against the mark (the estimate of θ_2 implies an elasticity of -6.5 percent). Results support the utilization of the Generalised Autoregressive Conditionally Heteroskedastic process of the first order for the variance equation: the parameters of the latter — the ω's, α's and β's estimated for each univariate model — are all statistically significant, and the persistence of the model, measured by the average value of $(\alpha + \beta)$, equals 0.97, thus implying a half-life for shocks of close to 22 days.[9]

In the mean model, the coefficient μ_1 records the impact of bad news and μ_2 the impact of good news on the daily change in the lira-mark exchange rate, with a positive sign

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[8] In econometric practice it is customary to estimate a single-equation model in a top-down procedure. In practice one starts by estimating the full model and sequentially eliminates the unsignificant variables. However we are only interested in measuring that part of the variance of a financial series that can be attributed to each specific news items. This is the reason why we abandon the classical approach and follow, instead, a sequential estimation procedure. The individual variances due to individual news items will then be aggregated into an overall dummy. We thank a referee for raising this point.

[9] Half-life is a concept typically encountered in the stochastic volatility literature. It is defined as minus $\ln(2)/\ln(\alpha+\beta)$ with α and β denoting the parameters of a Garch(1,1) model and is employed to evaluate the period of time it takes half of a shock of size ε_t hitting the conditional variance at time *t* to be absorbed.

expected for the coefficients of both variables given the opposite sign used in the definition of the variables themselves. Estimates show that all of the bad news items have the correct sign, though only one half of them are statistically different from zero. Significant bad news includes public finance issues, institutional conflicts, political conflicts during the first government in our analysis, EMU issues and the debate on political corruption. Good-news dummies with statistically significant coefficients include public finance issues, consumer prices, political conflicts during the first government, electoral results and political events occurring during the second government and since the beginning of 1996. Bad news about inflation does not seem to have an impact on the exchange rate, but good news does. Electoral results and the easing of political tensions in 1995 positively affected the daily exchange rate return. Political events occurring in January-February 1996 and events relating to the second government in 1995 are associated with a positive effect on the exchange rate but not with a negative one. Bad news on EMU impacts negatively on exchange rates, but good news does not. In analogy with the results for the mean, the variance shows a remarkable asymmetry between the effects of bad and good news. Most good news items do not seem to have an impact on the volatility of the daily exchange rate return, with the exception of electoral results, easing of political tensions and news related to the second of the three governments. For bad news, public finance, consumer prices, institutional conflicts, political conflicts, the second government and the dismissal of the Justice Minister in the second government are all highly significant. The estimation of the mean and the variance models evidences that, in line with expectations, both bad and good news have an impact on the exchange rate, but only bad news has a significant impact on volatility.

3.2.2 Scheduled news

Scheduled news resembles unscheduled news concerning the impact on the conditional mean and the conditional variance (Table 2b; as in the case of unscheduled news, the results are commented upon mainly qualitatively in the text). In the model for the conditional mean, positive changes in the repo rate weaken the lira-mark exchange rate, while negative changes do not affect it. This contrasts with the traditional Mundell-Fleming scheme, according to which higher domestic interest rates lead to an appreciation in the exchange rate and could support an *increased-risk-premium* argument for Italian financial assets (whereas higher rates bring about an increase in the interest payment on public debt). This effect could instead derive by lower than expected official interest rate moves. The latter effect, i.e. negative repo rate changes being neutral with respect to the lira-mark exchange rate, could reveal that such moves have been already fully discounted by the market. In the case of the average T-bills rate, increases are found to strengthen the lira; once again, negative changes do not affect the exchange rate. It is important to recall that seeking traditional relations among economic variables in a high-frequency context does not *per se* lead to conclusive results since a relation may be positive and significant at a daily frequency and negative or not significant on a monthly or quarterly basis. In any case, the findings concerning the relation between interest and exchange rates above are not new to the high frequency approach to international finance: the positive relation between changes in the repo rate and exchange rate is generally assumed to follow traders' expectations, whereas the negative relation between the T-bill rate and the exchange rate can be attributed to portfolio adjustments of long-term investors.

Neither positive nor negative changes in the final consumer price index affect, when released, the exchange rate, possibly because final CPI data can be accurately anticipated on the basis of the preliminary release, which occurs approximately 15 days earlier. In contrast, preliminary CPI and PPI releases do have an impact on the exchange rate.[10]

In the conditional variance equation similar results are obtained. CPI releases do not affect the volatility of the exchange rate, but when preliminary data on CPI and PPI are released and when they signal that inflation is rising, volatility increases. Releases of decelerating preliminary CPI and PPI data do not have any impact on the exchange rate. The same results hold true for the repo and T-bill interest rates: positive changes increase the volatility of the exchange rate; negative changes do not.

The comparison of the results obtained for the mean and the variance models points out that only increases in inflation, as signalled in the release of preliminary CPI data and final PPI data, have an impact on the change and the volatility of the lira-mark exchange rate. The same holds true for changes in the repo and T-bill interest rates: only positive figures count;

[10] This seemingly confusing result could be the econometric proof of the well-known financial market rule *buy the noise, sell the fact*. This means that, in short term trading, market participants are likely to act on market rumours and to unwind their position as soon as the fact, to which the rumours refer, materializes.

negative changes have no effect. CPI releases are ineffective in both models since they are anticipated by market participants on the basis of their preliminary values.

So far, this analysis has excluded the presence of any regime shift in the conditional and the unconditional variance among the three different governments as well as that of any seasonal dayweek pattern in foreign exchange and interest rates. In the traditional financial literature, this phenomenon is reported to be highly significant, especially with respect to stock returns, and it will be analyzed in detail in the following section.

4. Multivariate analysis

This section is devoted to the estimation of a multivariate Garch model that includes not only the lira-mark exchange rate but also the mark-dollar exchange rate and the yield on Italian long-term government bonds (BTP). Moving from a univariate to a multivariate model is necessary for many reasons: first, exchange rates are determined simultaneously in the market and this is especially true for lira exchange rates which are well-known to be driven by movements in the mark-dollar rate. Hence, though a reduced-form approach is adopted throughout this paper, a seemingly unrelated regression model is the appropriate tool to accommodate non-scalar covariance matrices. Second, given the heteroskedastic nature of financial data, one cannot guarantee that the correlation matrix remains fixed through time, a requisite that can be appropriately tested within a multivariate Garch scheme. Last, it is only within a multivariate framework that one can carry out causality tests or evidence the presence of common factors in a set of individual variance series.

An initial step of the analysis is devoted to reducing the number of news variables, so as to keep the dimension of the multivariate system within manageable limits. This is important for two main reasons: first, an increase in the number of explanatory variables would slow the maximization algorithm as well as affect its accuracy; second, since dummies have many zero entries, multicollinearity could play a prominent role in the significance of the estimates, albeit the correlation matrix of the news variables has no entry larger than 0.36 in absolute value.

The dummies found to be statistically significant in the univariate analysis of the lira-mark exchange rate were aggregated into six variables: two summarise the effect of unscheduled news, both positive and negative, on exchange rate levels; another two summarise the effect of unscheduled news on the variance of the latter; the last two summarise the effect of (positive) scheduled news on the level and the variance of the exchange rate. With regard to the latter two variables, the distinction between positive and negative news events is maintained under the hypothesis that a positive news item raises the variance less than a negative news item of comparable importance. For this to hold, since both dummies in the variance equations take unit values, the difference between their respective coefficients must be significant, revealing an asymmetric reaction of the absolute price changes, as originally proposed by Black (1976).

Following this approach, the six dummies used in the multivariate analysis are defined on the basis of the estimates of μ and λ in the univariate framework, namely $\hat{\mu}$ and $\hat{\lambda}$. With reference to the *k*-th univariate equation (with k=1,2,...,12 for *DL* and *DV* and k=1,2,..,5 for *QL* and *QV*), the synthetic dummies are defined as:

$DL^{+} = \sum_{k=1}^{12} \hat{\mu}_{k1} News_k^{+}$ with $\hat{\mu}_{k1}$=0 if not significantly different from zero in the univariate case;

$DL^{-} = \sum_{k=1}^{12} \hat{\mu}_{k2} News_k^{-}$ with $\hat{\mu}_{k2}$=0 if not significantly different from zero in the univariate case;

$DV^{+} = \sum_{k=1}^{12} \hat{\lambda}_{k1} News_k^{+}$ with $\hat{\lambda}_{k1}$=0 if not significantly different from zero in the univariate case;

$DV^{-} = \sum_{k=1}^{12} \hat{\lambda}_{k2} News_k^{-}$ with $\hat{\lambda}_{k2}$=0 if not significantly different from zero in the univariate case;

$QL^{+} = \sum_{k=1}^{5} \hat{\mu}_{k1} News_k^{+}$ with $\hat{\mu}_{k1}$=0 if not significantly different from zero in the univariate case;

$QV^{+} = \sum_{k=1}^{5} \hat{\lambda}_{k1} News_k^{+}$ with $\hat{\lambda}_{k1}$=0 if not significantly different from zero in the univariate case.

The variables defined above (QL^- and QV^- were not evaluated since the five individual components were identically zero) are inserted in the specification of a trivariate system where the logarithmic change in the mark-dollar (*dmus*) and lira-mark (*dmit*) exchange rates, and the changes in the BTP yield (*btpy*) are modelled as a Vector autoregression where the variance of the error of each equation follows a Garch process. The synthetic dummies DL^+, DL^-, DV^+, DV^-, QL^+, QV^+, already identified with respect to the lira/mark exchange rate, were employed also as a measure of the impact of news on the BTP price and yield changes.

