Modeling Short-Term Interest Rate Spreads in the Euro Money Market*

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In the framework of a new money-market econometric model, we assess the degree of precision achieved by the European Central Bank (ECB) in meeting its operational target for the short-term interest rate and the impact of the U.S. subprime credit crisis on the euro money market during the second half of 2007. This is done in two steps. Firstly, the long-term behavior of interest rates with one-week maturity is investigated by testing for cobreaking and for homogeneity of spreads against the minimum bid rate (MBR, the key policy rate). These tests capture the idea that successful steering of very short-term interest rates is inconsistent with the existence of more than one common trend driving the one-week interest rates and/or with nonstationarity of the spreads among interest rates of the same maturity (or measured against the MBR). Secondly, the impact of several shocks to the spreads (e.g., interest rate expectations, volumes of open-market operations, interest rate volatility, policy interventions, and credit risk) is assessed by jointly modeling their behavior. We show that after August 2007, euro-area commercial banks started paying a premium to participate in the ECB liquidity auctions. This puzzling

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phenomenon can be understood by the interplay between, on the one hand, adverse selection in the interbank market and, on the other hand, the broad range of collateral accepted by the ECB. We also show that after August 2007, the ECB steered the "risk-free" rate close to the policy rate, but has not fully offset the impact of the credit events on other money-market rates.

JEL Codes: C32, E43, E50, E58, G15.

1. Introduction

Like other central banks of major currency areas, the European Central Bank (ECB) implements its monetary policy stance by steering very short-term interest rates. Differently from the U.S. Federal Reserve, the ECB does not announce an explicit target for its operational implementation of the monetary policy stance in the euro area.[1] However, the ECB provides refinancing to the banking system every week through its main refinancing operations (MROs), which are executed via variable-rate tender procedures with a pre-announced minimum bid rate (MBR). The level of the MBR is decided by the Governing Council of the ECB at its monthly policy meeting and signals the monetary policy stance for the euro area. Thus, the MBR can be seen as an implicit target for the weekly average of the overnight interest rate. Moreover, the MBR has been set, since April 1999, at the midpoint of the interest rate corridor defined by the rates on the two standing facilities offered by the ECB (marginal lending and deposit facilities). Given that the latter rates have an overnight maturity, they set bounds for the overnight interest rate in the euro area (± 100 basis points around the MBR, herein referred to as the interest rate corridor). Still, given that the interest rate corridor is relatively wide, it is unclear how much short-term interest rate volatility the ECB is willing to accept.

[1]For details of the operational framework for monetary policy implementation, see European Central Bank (2006b). For further institutional details of the euro money market, see Ewerhart et al. (2007).

The issue of the efficiency and effectiveness of the implicit, rather than explicit, operational targeting of a short-term interest rate is interesting from the perspective of the evaluation of different frameworks for the implementation of monetary policy and communication policy of the central bank. For instance, is the precision in achieving the target related to it being explicitly announced by the central bank? Another issue is that overnight interbank loans are unsecured, whereas lending by the ECB (and also by the U.S. Federal Reserve) is provided against collateral, which may complicate the quantitative assessment of the precision with which the target is met, as credit and funding/liquidity risk may affect market rates in a time-varying manner. In fact, in the euro area, spreads between very short-term money-market rates and the MBR have varied over time, sometimes displaying higher volatility or even a short-term increasing trend, which led the ECB to enact a policy of providing liquidity above the so-called benchmark amount (see European Central Bank 2002, 2006a). Which factors triggered the widening/narrowing of spreads? Was liquidity policy effective in moving the spreads in the desired direction?