In a subsequent phase, so as to account for the presence of regime shifts in the level of the unconditional variance, the multivariate Garch models were estimated again, introducing three dummies (each of them equalling unity in the time interval spanned by one government and zero in the time interval spanned by the remaining two) in the specification of the intercept in the equations for the conditional variance (see equations (6) and (7) reported in the model presentation (1)-(10) below). The same dummies were employed to evidence changes in the coefficients of the conditional variance equations, a signal of the modifications of the latter following the arrival of news. To identify the presence of seasonal heteroskedasticity, the individual dummies in the univariate Garch models presented in the last section have been replaced by the residual of a regression of each individual dummy on the five weekday dummies; the new synthetic dummies for the conditional mean (DL^+_R, DL^-_R, QL^+_R), which are immune from seasonality, are, as before, obtained as aggregation of the individual seasonally-adjusted dummies with weights provided by the associated coefficient in the new univariate Garch model. The multivariate schemes are then estimated again employing the seasonally-adjusted dummies in the conditional mean equations; as for the conditional variance equations, four dayweek dummies are employed along with the synthetic dummies DV^+, DV^-, QV^+ evaluated before.

4.1 Multivariate Garch models

The estimation of multivariate Garch schemes becomes very difficult when the dimension of the system is increased to five variables. Though most of the results referred to multivariate schemes revealed an almost perfect diagonality in the C, A, B matrices defined below, the full non-diagonal specification is employed to carry out a preliminary identification of the model and to test the hypothesis that the conditional correlation coefficients are stable over time, a situation which would simplify the estimation.

According to the formulation put forward by Engle and Kroner (1995), the multivariate Garch(1,1) can be written as:

$$y_t = \mu + \theta \cdot DL^+_t + \pi \cdot DL^-_t + \gamma \cdot QL^+_t + \Pi_t$$

$$\Pi_t \mid I_{t-1} \sim MN(0,H_t)$$

$$H_t = C' \cdot C + A' \cdot \varepsilon_{t-1} \cdot \varepsilon_{t-1}' \cdot A + B' \cdot H_{t-1} \cdot B$$

where C, A and B are $n \cdot n$ matrices and C', A' and B' are their transposes; H_t is the $n \cdot n$ time-varying covariance matrix; Π_t is the $n \cdot 1$ error term; θ, π and γ are $n \cdot 1$ vectors; DL^+, DL^- and QL^+ are the synthetic dummies of the mean equation as defined before. The synthetic dummies DV^+, DV^- and QV^+ are not inserted in the conditional covariance equations, since, at this stage, the analysis aims only at identifying the structure of the model, regardless of the effect of the temporary shocks induced by the occurrence of news on the conditional variance. For the trivariate model where y_t, the vector of independent variables, is made up of the (log) changes in the mark/dollar and the lira/mark exchange rates as well as the BTP yield changes, the assumption of diagonality cannot be rejected, since the off-diagonal terms of the C, A, B matrices were not statistically different from zero. The estimated matrices are reported below together with t-statistics (only the significant coefficients are reported; θ, π and γ are not shown since they are not relevant at this stage of the analysis):

$$A = \begin{pmatrix} .205(14.94) & 0 & 0 \\ 0 & .255(25.47) & 0 \\ 0 & 0 & .210(8.26) \end{pmatrix}$$

$$B = \begin{pmatrix} .979(385.71) & 0 & 0 \\ 0 & .969(368.21) & 0 \\ 0 & 0 & .963(124.4) \end{pmatrix}$$

$$C = \begin{pmatrix} 0 & 0 & 0 \\ 0 & 2.17 \cdot 10^{-4}(1.82) & 0 \\ 0 & 0 & 1.53 \cdot 10^{-3}(5.56) \end{pmatrix}.$$

Figures 2 to 4 show the estimates of the conditional correlations, the conditional covariances and the conditional standard deviations. The conditional covariances (Figure 3) unambiguously vary over time, with especially large changes around September 1994 and in the period from April to September 1995. However, the conditional correlations are rather stable through time, since variances change significantly when covariances do, lending support to the hypothesis that they vary proportionally to the conditional standard deviations (as one can appreciate comparing Figures 3 and 4). On the basis of these results, that hold also when the BTP yield is replaced by the BTP futures price, a trivariate diagonal Garch model with constant conditional correlations is estimated adopting the following structure:

conditional mean equations

(1) $\ln(dmus_t / dmus_{t-1}) = x_{1t} = \mu_{10} + \varepsilon_{1t}$

(2) $$\begin{aligned}\ln(dmit_t / dmit_{t-1}) = x_{2t} = \mu_{20} + \mu_{21}DL_t^+ + \mu_{22}DL_t^- + \mu_{23}QL_t^+ + \xi_1 x_{1,t-1} + \\ \theta_{21}\ln(BTdmit_{t-1} / dmit_{t-1}) + \theta_{22}\ln(BTdmus_{t-1} / dmus_{t-1}) + \varepsilon_{2t}\end{aligned}$$

(3) $$\begin{aligned}btpy_t - btpy_{t-1} = x_{3t} = \mu_{30} + \mu_{31}DL_t^+ + \mu_{32}DL_t^- + \mu_{33}QL_t^+ + \xi_2 x_{1,t-1} + \\ \theta_{31}\ln(BTdmit_{t-1} / dmit_{t-1}) + \theta_{32}\ln(BTdmus_{t-1} / dmus_{t-1}) + \varepsilon_{3t}\end{aligned}$$

distributional assumption

(4) $$\Gamma_t = \begin{pmatrix}\varepsilon_{1t}\\ \varepsilon_{2t}\\ \varepsilon_{3t}\end{pmatrix} \Big| \; I_{t-1} \sim N(\eta, H_t); \quad \eta = \begin{pmatrix}0\\0\\0\end{pmatrix}; \qquad H_t = \begin{pmatrix}\sigma_{1t}^2 & \sigma_{12t} & \sigma_{13t}\\ \sigma_{12t} & \sigma_{2t}^2 & \sigma_{23t}\\ \sigma_{13t} & \sigma_{23t} & \sigma_{3t}^2\end{pmatrix}$$

conditional covariance equations

(5) $\sigma_{1t}^2 = \omega_1 + \alpha_1(\varepsilon_{1t-1} / E_{t-1})^2 + \beta_1\sigma_{1t-1}^2$

(6) $\sigma_{2t}^2 = \omega_2 + \alpha_2(\varepsilon_{2t-1} / E_{t-1})^2 + \beta_2\sigma_{2t-1}^2 + \lambda_{21}DV_t^+ + \lambda_{22}DV_t^- + \lambda_{23}QV_t^+ + \psi_1\varepsilon_{1,t-1}^2$

(7) $\sigma_{3t}^2 = \omega_3 + \alpha_3(\varepsilon_{3t-1} / E_{t-1})^2 + \beta_3\sigma_{3t-1}^2 + \lambda_{31}DV_t^+ + \lambda_{32}DV_t^- + \lambda_{33}QV_t^+ + \psi_2\varepsilon_{1,t-1}^2$

(8) $\sigma_{12t} = \omega_{21}\sigma_{1t-1}\sigma_{2t-1}$

(9) $\sigma_{23t} = \omega_{23}\sigma_{2t-1}\sigma_{3t-1}$

(10) $\sigma_{13t} = \omega_{13}\sigma_{1t-1}\sigma_{3t-1}$,

where x_1, x_2 and x_3 are the three endogenous variables; DL^+ , DL^- , DV^+ and DV^- the synthetic dummies summarizing the impact of bad and good unscheduled news on the conditional mean and the conditional variance; QL^+ and QV^+ the synthetic dummies summarizing the impact of bad scheduled news on the conditional mean and the conditional variance;[11] ε_1, ε_2 and ε_3 three error terms conditionally jointly distributed as a multivariate normal with mean η and conditional variance H_t, which is a 3×3 symmetric positive definite matrix; finally E is a variable recording the number of days elapsed between subsequent observations. Equations (8) to (10) are the conditional covariances between the logarithmic rates of change of the mark-dollar and the lira-mark exchange rates, between the logarithmic change in the lira/mark exchange rate and the BTP yield, and between the logarithmic change in the mark-dollar exchange rate and the BTP yield, respectively. They are built so as to keep the conditional correlation coefficients fixed at ω_{21}, ω_{23} and ω_{13} respectively (Bollerslev, 1990). It is important to point out that equations (2)-(3) and (6)-(7) are structured in such a way as to allow the conditional means and the conditional variances of the lira/mark exchange rate and of the BTP yield (or the BTP futures price when (3) is replaced by (3')) to be Granger-caused by the conditional mean and variance of the mark/dollar exchange rate. This assumption allows us to test whether the frequently reported evidence that the Italian financial market is led — over long horizons — by developments in the US and German markets, also holds at shorter frequencies. The parameters are estimated by the numerical maximization of the likelihood function of the system.

The same multivariate analysis is replicated replacing the daily rates of change in the BTP with the daily logarithmic rates of change in the price of the BTP futures traded on LIFFE (*btpf*). This model permits the testing of the hypothesis that the futures market is more efficient than the spot market in processing information. The model is analogous to the previous model but replaces equation (3) with:

$$\begin{aligned}(3')\quad & \ln(btpf_t / btpf_{t-1}) = x_{3t} = \mu_{30} + \phi_3 x_{3t-1} + \mu_{31} DL_t^+ + \mu_{32} DL_t^- + \mu_{33} QL_t^+ + \xi_3 x_{1,t-1} + \\ & \theta_{31} \ln(BTdmit_{t-1} / dmit_{t-1}) + \theta_{32} \ln(BTdmus_{t-1} / dmus_{t-1}) + \varepsilon_{3t}.\end{aligned}$$

In this case all the parameters regarding the first two variables (the mark/dollar and the lira/mark exchange rates) have been kept equal to the values they had in the previous

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[11] In the univariate model the release of good news — i.e. the single dummies associated to the coefficients μ^- and λ^- — is not significant: hence the dummies DL^- and DV^- are nil and eliminated from the analysis.

specification (XXY; see Section 4.2) in order to avoid the misleading circumstance of having two sets of coefficients for the same equation.

4.2 Baseline model

We report hereafter the results of the model described in the previous section (Table 3, lines a)). One model is for the lira-mark and the mark-dollar exchange rates as well as the first difference of the 10-year Government bond yield (BTP). The other model replaces BTP yield with the daily rates of change in the price of its futures contract traded at LIFFE while — as previously stated — leaving the parameters concerning the first two variables unchanged. The two models are henceforth called mark-dollar/lira-mark/BTP-yield (XXY) and mark-dollar/lira-mark/BTP-futures (XXF).