In order to answer the first question, we assess whether the ECB successfully met its implicit operational target for the short-term interest rate. The proposed approach relies on a new fractionally integrated factor vector autoregressive (FI-F-VAR) model, allowing for joint modeling of the common long- and medium-term dynamics of interest rates. The analysis is done in two steps. First, we investigate the source of nonstationarity in short-term money-market rates; second, we test for persistency in the deviations of these rates from the MBR. In the proposed approach, the long-term analysis concerns the source of nonstationarity in the series. Under successful monetary policy implementation, the MBR should determine long-term fluctuations in all short-term interest rates, thereby being the common trend driving the market interest rates. The investigation of the long-term multivariate interest rate structure relies on the use of cobreaking techniques considering the MBR as a deterministic step function. In fact, the MBR is a step function with jumps. From a medium- to longer-term perspective, these jumps are stochastic, as their size and timing are not known in advance: they depend on the business cycle and on the monetary policy "rule" followed by the ECB. However, from a very short-term perspective (e.g., Tuesday on Tuesday) the size and the timing of the jump in the MBR are known in advance, given that the policy announcement is made on a Thursday and the effective implementation is on the following Wednesday.[2] If the common-break process can be identified with the MBR, then, under homogeneity in the cobreaking relationships, the spreads of the various interest rates against the MBR should not be affected by any other source of deterministic nonstationarity. In the second step, the underlying hypothesis is that successful monetary policy implementation should generate stationary and mean-reverting spreads with the tightness of the control inversely related to the degree of persistency in the spreads. Moreover, spreads should be either $I(0)$ or fractionally integrated ($I(d)$, with $0 < d < 0.5$). Hence, the investigation of the medium-term multivariate structure of interest rate spreads relies on the use of (fractional) cointegration techniques. The identified cointegrating vectors have an economic interpretation, which is useful for interpreting the sources of disequilibrium in the money market. The second set of questions is answered by means of forecast-error-variance decomposition and impulse-response analysis where the impact of several shocks (e.g., market expectations, size of operations, overnight rate uncertainty, policy interventions, and credit risk) are tested, also shedding some light on the short- to medium-term determinants of the dynamics of the spreads.

Our main findings are twofold: First, we cannot reject that the MBR is the common trend driving short-term money-market rates, which supports the idea that implicit targeting is an effective communication tool for steering interest rates. Second, we find that spreads against the MBR are highly persistent, which shows that the steering of short-term rates by the ECB is not tight. This may be due (at least partially) to the inefficiency of the implicit targeting, which has to be traded off with other characteristics of the

[2] This is strictly true only after the March 2004 reform of the operational framework of the ECB. In fact, before March 2004 the MBR and the rates of the standing facilities were announced on Thursday and became effective the next day (Friday). This difference should not affect our main results. See European Central Bank (2003, 2005) for details and assessment of the reform.

operational framework.[3] Third, the increasing liquidity deficit in the euro area, expectations of increases in policy rates (or the slope of the yield curve), and volatility of the overnight interest rate put an upward pressure on the spreads; these are effectively counteracted by liquidity policy. Fourth, the long- and medium-term structures of the euro money market do not seem to have been affected by the U.S. subprime credit crisis. However, its short-term dynamics has been significantly disturbed: within the FI-F-VAR model, this can be interpreted as deviations from equilibrium resulting from credit-risk shocks. In this context, we show that after August 2007, the ECB successfully steered the "risk-free" rate close to the MBR, but has not fully offset the impact of the credit events on the other money-market rates.

The remainder of the paper is structured as follows. Section 2 presents the economics of money-market spreads in the euro area. Section 3 presents the econometric methodology. Section 4 reports the main empirical results on money-market spreads. The impact of shocks is presented in section 5. Section 6 concludes.

2. The Economics of Short-Term Interest Rate Spreads in the Euro Area

The *martingale hypothesis* has been the baseline reference for modeling the overnight interest rate in the euro area. It says that when monetary policy is implemented within an interest rate corridor with reserve requirements and averaging, the overnight interest rate on any day within the reserve maintenance period, on_t, should be equal to the rate expected to prevail on the last day of the reserve maintenance period (day T), given the current available information, $E_t(on_T)$. This result can be derived from the seminal work of Poole

[3] Other relevant characteristics are as follows. First, there is the low frequency of open-market operations; weekly operations are less costly than daily operations, namely if the objective of the central bank is to provide refinancing to a large number of banks. Second, there is the size of the reserve requirement; when it is large (above optimal working balances) and remunerated, liquidity shocks do not affect overnight rates on a day-to-day basis because the former are absorbed by variations in the current accounts of the commercial banks with the central bank (except on the last day of the reserve maintenance period). Given the above, and the length of the reserve maintenance period (about one month), short-term interest rates in the euro area are sticky.

(1968) and has been applied to the institutional context of the euro area by Pérez-Quirós and Mendizábal (2006) and Välimäki (2003) (see also Whitesell 2006). Moreover, the expected overnight rate for the last day of the reserve maintenance period should be equal to the probability-weighted average of the rates of the two standing facilities. If policy rates are not changed, the spread of the overnight interest rate against the policy rate can be expressed as follows:

$$\begin{aligned} on_t - MBR &= E_t(on_T) - MBR \\ &= [l \cdot P_t(ML_T)] + [d \cdot P_t(DF_T)] - MBR, \end{aligned} \tag{1}$$

where l is the marginal lending rate, d is the deposit facility rate, and $P_t(.)$ denotes the probability of marginal lending (ML) or recourse to the deposit facility (DF) on the last day of the maintenance period (T), conditional on information available at time t. Equation (1) says that if the reserve maintenance period ends with a liquidity shortage, banks must borrow (overnight) from the lending facility of the ECB at a penalty rate (l); $P_T(ML_T)$ is the probability of such an event. If the reserve maintenance period ends with a liquidity surplus, banks transfer the surplus (overnight) to the deposit facility of the ECB at a penalty rate (d); $P_T(DF_T)$ is the probability of such an event. The latter is equivalent to banks having excess reserves remunerated at a penalty rate.

With a symmetric interest rate corridor, $MBR = (l + d)/2$. If the central bank targets zero net recourse to standing facilities and has unbiased forecasts of aggregate liquidity needs, $P_t(ML_T) = P_t(DF_T) = 0.5$. Substituting the two in equation (1) gives

$$on_t - MBR = \frac{l + d}{2} - MBR = 0, \tag{2}$$

leading to the prediction that the central bank should meet its (implicit or explicit) target without any further action or communication device beyond the following: (i) the announcement of the interest rate corridor and of the MBR, and (ii) the liquidity policy. In practice, a positive spread may still exist due to the unsecured nature of the interbank market or market misperceptions about the liquidity policy of the central bank.

Switching from the overnight interest rate to one-week interest rates, the spread between the one-week EONIA swap rate and the policy rate, $w_t^{swap} - MBR$, is of particular interest. Consider the last week of the reserve maintenance period:

$$w_t^{swap} - MBR = \frac{1}{7}\sum_{s=1}^{7} E_t(on_{t+s}) - MBR = \delta^{swap} + \varepsilon_t^{swap}, \qquad (3)$$

where δ^{swap} is a constant and ε_t^{swap} is a stochastic component, with $E(\varepsilon_t^{swap}) = 0$. From equation (1), by the law of iterated expectations, $\sum_{s=1}^{7} E_t(on_{t+s}) = 7 \cdot E_t(on_T)$, where $T = t + 7$ is the last day in the reserve maintenance period. Substituting in equation (3) gives

$$E_t(on_T) - MBR = \delta^{swap} + \varepsilon_t^{swap}. \qquad (4)$$

The tightness of the ECB's control over the overnight interest rate can be assessed against the benchmark of perfect control, which implies $\delta^{swap} = 0$, and white-noise deviations from target, ε_t^{swap}, $E(\varepsilon_t^{swap}) = 0$, and $E(\varepsilon_i^{swap}, \varepsilon_j^{swap}) = 0$ $(i \neq j)$. In assessing monetary policy implementation, we focus on the persistence and the underlying factors of the deviations of the short-term money-market rate from the target rate. Volatility is only one of those factors.

For other one-week interest rates, w_t^i, we define

$$w_t^i - MBR = \frac{1}{7}\sum_{s=1}^{7} E_t(on_{t+s}) - MBR = \delta^i + \varepsilon_t^i, \qquad (5)$$

where $i = mar$, war, $depo$, and $repo$ (respectively, marginal and weighted average MRO rates, deposit rate, and repo rate); δ^i is a (small) term premium or discount reflecting the (slightly) longer maturity of the one-week interest rate or differences in credit risk and liquidity risk; and ε_t^i is a stochastic component, with $E(\varepsilon^i) = 0$. Weekly rates are set by market participants in a forward-looking manner. The horizon for expectations considered is the same for all market rates corresponding to the maturity and the frequency of the regular open-market operations of the ECB (MRO). The spreads, $w_t^i - MBR(i \neq swap)$, do not have a clear interpretation; however, the implied spreads among money-market rates have a meaningful economic content. The focus of our analysis is on three spreads that measure credit risk, bid shading, and the (relative) cost of (liquid) collateral.