There is no causal relation between the conditional mean and variance of the mark/dollar exchange rate and the conditional variance of the lira/mark rate and the BTP yield or the BTP futures: the coefficients ξ_1, ξ_2, ξ_3, ψ_1 and ψ_2 were not statistically significant, and were not shown in Table 2 in order to save space. Given the small t-statistics, such causal links were omitted from all subsequent modifications of the baseline model, though we preserved the trivariate nature of the econometric scheme.

In the XXY model, all the coefficients of the conditional variances are highly significant, lending support to the appropriateness of the Garch(1,1) scheme; in all cases the persistence of conditional variances is rather high, 0.955 for the BTP yield, 0.953 for the BTP futures, 0.904 for the mark/dollar and 0.990 for the lira/mark.[12] As concerns the mean equations, the lira-mark exchange rate yields coefficients that are similar to those obtained in the univariate case, with the exception of the lagged variable ln(*BTdmus/dmus*) whose coefficient is now not significantly different from zero. Unscheduled news, both positive and

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[12] A wide body of literature has examined the properties of the integrated Garch(1,1) model, a particular case of the traditional Garch(1,1) scheme, in which the sum of α and β (see equation (6)), is unity. This model is generally thought to be non stationary, since forecasts would tend to explode as long as the time horizon increases. Nelson (1990) has shown that the limiting behavior of σ_t^2 is determined by the intercept of the variance equation as well as by the sign of $E[\ln(\beta+\alpha z_t^2)]$ where $E(\cdot)$ is the expectations operator and z_t is an i.i.d. variable. In any case, the ergodicity condition differes from the stationarity condition. In any case, we are not interested in forecasts of the conditional variance; hence the issue concerning the presence of a unit root in its generating process is ignored in this paper.

negative, has a significant impact on the conditional mean of the lira/mark rates of change, while BTP yield changes are only affected by the occurrence of negative unscheduled news. Scheduled news affects the conditional variance of the lira/mark exchange rate; unscheduled news affects it only in the case of bad news. Positive unscheduled news affects the conditional variance of the BTP yield changes, negative news does not; scheduled news affects the variance of the BTP yield.

In the case of the XXF model it is important to recall that the futures are quoted in terms of price rather than yield and thus some of the coefficients have the opposite sign when compared to the XXY model. As concerns the futures price changes, unscheduled news affects the conditional mean, in the case of both good and bad news; positive scheduled news has no influence on it. The conditional variance is only affected by positive unscheduled news.

The multivariate analysis evidences a positive and sizeable relationship between the lira-mark exchange rate and the revision of the same variable which took place in the US (*Btdmit/dmit*). The same variable is also highly significant and sizeable in the case of BTP yield changes, with a coefficient close to 1, implying that the afternoon revision of the lira/mark exchange rate has full impact on the bond market; this link may be hard to explain with standard economic arguments. As in the univariate analysis, there exists a negative correlation between changes in the lira-mark exchange rate and changes in the previous afternoon's value of the mark-dollar exchange rate, as revealed by the coefficient θ_2; this implies that a weakening mark against the dollar in the (East Coast time) afternoon results in a strengthening lira against the mark throughout the next day (central Europe time).

Figures 5-7 show, for each of the three assets, the overall conditional standard deviation and the part due to the effect of the news. The contribution of news to the overall variability is not negligible in the case of the lira/mark exchange rate, especially in the period between March and May 1995; it is less so in the case of the BTP yield, averaging 5 percent, i.e. one third of the overall conditional standard deviation; it becomes very low in the case of the BTP futures price, averaging 2 percent, i.e. one fifth of the overall variability of the series, which could support the theory of greater informational efficiency in the futures market. To summarise, these results imply that news has a significant effect on the conditional means of the three variables. Concerning the conditional second moments, news has a relevant impact on the conditional second moment of the lira/mark exchange rate, a much less pronounced effect in the case of the BTP yield and a marginal influence in the case of the futures price, which could be informationally superior.

4.3 Regime shifts in the unconditional and the conditional variance

When a relevant event takes place, the parameters of an econometric model may change in reaction to the new market conditions. We tested for the existence of breaks in the conditional variance equations with reference both to the estimated unconditional variance, i.e. the level to which it reverts in the long-run, and to the persistence of the variance, i.e. its reaction function to the arrival of shocks.

To evidence the presence of shifts in the unconditional variance we modified equations (6) and (7) as reported in (6') and (7') below:

$$(6')\ \sigma^2_{2t} = \omega_2(1+\varphi_{21}FG_t + \varphi_{22}SG_t) + \alpha_2(\varepsilon_{2t-1}/N_{t-1})^2 + \beta_2\sigma^2_{2t-1} + \lambda_{21}DV^+_t + \lambda_{22}DV^-_t + \lambda_{23}QV^+_t$$

$$(7')\ \sigma^2_{3t} = \omega_3(1+\varphi_{31}FG_t + \varphi_{32}SG_t) + \alpha_3(\varepsilon_{3t-1}/N_{t-1})^2 + \beta_3\sigma^2_{3t-1} + \lambda_{31}DV^+_t + \lambda_{32}DV^-_t + \lambda_{33}QV^+_t\,.$$

Such equations are specified so that the intercept of the conditional variance is affected by the sample either being the overall period, the period in which the first government (FG) was in charge or the period in which the second government (SG) was in charge, respectively. Since the unconditional variance is a function of ω_i, α_i, β_i, λ_{i1}, λ_{i2} (i = 1, 2) as well as $\varphi_{i,1}$ - $\varphi_{i,2}$, it will change during the three different regimes when these last four parameters are significantly different from zero; their values are reported in Table 3, lines b). In the XXY model, the unconditional variance of the lira/mark exchange rate is only marginally higher during the first government relative to the whole sample, but does not change during the second and third; the unconditional variance of the bond yield changes is instead significantly different during the second government.[13] In the XXF model, the conditional variance of the BTP futures price is significantly different during the three governments, higher in the second and lower in the first; further, when the three regime dummies are included in the conditional variance specification, the effect of the synthetic dummies (DV^+, DV^- and QV^+) is no longer

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[13] See Bollerslev (1986) for the way in which the unconditional variance is obtained. In the standard Garch(1,1) it is obtained as $\omega/(1-\alpha-\beta)$.

significant, a further indication of the superior informational efficiency of the derivatives market, once one moves towards a more adequate specification of the model.

To test whether the persistence of the system is affected by the occurrence of news, the specification of the Garch schemes has instead been changed by replacing equations (6) and (7) with the following:

(6''') $$\sigma^2_{2t} = \omega_2 + \alpha_2(\varepsilon_{2t-1} / N_{t-1})^2 + \beta_2(1+\gamma_1 \cdot Ind_t)\sigma^2_{2t-1} + \lambda_{21}DV^+_t + \lambda_{22}DV^-_t + \lambda_{23}QV^+_t$$

(7''') $$\sigma^2_{3t} = \omega_3 + \alpha_3(\varepsilon_{3t-1} / N_{t-1})^2 + \beta_3(1+\gamma_2 \cdot Ind_t)\sigma^2_{3t-1} + \lambda_{31}DV^+_t + \lambda_{32}DV^-_t + \lambda_{33}QV^+_t,$$

where the coefficients β_2 and β_3 have been modified so as to allow for the presence of a shift coinciding with the occurrence of a news release. The variable Ind_t is built so as to be unity at the time of release of a news item (scheduled or unscheduled) and nil when there is no news. The results, reported in Table 3, lines d), show that the volatility of the lira/mark exchange rate is higher on days when both scheduled and unscheduled news occurs; in these cases the persistence of the model increases by 0.03. The generating process of the variance of the BTP yield and of the price of the BTP futures are again only marginally higher, by 0.01 and 0.06 respectively, when news occurs; however, the t-ratios of γ_1 and γ_2 are not significant.

4.4 Seasonality

In this section the multivariate Garch scheme is estimated again so as to account for the presence of a weekday seasonality. Compared to the model reported in (1)-(10), equations (2), (3), (6) and (7) have been replaced by the following (2''), (3''), (6'') and (7''):

(2'') $$\begin{aligned}&\ln(dmit_t / dmit_{t-1}) = x_{2t} = \mu_{20} + \mu_{21}DL^+_t\_R + \mu_{22}DL^-_t\_R + \mu_{23}QL^+_t\_R + \\ &\quad \theta_{21}\ln(BTdmit_{t-1} / dmit_{t-1}) + \theta_{22}\ln(BTdmus_{t-1} / dmus_{t-1}) + \varepsilon_{2t}\end{aligned}$$

(3'') $$\begin{aligned}&btpy_t - btpy_{t-1} = x_{3t} = \mu_{30} + \mu_{31}DL^+_t\_R + \mu_{32}DL^-_t\_R + \mu_{33}QL^+_t\_R + \\ &\quad \theta_{31}\ln(BTdmit_{t-1} / dmit_{t-1}) + \theta_{32}\ln(BTdmus_{t-1} / dmus_{t-1}) + \varepsilon_{3t}\end{aligned}$$

(6'') $$\begin{aligned}&\sigma^2_{2t} = \omega_2 + \alpha_2(\varepsilon_{2t-1} / N_{t-1})^2 + \beta_2\sigma^2_{2t-1} + \lambda_{21}DV^+_t + \lambda_{22}DV^-_t + \lambda_{23}QV^+_t + \\ &\gamma_1 D_{1,t} + \gamma_2 D_{2,t} + \gamma_3 D_{3,t} + \gamma_4 D_{4,t}\end{aligned}$$

(7'') $$\begin{aligned}&\sigma^2_{3t} = \omega_3 + \alpha_3(\varepsilon_{3t-1} / N_{t-1})^2 + \beta_3\sigma^2_{3t-1} + \lambda_{31}DV^+_t + \lambda_{32}DV^-_t + \lambda_{33}QV^+_t + \\ &\gamma_5 D_{1,t} + \gamma_6 D_{2,t} + \gamma_7 D_{3,t} + \gamma_8 D_{4,t}\end{aligned}$$

where, as before, x_1, x_2 and x_3 are the three endogenous variables; DL^+_R, DL^-_R, QL^+_R the synthetic seasonally-adjusted dummies summarizing the occurrence of bad and good news on the conditional mean; DV^+, DV^- and QV^+ are the same synthetic conditional variance dummies as before; D_1, D_2, D_3 and D_4 are four dummies equal to one on Monday, Tuesday, Wednesday and Friday, respectively, and nil elsewhere (the seasonal effect of Thursday is obtained from the difference between the intercepts ω_2 and ω_3 and the four dummies). The results of this estimation, reported in Table 3, lines c), evidence that a higher-than-average volatility is observed on Monday with respect to the lira/mark exchange rate, and on Wednesday and Friday with respect to the BTP yield; BTP futures are not subject to a statistically significant weekday seasonal pattern. Even though significant, the seasonal coefficients are so small that they do not represent an important component of the conditional variances. Despite this, it is interesting to note that the introduction of seasonal dummies eliminates the impact of news on the conditional mean of the BTP yield and reduces the impact exercised by the afternoon movements of the exchange rates; however, it adds something to the coefficients of the synthetic variables in the conditional variance equation. These results may signal that part of the impact of news on the conditional mean and variance of the BTP yield changes can be attributed to a fixed seasonal effect (however being impossible to discriminate between seasonality and the presence of a given news item occurring systematically in a specific day of the week), hence lowering the component of volatility specifically related to news.

4.5 Reaction of the conditional variance to adverse events

The multivariate Garch model can be used to assess the short term reaction of the conditional covariance matrix of the exchange and long-term interest rates to the occurrence of shocks — *news* or large destabilizing events in general. To this end we adopt an *impulse response* framework referred not to the level of the three variables but to their conditional second moments.

In the case of unscheduled news, the system estimated before (the baseline model) and specified in equations (1)-(10) is perturbed by two simultaneous negative shocks equal to five times the maximum of the historical standard deviations of the synthetic dummies DL^- and DV^- respectively — hence a rare event corresponding to the diffusion of very bad news.[14] The effects of the shocks are assessed by evaluating equations (1)-(10) recursively, keeping the parameters fixed at their full-sample values — and examining the path of conditional variances and covariances for fifty out-of-sample daily simulations. In order to measure the impact of a shock coming from scheduled news, since the variable QV^- was identically nil over the sample, the QV^+ variable was employed in its place (hence assuming no *leverage effect*); obviously, since QL^- was nil, there is no shock on the conditional mean of the exchange and long term interest rates in this case.

Figure 8 shows the effects of the shock on the conditional standard deviation of the lira/mark exchange rate and the BTP yield, as well as on the covariance between these two variables. The standard deviation of the exchange rate increases from 4.2 to 5.4 percent on the day in which the *extreme* news is recorded, a jump of nearly 20 percent, reasonably small compared to the dimension of the shock. After the initial impact, volatility starts to decline gradually, reaching the new equilibrium level in about one month; convergence to the new equilibrium of the BTP yield variance and of the covariance between the latter and the exchange rate is faster. The conclusions are basically unchanged for the impulse response functions in the case of scheduled news, as reported in Figure 9.[15]

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[14] The size chosen for the shock, five times the maximum historical standard deviation of the overall dummies, may seem a very rare occurrence. However, a wide body of literature has unambiguously shown that the distribution of financial price log-changes is leptokurtic (Bollerslev, Engle and Nelson, 1994), so that traditional inference is not adequate. In other words, the traditional $(\mu\text{-}3\sigma,\mu\text{+}3\sigma)$ interval around the mean of a series does not contain, as usually happens, 99 percent of observations.

[15] Since this paper focuses mainly on the relation between news and the conditional variance of financial price changes, there is no attempt to quantify the impact of news on the conditional mean of such price changes. This strategy, as a referee has observed, may lead the reader towards the conclusion that there is no significant impact of news on the first moment of financial prices. The papers which deal with identifying such effect have employed, at least very recently, intradaily data and have evidenced that the impact of news vanishes within one hour (see also the Introduction to this paper); also for this circumstance we focused on the second rather than the first order conditional moments. The interested reader can nonetheless recover the elasticity between news and price changes from the coefficients reported in Table 3. To provide a benchmark, when the five-standard-deviation shock hits the system (1)-(10), the difference between the simulated and the observed level of the lira/mark exchange rate is close to 1 percent (more or less 10 lire given the actual fixed parity in the euro).

5. Conclusions

This paper provides a quantitative assessment of the impact of the diffusion of news on the daily volatility of the lira exchange rate and long-term interest rates. The analysis spans a period — March 1994 to November 1996 — that was particularly turbulent in Italy for both political and economic reasons.

In this econometric analysis news is divided into two categories: scheduled (macroeconomic data released on dates and times known by market participants) and unscheduled (declarations by government officials and political leaders providing relevant new information about the economic and political outlook). Each unscheduled news item is transformed into a dummy variable taking the value plus (minus) one on the day of release according to whether it can be expected to affect negatively (positively) the economic outlook.

The first part of the econometric analysis is devoted to the identification of the scheduled and unscheduled news items that had a significant impact on the exchange rate (T-bill auctions, repo operations, preliminary information on CPI and PPI, final CPI figures and a number of other unscheduled news items were found to be significant) via a simple Garch model of the exchange rate of the lira vis-à-vis the Deutsche mark. In these models scheduled and unscheduled news were added either to the equation for changes in the exchange rate or to the equation for its time-varying second moment.

In the second part of the analysis, a trivariate Garch model is estimated for the lira/mark and mark/dollar exchange rates and, alternatively, for the long-term interest rates or the price of the futures contract on the corresponding benchmark government bond; this scheme is the main tool by which the impact of news is measured. An impulse response analysis of the conditional variance of the three dependent variables (exchange rates and BTP yield) to a very large shock (5 times the standard deviation of the dummies) to the arrival of news is also carried out on the basis of the estimates of the Garch model. The estimated models and the results of the impulse response analysis concord that news has a significant impact on the fluctuations in the lira exchange rate and long-term interest rates; seasonality and regime shifts are also evidenced in the time series of the conditional variance, effects which, once accounted for, decrease the overall effect of news. Contrary to widely-held views, the impact of news on the conditional second moments was quite small in size albeit statistically different from zero. This suggests that market participants react to the diffusion of news, generally regarding it as informative of permanent changes in fundamental variables rather than as erratic episodes of instability that permanently add to the uncertainty in price formation of lira-denominated assets.

Table 1

MACROECONOMIC ANNOUNCEMENTS (*)

Variable	GMT time of release	Day of the week/month	Reporting agency	Frequency
Average repo rate	irregular	the offer is announced with closed markets at 7:30 and done at 8:00, the repo rate is announced shortly after	Bank of Italy	-
3-month T-bill auction average rate	-	according to an announced auction calendar, around the middle and the end of the month	Bank of Italy	bi-weekly
Preliminary cities' CPI, M/M and Y/Y	16:30	two consecutive days between the 18th and the 22nd for same month data	Istat	monthly
Final CPI, M/M and Y/Y	8:00	one day between the 4th and the 6th for previous month data	Istat	monthly
Final PPI, M/M	8:00	one day between the 13th and the 19th for two months previous data	Istat	monthly

(*) CPI is Consumer Price Index; PPI is Producer Price Index; M/M is month-over-month and Y/Y year-over-year.

Table 2a

UNIVARIATE GARCH(1,1) FOR THE LIRA-MARK EXCHANGE RATE UNSCHEDULED NEWS CASE

dummy → param ↓	PF mean	PF variance	CP mean	CP variance	IC mean	IC variance	PC mean	PC variance	ER mean	ER variance	PE mean	PE variance
μ_0	-1.61e-5 (-0.12)	4.49e-5 (0.35)	1.13e-4 (1.00)	6.76e-5 (0.55)	3.80e-6 (0.03)	3.56e-5 (0.30)	1.49e-5 (0.12)	-4.42e-6 (-0.03)	5.08e-5 (0.42)	7.46e-6 (0.06)	5.77e-5 (0.49)	5.68e-5 (0.48)
ϕ	0.059 (2.35)	0.048 (1.53)	0.048 (1.82)	0.046 (1.54)	0.040 (1.38)	0.044 (1.51)	0.031 (1.18)	0.047 (1.78)	0.032 (1.17)	0.037 (1.27)	0.043 (1.46)	0.046 (1.52)
θ_1	0.900 (22.06)	0.933 (21.00)	0.914 (22.57)	0.937 (22.98)	0.926 (21.90)	0.934 (22.34)	0.900 (22.13)	0.914 (21.52)	0.926 (22.32)	0.917 (22.89)	0.931 (22.62)	0.936 (22.87)
θ_2	-0.113 (-4.81)	-0.067 (-2.54)	-0.099 (-4.35)	-0.066 (-2.92)	-0.075 (-3.29)	-0.068 (-2.99)	-0.070 (-2.95)	-0.075 (-2.83)	-0.072 (-3.23)	-0.085 (-3.10)	-0.067 (-3.03)	-0.066 (-2.99)
μ_1	8.27e-4 (2.59)	-	-2.04e-4 (-0.27)	-	1.69e-3 (3.03)	-	1.75e-3 (4.58)	-	5.92e-4 (0.40)	-	9.33e-4 (0.28)	-
μ_2	7.62-4 (2.06)	-	1.51e-3 (3.00)	-	1.98e-3 (0.52)	-	3.08e-3 (4.40)	-	1.83e-3 (3.88)	-	2.61e-3 (2.27)	-
$\sqrt{\omega}$	7.05e-4 (14.76)	6.01e-4 (5.36)	6.65e-4 (14.64)	8.59e-4 (12.19)	8.69e-4 (12.99)	7.92e-4 (13.07)	6.46e-4 (9.03)	5.84e-4 (6.46)	7.58e-4 (12.64)	7.94e-4 (9.16)	8.56e-4 (13.49)	8.83e-4 (13.55)
$\sqrt{\alpha}$	0.348 (18.67)	0.382 (15.26)	0.355 (18.21)	0.390 (14.20)	0.397 (15.07)	0.381 (14.95)	0.338 (16.45)	0.356 (17.15)	0.367 (16.81)	0.362 (15.14)	0.396 (15.96)	0.398 (15.74)
$\sqrt{\beta}$	0.924 (123.83)	0.918 (86.29)	0.925 (123.53)	0.898 (73.64)	0.898 (73.44)	0.909 (85.02)	0.931 (107.79)	0.927 (102.69)	0.915 (100.15)	0.908 (73.77)	0.900 (78.19)	0.895 (73.73)
$\sqrt{\lambda_1}$	-	-1.32e-3 (-6.22)	-	1.41e-3 (2.85)	-	-1.10e-3 (-2.29)	-	-1.33e-3 (-6.15)	-	2.76e-6 (0.00)	-	-2.28e-3 (-0.76)
$\sqrt{\lambda_2}$	-	-1.15e-6 (0.00)	-	-5.76e-4 (-1.07)	-	3.00e-6 (0.00	-	6.25e-7 (0.00)	-	-2.82e-3 (-5.88)	-	1.76e-3 (1.99)