Credit Risk. Credit risk is captured by the spread between the one-week interbank deposit rate (w_t^{depo}, unsecured lending) and the one-week interbank repo rate (w_t^{repo}, secured or collateralized lending):

$$\begin{aligned} w_t^{depo} - w_t^{repo} &= \left(w_t^{depo} - MBR\right) - \left(w_t^{repo} - MBR\right) \\ &= \delta^{cr} + \varepsilon_t^{cr}, \end{aligned} \tag{6}$$

where $\delta^{cr} = \delta^{depo} - \delta^{repo}$, the difference of the two spreads against the MBR, measures *credit risk*; $\varepsilon_t^{cr} = \varepsilon_t^{depo} - \varepsilon_t^{repo}$, $E(\varepsilon_t^{cr}) = 0$. When the assets of the banks are observed with certainty, δ^{cr} should be close to zero in line with both structural and intensity models of credit-risk pricing (Duffie and Singleton 2003, ch. 3). However, when the assets of the banks are imperfectly observed, namely due to accounting uncertainty, δ^{cr} can be positive even for short maturities (Duffie and Lando 2001; Duffie and Singleton 2003, ch. 5). Moreover, credit risk may vary over the (business) interest rate cycle, increasing when interest rates are increasing and, conversely, decreasing when interest rates are decreasing; these variations should be captured by the residual component, ε_t^{cr}.

Bid Shading. The spread between the one-week EONIA swap rate (w_t^{swap}) and the marginal MRO rate (w_t^{mar}) captures the discount of (marginal) bid rates relative to the unobserved true marginal valuations, the latter proxied by the one-week EONIA swap rate:

$$\begin{aligned} w_t^{swap} - w_t^{mar} &= \left(w_t^{swap} - MBR\right) - \left(w_t^{mar} - MBR\right) \\ &= \delta^{bs} + \varepsilon_t^{bs}, \end{aligned} \tag{7}$$

where $\delta^{bs} = \delta^{swap} - \delta^{mar}$, the difference of the two spreads against the MBR, measures *bid shading*; $\varepsilon_t^{bs} = \varepsilon_t^{swap} - \varepsilon_t^{mar}$, $E(\varepsilon_t^{bs}) = 0$. Bid shading may vary with uncertainty about tender outcomes (i.e., allotment share and tender rates). Unfortunately, no general predictions on the impact of rate uncertainty can be made based on existing multiunit, discriminatory pricing, auction theory. Nevertheless, Ewerhart, Cassola, and Valla (2006) show that for the special case of declining linear marginal valuations and symmetric bidders, in the discriminatory auction the marginal bid rate, the average bid rate, marginal valuations, and bid shading are all expected to increase with a simultaneous and proportional expansion in the liquidity needs of banks and the allotment volume. Hortaçsu (2002), also for some special cases, derives results suggesting that bid shading should increase with rate and liquidity uncertainty and decrease with the number of participants in the auction.

Relative Cost of Liquid Collateral. Consider the spread between the marginal MRO rate (w_t^{mar}) and the one-week repo rate (w_t^{repo}):

$$\begin{aligned} w_t^{mar} - w_t^{repo} &= \big(w_t^{mar} - MBR\big) - \big(w_t^{repo} - MBR\big) \\ &= \delta^{cc} + \varepsilon_t^{cc}, \end{aligned} \tag{8}$$

where $\delta^{cc} = \delta^{mar} - \delta^{repo}$, the difference of the two spreads against the MBR, measures the *relative cost of liquid collateral*; $\varepsilon_t^{cc} = \varepsilon_t^{mar} - \varepsilon_t^{repo}$, $E(\varepsilon_t^{cc}) = 0$. This spread should be positive, $\delta^{cc} > 0$, insofar as commercial banks pledge for central bank operations securities that are less liquid than those accepted for private repo transactions (see Ewerhart, Cassola, and Valla 2006).

3. Econometric Methodology

In our empirical analysis, we jointly model the dynamics of short-term interest rates in levels and test for a single common-break process, for stationarity of spreads against the policy rate (MBR), and for homogeneity of cointegrating relationships among money-market rates. We consider the following fractionally integrated factor vector autoregressive (FI-F-VAR) model (Morana 2007b):

$$x_t = \Lambda_\mu \mu_t + \Xi z_t + \Lambda_f f_t + C(L)(x_{t-1} - \Lambda_\mu \mu_{t-1}) + v_t \tag{9}$$

$$D(L) f_t = \eta_t, \tag{10}$$

where x_t is an n-variate vector of long-memory processes;[4] f_t is an r-variate vector of stationary long-memory factors ($I(d)$, $0 < d_i <$

[4]See Baillie (1996) for an introduction to long-memory processes.