dummy→ param ↓	GD1 mean	GD1 variance	S1 mean	S1* variance	QUM mean	QUM variance	CG mean	CG variance	DU17 mean	DU17 variance	MA mean	MA variance
μ_0	7.76e-5 (0.64)	3.45e-5 (0.30)	8.27e-5 (0.66)	4.65e-5 (0.36)	-7.09e-5 (-0.78)	6.68e-5 (0.51)	-3.42e-5 (-0.28)	5.34e-5 (0.44)	3.95e-5 (0.33)	2.23e-5 (0.18)	4.47e-5 (0.38)	5.07e-5 (0.43)
ϕ	0.038 (1.30)	0.032 (1.11)	0.036 (1.29)	0.049 (1.56)	0.015 (1.30)	0.061 (1.92)	0.032 (1.12)	0.044 (1.51)	0.037 (1.28)	0.042 (1.47)	0.047 (1.61)	0.046 (1.56)
θ_1	0.918 (20.93)	0.927 (20.32)	0.919 (22.67)	0.943 (21.72)	0.918 (22.78)	0.965 (21.89)	0.920 (22.81)	0.936 (22.68)	0.916 (22.18)	0.923 (22.77)	0.938 (22.97)	0.933 (22.43)
θ_2	-0.066 (-2.98)	-0.075 (-3.14)	-0.083 (-3.74)	-0.073 (-3.06)	-0.118 (-4.85)	-0.094 (-3.71)	-0.074 (-3.39)	-0.066 (-2.91)	-0.075 (-3.12)	-0.078 (-3.24)	-0.072 (-3.22)	-0.072 (-3.26)
μ_1	6.68e-4 (1.31)	-	8.81e-4 (1.67)	-	1.75e-3 (2.70)	-	2.26e-3 (3.63)	-	2.34e-2 (10.86)	-	2.72e-3 (1.80)	-
μ_2	3.13e-3 (4.32)	-	2.03e-3 (3.29)	-	-2.97-4 (-0.43)	-	1.99e-2 (0.00)	-	-	-	4.80e-3 (3.86)	-
$\sqrt{\omega}$	8.25e-4 (14.03)	1.04e-3 (14.11)	8.56e-4 (13.30)	9.44e-4 (12.29)	1.29e-4 (14.16)	1.01e-3 (12.09)	8.70e-4 (13.72)	8.76e-4 (13.36)	9.28e-4 (12.26)	1.04e-3 (14.05)	9.39e-4 (13.40)	9.28e-4 (13.82)
$\sqrt{\alpha}$	0.385 (13.37)	0.325 (8.40)	0.400 (15.85)	0.440 (14.63)	0.325 (20.95)	0.466 (14.47)	0.415 (15.46)	0.399 (15.39)	0.390 (15.11)	0.399 (15.72)	0.429 (15.65)	0.408 (15.51)
$\sqrt{\beta}$	0.906 (78.83)	0.882 (54.97)	0.897 (76.42)	0.885 (62.17)	0.954 (223.42)	0.874 (55.21)	0.892 (71.67)	0.898 (74.85)	0.893 (69.90)	0.876 (70.31)	0.881 (64.12)	0.886 (66.96)
$\sqrt{\lambda_1}$	-	-4.42e-3 (-6.80)	-	-1.11e-4 (-0.02)	-	-2.77e-6 (-0.02)	-	-3.60e-6 (0.00)	-	-1.58e-2 (-5.52)	-	2.51e-3 (3.00)
$\sqrt{\lambda_2}$	-	3.43e-3 (5.12)	-	-2.91e-6 (0.00)	-	3.74e-6 (0.00)	-	-4.50e-2 (0.00)	-	-	-	-3.88e-3 (-1.77)

PF = Public Finance, CP = Consumer Prices, IC = Institutional Conflicts, PC = Political Conflicts, ER = Electoral results, PE = Easing of political tensions; GD1 = Mr. Dini government, S1 = Political opposition easing, QUM = EMU issues, QUMIN = EMU and international environment, CG = Bribesville's debate, DU17 = Black Friday, 17 March 1995, MA = Justice Minister Mancuso removal.

In parantheses t-statistics ($t_{654,\alpha=0.25}$=1.96).

* Only from December 21, 1995 onwards.

Table 2b

UNIVARIATE GARCH(1,1) FOR THE LIRA-MARK EXCHANGE RATE SCHEDULED NEWS CASE

dummy → param ↓	REPO mean	REPO variance	T-bills mean	T-bills variance	CPI mean	CPI variance	prel CPI mean	prel CPI variance	PPI mean	PPI variance
μ_0	-2.80e-5 (-0.22)	6.72e-5 (0.50)	9.23e-5 (0.77)	2.48e-5 (0.20)	7.69e-5 (0.65)	-5.76e-5 (-0.47)	8.71e-5 (0.75)	4.97e-5 (0.39)	5.98e-5 (0.48)	4.11e-5 (0.33)
ϕ	0.046 (1.65)	0.056 (1.81)	0.048 (1.66)	0.052 (1.81)	0.044 (1.53)	0.018 (0.60)	0.034 (1.20)	0.046 (1.54)	0.044 (1.44)	0.035 (1.17)
θ_1	0.934 (22.53)	0.934 (20.77)	0.945 (23.56)	0.965 (24.70)	0.936 (22.44)	0.957 (21.97)	0.931 (22.72)	0.950 (23.05)	0.934 (22.52)	0.911 (21.35)
θ_2	-0.066 (-2.93)	-0.066 (-2.32)	-0.071 (-3.16)	-0.069 (-3.01)	-0.064 (-2.84)	-0.037 (-1.58)	-0.065 (-2.88)	-0.059 (-2.57)	-0.071 (2.96)	-0.091 (-3.34)
μ_1	3.59e-3 (2.27)	-	-7.50e-4 (-3.62)	-	-3.46e-3 (-0.47)	-	-8.53e-3 (-4.30)	-	-2.76e-3 (-3.11)	-
μ_2	2.00e-3 (0.88)	-	-1.53e-3 (-1.13)	-	-2.59e-3 (-0.49)	-	-3.78e-3 (-0.70)	-	1.43e-3 (0.35)	-
$\sqrt{\omega}$	8.73e-4 (12.95)	7.58e-4 (7.20)	9.83e-4 (11.40)	1.11e-3 (10.84)	8.73e-4 (13.35)	9.47e-4 (12.26)	8.39e-4 (13.32)	9.46e-4 (12.58)	8.56e-4 (12.87)	4.36e-4 (4.43)
$\sqrt{\alpha}$	0.403 (14.36)	0.385 (12.20)	0.444 (13.57)	0.462 (12.49)	0.401 (15.84)	0.443 (14.79)	0.383 (15.52)	0.408 (13.37)	0.395 (15.47)	0.297 (17.11)
$\sqrt{\beta}$	0.895 (73.52)	0.905 (71.99)	0.872 (47.26)	0.841 (33.75)	0.897 (76.10)	0.881 (61.08)	0.905 (82.08)	0.887 (60.79)	0.900 (77.10)	0.950 (135.63)
$\sqrt{\lambda_1}$	-	4.88e-3 (7.74)	-	2.80e-3 (3.83)	-	2.20e-6 (0.00)	-	4.63e-3 (3.44)	-	3.25e-3 (4.26)
$\sqrt{\lambda_2}$	-	6.37e-7 (0.00)	-	1.23e-6 (0.00)	-	3.60e-5 (0.00)	-	-2.36e-6 (0.00)	-	-1.05e-3 (-0.30)

Repo = average repo rate; T-bills = average 3 month discount bills rate; CPI = Consumer Price Index; prel CPI = preliminary Consumer Price Index; PPI = Producer Price Index.
In parentheses t-statistics ($t_{753,\alpha=0.25}$=1.96).

Table 3

MULTIVARIATE GARCH(1,1)

		μ_{i0}	μ_{i1}	μ_{i2}	μ_{i3}	θ_{i1}	θ_{i2}	$\sqrt{\omega_i}$	$\sqrt{\alpha_i}$	$\sqrt{\beta_i}$	$\sqrt{\lambda_{i1}}$	$\sqrt{\lambda_{i2}}$	$\sqrt{\lambda_{i3}}$	ω_{12}	ω_{23}	ω_{13}	Mon	Tue	Wed	Fri	γ_1	γ_2	φ_1	φ_2
mark/dollar	a)	-9.62e-5	-	-	-	-	-	2.09e-3	0.266	0.913	-	-	-	-0.582	0.536	-0.236	-	-	-	-	-	-	-	-
		(-0.37)	-	-	-	-	-	(10.93)	(9.51)	(61.42)	-	-	-	(-19.33)	(19.51)	(-4.42)	-	-	-	-	-	-	-	-
lira/mark																								
baseline	a)	1.01e-4	0.093	0.290	0.775	0.805	-0.217	8.16e-4	0.398	0.912	0.018	0.091	0.671	-0.582	0.536	-0.236	-	-	-	-	-	-	-	-
		(0.69)	(4.59)	(5.87)	(3.89)	(19.48)	(-6.02)	(6.16)	(11.07)	(65.33)	(0.11)	(2.15)	(5.28)	(-19.33)	(19.51)	(-4.42)	-	-	-	-	-	-	-	-
regime	b)	1.66e-4	0.121	0.267	0.712	0.793	-0.248	5.96e-4	0.322	0.918	0.041	0.118	0.629	-0.594	0.530	-0.279	-	-	-	-	2.70	0.06	-	-
shifts		(1.28)	(7.86)	(6.09)	(3.53)	(20.54)	(-6.72)	(4.02)	(8.97)	(68.71)	(0.62)	(3.85)	(5.20)	(-20.54)	(19.58)	(-5.56)	-	-	-	-	(1.70)	(0.12)	-	-
seasonality	c)	4.37e-5	0.087	0.159	0.249	0.822	-0.233	9.56e-4	0.374	0.913	0.153	0.204	0.95	-0.605	0.570	-0.281	0.001	7.74e-4	3.14e-4	-3.76e-4	-	-	-	-
		(0.24)	(4.43)	(3.15)	(1.57)	(18.87)	(-6.54)	(2.40)	(10.76)	(61.73)	(3.51)	(3.95)	(9.36)	(-20.52)	(19.31)	(-4.85)	(2.37)	(1.09)	(0.15)	(-0.22)	-	-	-	-
jump of the	d)	8.45e-5	0.149	0.364	0.613	0.792	-0.225	9.23e-4	0.353	0.884	0.051	-0.016	0.646	-0.578	0.517	-0.220	-	-	-	-	-	-	0.178	-
persistence		(0.54)	(7.14)	(5.56)	(2.37)	(17.43)	(-6.12)	(7.15)	(10.14)	(60.82)	(0.56)	(-0.05)	(4.07)	(-19.21)	(18.60)	(-4.36)	-	-	-	-	-	-	(3.28)	-
BTP yield																								
baseline	a)	-3.49e-4	0.058	0.425	0.599	0.992	-0.114	1.93e-3	0.313	0.926	0.296	0.214	0.740	-0.582	0.536	-0.236	-	-	-	-	-	-	-	-
		(-0.88)	(1.39)	(2.27)	(0.66)	(8.38)	(-0.10)	(4.94)	(7.87)	(66.02)	(4.97)	(-1.97)	(3.66)	(-19.33)	(19.51)	(-4.42)	-	-	-	-	-	-	-	-
regime shifts	b)	-5.91e-4	0.033	0.393	0.343	0.869	-0.132	1.66e-3	0.273	0.933	0.166	0.089	0.349	-0.594	0.530	-0.279	-	-	-	-	0.584	1.149	-	-
		(-1.68)	(0.97)	(2.38)	(0.53)	(8.14)	(-1.23)	(4.99)	(7.68)	(64.00)	(1.78)	(0.41)	(1.42)	(-20.54)	(19.58)	(-5.56)	-	-	-	-	(1.33)	(2.03)	-	-
seasonality	c)	-4.98e-4	0.026	0.305	0.341	0.884	-0.184	1.38e-4	0.335	0.911	0.008	0.196	1.09	-0.605	0.570	-0.281	-0.002	-0.022	0.0045	0.0049	-	-	-	-
		(-0.97)	(0.63)	(1.22)	(0.32)	(6.73)	(-1.36)	(4.24)	(5.99)	(34.03)	(0.01)	(3.47)	(8.76)	(-20.52)	(19..31)	(-4.85)	(-0.98)	(-1.15)	(3.61)	(3.68)	-	-		
jump of the	d)	-5.42e-4	0.023	0.540	0.300	1.058	-0.011	1.77e-3	0.315	0.905	0.051	-0.016	0.646	-0.578	0.517	-0.220	-	-	-	-	-	-	-	0.093
persistence		(-1.50)	(0.47)	(3.17)	(0.38)	(9.69)	(-0.10)	(4.11)	(8.49)	(55.46)	(0.56)	(-0.04)	(4.08)	(-19.21)	(18.60)	(-4.36)	-	-	-	-	-	-	-	(1.66)
BTP futures																								
baseline	a)	-3.49e-4	-4.12e-3	-4.03e-2	-5.34e-2	-3.04e-2	-1.23e-2	1.38e-4	0.315	0.934	1.44e-2	1.73e-4	2.19e-2	-0.582	-0.480	0.302	-	-	-	-	-	-	-	-
		(-0.88)	(-2.19)	(-4.38)	(-1.17)	(-4.62)	(-2.68)	(5.14)	(6.17)	(47.95)	(3.36)	(0.00)	(0.67)	(-19.33)	(-16.41)	(7.68)	-	-	-	-	-	-	-	-
regime shifts	b)	3.29e-5	-1.71e-3	-4.14e-2	-5.14e-2	-3.34e-2	-1.68e-2	1.31e-4	0.257	0.927	1.24e-2	6.44e-3	3.10e-2	-0.582	-0.521	0.331	-	-	-	-	0.952	2.01	-	-
		(1.58)	(-1.20)	(-3.79)	(-1.10)	(-4.69)	(-2.36)	(3.93)	(4.11)	(31.19)	(1.50)	(0.31)	(1.03)	(-)	(-17.76)	(8.76)	-	-	-	-	(2.00)	(2.72)	-	-
seasonality	c)	-7.24e-5	-2.93e-3	-4.70e-2	-4.25e-2	-4.41e-2	-1.70e-2	2.68e-4	0.121	0.968	6.36e-3	-3.34e-2	-2.54e-2	-0.582	-0.594	0.352	1.47e-4	-1.35e-5	-1.59e-4	1.26e-4	-	-	-	-
		(-1.27)	(-0.96)	(-1.90)	(-0.29)	(-2.43)	(-1.01)	(3.74)	(3.78)	(47.41)	(0.10)	(-1.78)	(-0.25)	(-)	(-28.82)	(8.12)	(0.18)	(0.00)	(-0.17)	(0.14)	-	-	-	-
jump of the	d)	-8.44e-7	-1.17e-3	-0.035	-0.062	-4.06e-2	-1.75e-2	1.82e-4	0.304	0.925	-1.03e-2	1.01e-3	-1.64e-2	-0.582	-0.456	0.288	-	-	-	-	-	-	-	0.252
persistence		(-0.02)	(-0.44)	(-2.62)	(-0.90)	(-4.13)	(-1.90)	(3.95)	(3.96)	(31.41)	(-0.57)	(0.00)	(-0.20)	(-)	(-11.79)	(6.13)	-	-	-	-	-	-	-	(1.60)

In parentheses, *t* statistics. The parameters refer to equations (1)-(10) in the text and their modifications. The baseline model for the conditional means and covariances of the mark/dollar, the lira/mark and the BTP yield (or the BTP futures) is reported in lines a). In lines b) it is modified to account for regime shifts of the unconditional variance across the three governments; in lines c) the basline model takes into account the presence of weekday seasonality; in lines d) we test for the presence of changes in the conditional volatility and its persistence in days characterized by the occurrence of scheduled or unscheduled news.

Figure 1

TIME FRAMEWORK REPRESENTATION

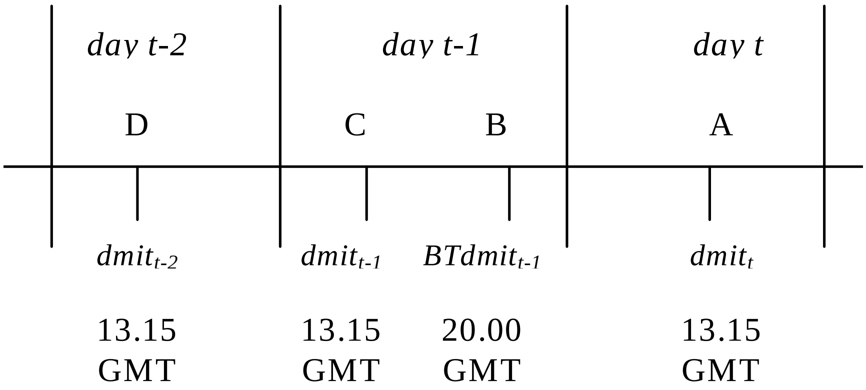


Figure 2

CONDITIONAL CORRELATIONS

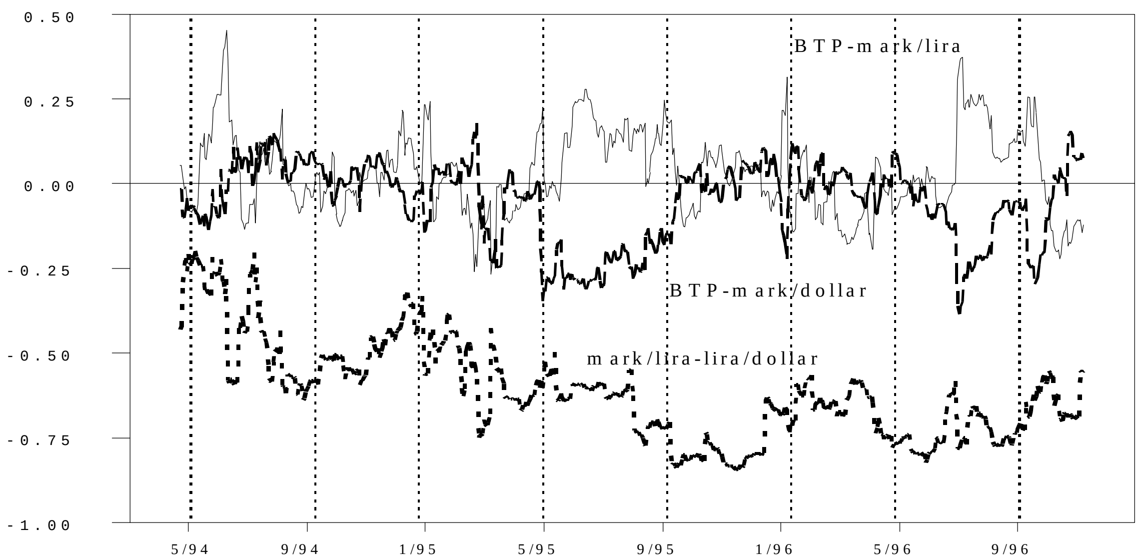


Figure 3

CONDITIONAL COVARIANCES

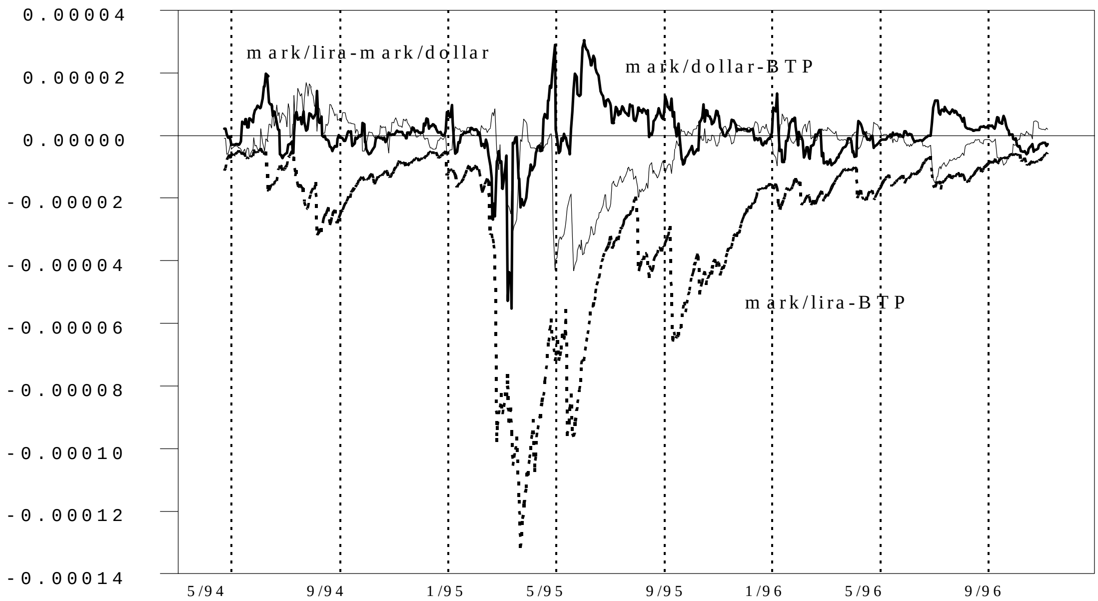


Figure 4

CONDITIONAL STANDARD DEVIATIONS
(percent value on a yearly basis)

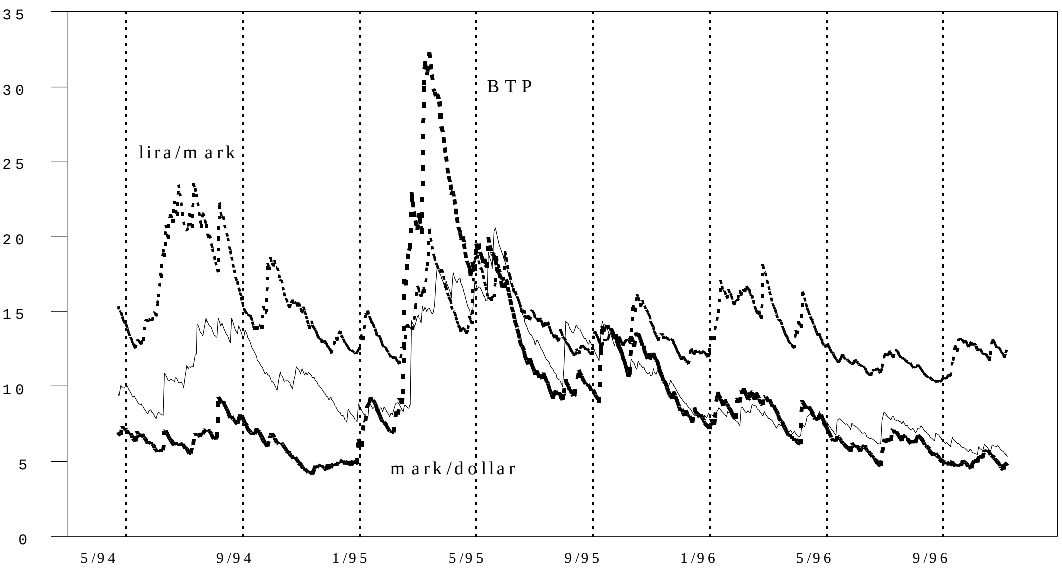


Figure 5

CONDITIONAL STANDARD DEVIATION OF LIRA/MARK EXCHANGE RATE

(overall and component due to scheduled and unscheduled news; percent values)

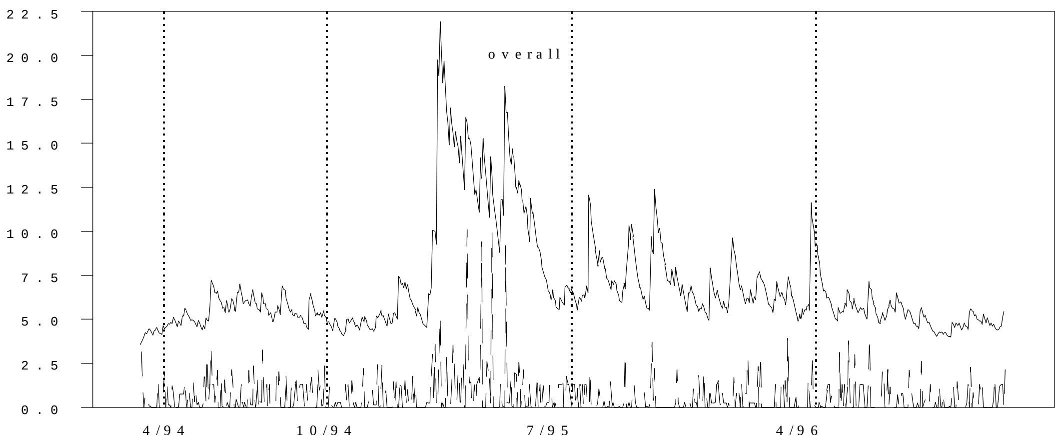


Figure 6

CONDITIONAL STANDARD DEVIATION OF THE BTP YIELD

(overall and component due to scheduled and unscheduled news; percent values)

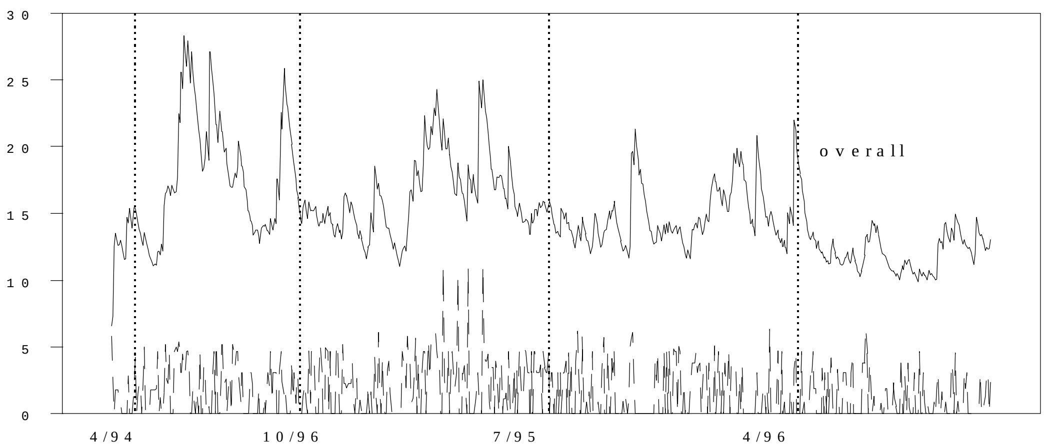


Figure 7

CONDITIONAL STANDARD DEVIATION OF THE BTP FUTURES

(overall and component due to scheduled and unscheduled news; percent values)

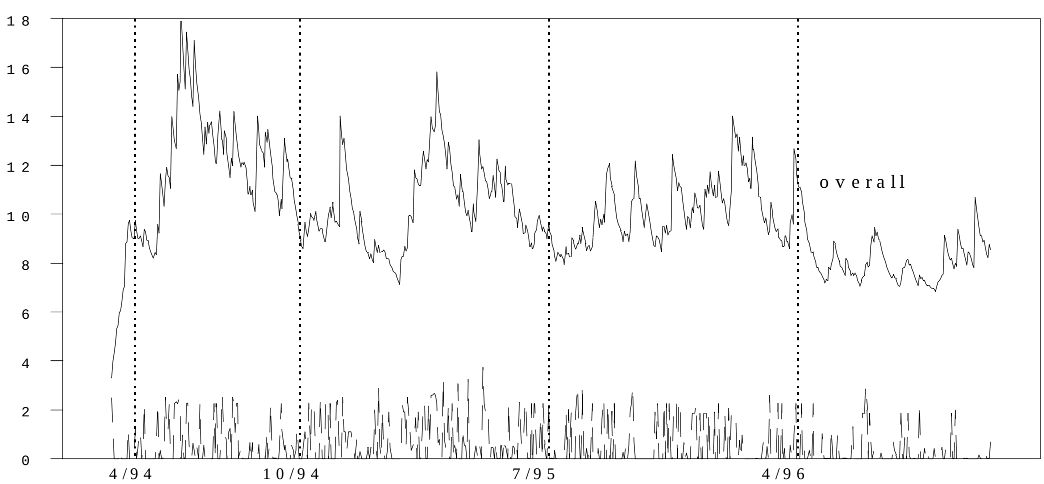


Figure 8

IMPULSE ANALYSIS OF THE MULTIVARIATE GARCH MODEL: UNSCHEDULED NEWS

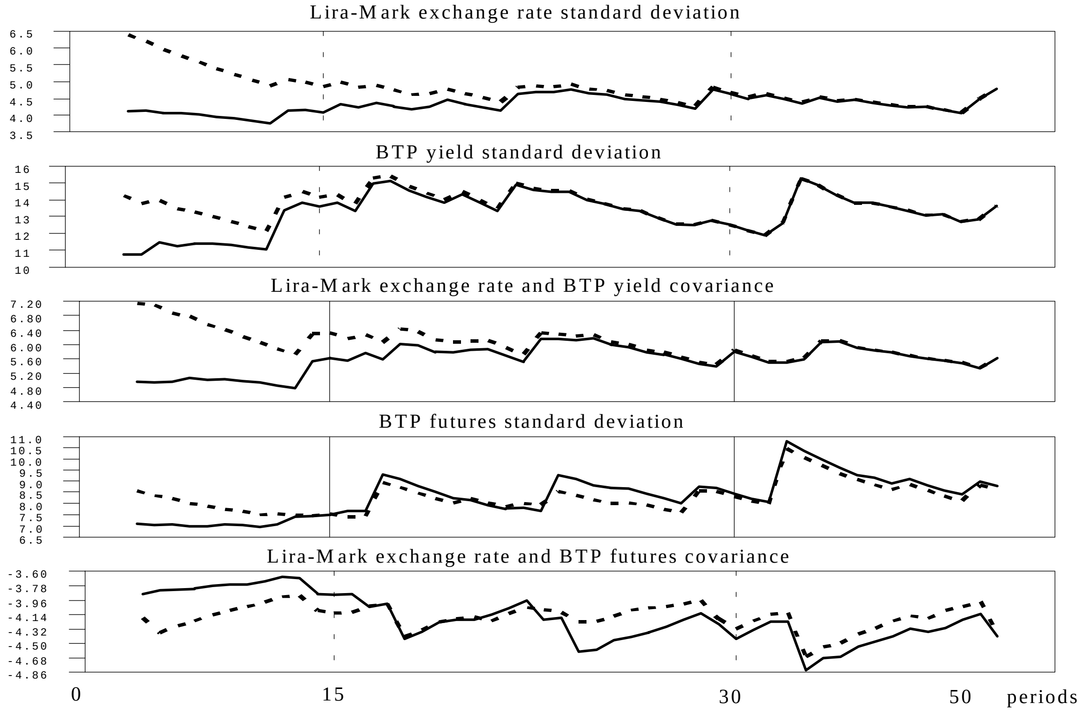
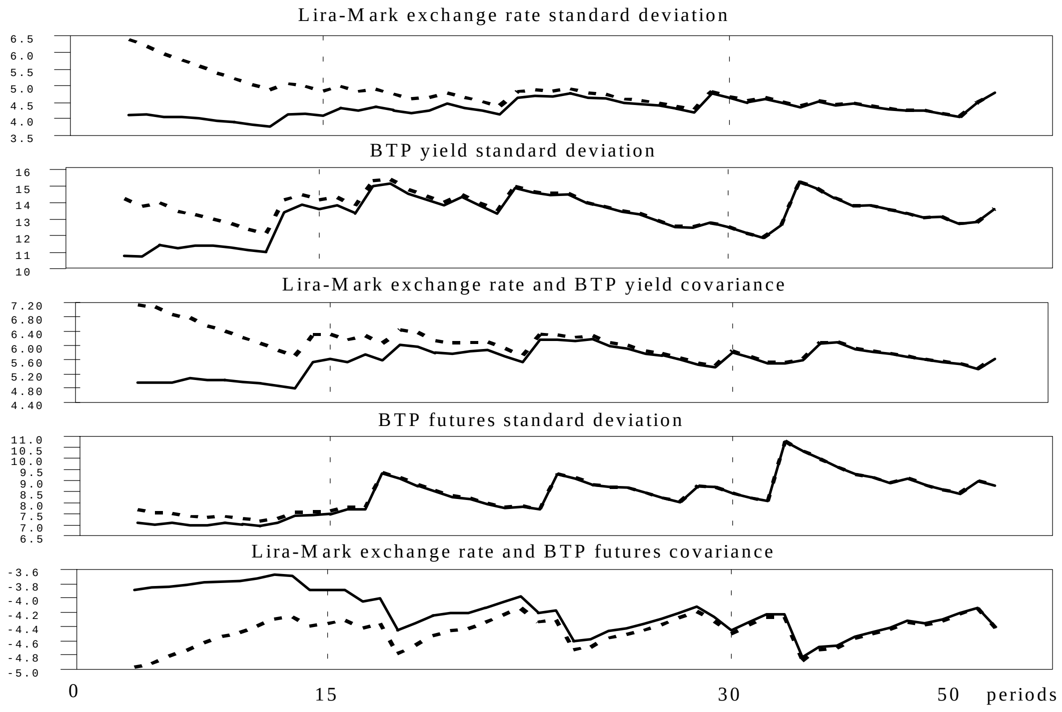


Figure 9

IMPULSE ANALYSIS OF THE MULTIVARIATE GARCH MODEL: SCHEDULED NEWS



Bold lines are observed values, dotted lines are simulated values.

Appendix

Setup of the dummies

As a reminder to the non-Italian reader we recall the main political events that occurred in Italy between 28 March 1994 and November 1996. In the wake of corruption scandals that led to the deepest Italian political crisis since World War II, general elections were held on 28 March 1994 resulting in the victory of a center-right coalition (*Polo delle Libertà*) led by Mr. Berlusconi. During our sample period, three Governments were in power (Berlusconi, Dini and Prodi, respectively) and there were two full-fledged government crises. The first occurred in December 1994-January 1995, when members of the ruling coalition, the Northern League, revoked their support to the government; the second was in January-February 1996, when the Government led by Mr. Dini resigned. The Parliamentary coalitions backing the three Governments were characterised by very different degrees of cohesion; this turned out to be a fundamental factor in the way the markets reacted to announcements of domestic policies and to international shocks. Moreover, the US dollar — usually regarded as an important contributing factor in exchange rate dynamics among European currencies — changed significantly during the time period: from broadly stable during the first government, to falling and then remaining weak during the second and strengthening during the third.

In this turbulent environment, a large number of news items could be envisaged. The construction of a dummy recording the occurrence of a news item implies some degree of arbitrariness since it requires a) selecting topics that can affect financial markets and b) calibrating, in the sense explained below, the news itself. Furthermore, since no news-sampling can be exhaustive in itself, a hopefully unbiased criterion has to be chosen in order to select those news items relevant to the lira/mark exchange rate as well as to Italian fixed-income assets.

In this paper unscheduled news, i.e. information whose time of release is uncertain and not embodied into observed prices, was chosen by the following method: i) the headlines concerning Italian political and economic events appearing in the first two pages of Il Sole-24 Ore, the leading Italian financial newspaper were collected; ii) the above selection was completed with news items reported in the Financial Times, both in the European section and in the Reuters news bank. The selection reported in i) and ii) rests heavily on the typology of news used in the literature (see footnote 2) and in the financial press, as well as on conversations held with asset managers and traders.

Each news items was then scaled according to the following approach: if it was expected to weaken the exchange rate we assigned it a unit value; if it was expected to strengthen the exchange rate we assigned it the value -1; nil means no news. Hence, good news about public finance issues is expected to strengthen the lira-mark exchange rate and the associated dummy receives a value of -1; if a declaration is expected to weaken the lira-mark exchange rate, the associated dummy receives a value of 1.

The following categories of unscheduled news were selected:

- Public Finance: all the news in this category describes events likely to affect the Government deficit directly (e.g. variations in interest rates) and disclosures about it.
- Consumer Prices: here, most news is scheduled but there are also unscheduled events (e.g. CPI and PPI releases, but also labour costs, raw material costs, indirect taxes news).
- Institutional Conflicts: this is a very Italian-specific kind of news, referring to the balance of political power between the President of the Republic, Parliament and the Premier in the newly established Italian bipartisan setup.
- Political Conflicts.
- Electoral results: though electoral results are scheduled, in that they are predetermined, there is generally a flow of information from declarations made by of political leaders, concerning the importance of the observed shares, that typically continues for two or three days after the vote; this is why they are regarded as unscheduled.
- Easing of Political Tensions during the Dini Government (January 16-December 29, 1995): refers to all political conflicts and easings of tensions in this period between the Government coalition and the opposition.
- Dini government: this contains a number of news events related to political action by the Dini government.
- Political debate from the beginning of 1996.
- EMU events directly relevant to Italy: this category refers mostly to European doubt about Italy’s entrance in the EMU.

- EMU and the international environment: this news refers to the international market's decoding of events leading to the EMU.
- Debate on judiciary procedures: this news refers to clashes between the Berlusconi government and the Pool of Milan magistrates who bore responsibility for legal suits against corrupted politicians and eventually against Berlusconi himself.
- Black Friday, 17 March 1995: this is a vector equal to one on Black Friday and -1 on the three subsequent working days.
- Justice Minister Mancuso's actions and his Parliamentary removal.

For *scheduled news*, we use the difference between the observed and the previous value of each variable, under the assumption that the latter proxy expectations (here we implicitly assume that markets have static expectations or that series follow random walks and agents have rational expectations). Scheduled news regards macroeconomic, financial and monetary data releases. We choose each of them according to the relative importance market agents assign to it. We select the average repurchase agreement (repo) rate and the three-month bill auction rate which generally are considered to have the largest impact on money market rates and, consequently, on the spot exchange rate market; also we consider in this category the preliminary and the final consumer price index as well as the final producer price index, which all affect market perception of a likely swing in the stance of monetary policy.

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