0.5, $i = 1, \ldots, r$); μ_t is an m-variate vector of common deterministic break processes; v_t is an n-variate vector of zero-mean idiosyncratic i.i.d. shocks; η_t is an r-variate vector of common zero-mean i.i.d. shocks; $E[\eta_t v_{is}] = 0$ all i, t, s; Λ_f and Λ_μ are, respectively, $n \times r$ and $n \times m$ matrices of loadings; $C(L) = C_1 L + C_2 L^2 + \ldots$, is a finite-order matrix of polynomials in the lag operator with all the roots outside the unit circle; C_i $i = 1, \ldots$, is a square matrix of coefficients of order n; $D(L) = diag\{(1-L)^{d_1}, (1-L)^{d_2}, \ldots, (1-L)^{d_r}\}$ is a diagonal matrix in the polynomial operator of order r; z_t is a q-variate vector of exogenous/shock variables; and Ξ is an $n \times q$ matrix of parameters. The fractional-differencing parameters d_i, as well as the μ_t and f_t factors, are assumed to be known, although they need to be estimated. This is not going to affect the asymptotic properties of the estimator, since consistent estimation techniques are available for all the parameters and unobserved components.

The Fractional VAR Form. By taking into account the binomial expansion[5] in equation (10) and substituting into equation (9), the infinite-order vector autoregressive representation for the factors f_t and the series x_t can be written as

$$\begin{aligned} \begin{bmatrix} f_t \\ x_t - \Lambda_\mu \mu_t \end{bmatrix} &= \begin{bmatrix} 0 \\ \Xi \end{bmatrix} z_t + \begin{bmatrix} \Phi(L) & 0 \\ \Lambda_f \Phi(L) & C(L) \end{bmatrix} \\ &\quad \times \begin{bmatrix} f_{t-1} \\ x_{t-1} - \Lambda_\mu \mu_{t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_{f_t} \\ \varepsilon_{x_t} \end{bmatrix}, \end{aligned} \tag{11}$$

where $D(L) = I - \Phi(L)$, and $\Phi(L) = \Phi_0 L^0 + \Phi_1 L^1 + \Phi_2 L^2 + \ldots$, Φ_i, $\forall i$ is a square matrix of coefficients of dimension r,

$$\begin{bmatrix} \varepsilon_{f_t} \\ \varepsilon_{x_t} \end{bmatrix} = \begin{bmatrix} I \\ \Lambda_f \end{bmatrix} \eta_t + \begin{bmatrix} 0 \\ v_t \end{bmatrix},$$

with variance-covariance matrix

$$E\varepsilon_t \varepsilon_t' = \Sigma_\varepsilon = \begin{bmatrix} \Sigma_\eta{}' & \Sigma_\eta{}' \Lambda_f' \\ \Lambda_f \Sigma_\eta{}' & \Lambda_f \Sigma_\eta{}' \Lambda_f' + \Sigma_v \end{bmatrix},$$

where $E\eta_t \eta_t' = \Sigma_\eta$ and $E v_t v_t' = \Sigma_v$.

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[5] $(1-L)^d = \sum_{j=1}^{\infty} \rho_j L^j$, $\rho_j = \frac{\sum_{k=0}^{\infty} \Gamma(j-d) L^j}{\Gamma(j+1)\Gamma(-d)}$, where $\Gamma(\cdot)$ is the gamma function.

Estimation. The estimation problem may be written as follows:

$$\min_{\mu_1,\ldots,\mu_m,f_1,\ldots,f_r,\Lambda_f,\Lambda_\mu,C(L),} T^{-1}\sum_{t=1}^{T}\varepsilon'_{x_t}\varepsilon_{x_t},$$

where $\varepsilon_{x_t} = [I - C(L)L][x_t - \Lambda_\mu \mu_t] - [\Lambda_f \Phi(L)L]f_t - \Xi z_t$. Yet, since the infinite-order representation cannot be handled in estimation, a truncation to a suitable large lag for the polynomial matrix $\Phi(L)$ is required. Hence, $\Phi(L) = \sum_{j=0}^{p} \Phi_j L^j$.

The estimation problem can then be solved following an iterative process, described in Morana (2007b).

The above approach can be understood as a generalization of the factor VAR approach proposed by Stock and Watson (2005), allowing for both deterministic and long-memory stochastic factors. Stock and Watson (2005) provide details about the asymptotic properties—i.e., consistency and asymptotic normality—of the estimation procedure for the case of $I(0)$ variables. Although no theoretical results are currently available for long-memory processes, Monte Carlo evidence provided in Morana (2007a) fully supports the use of the principal component analysis (PCA) for long-memory processes.[6] Moreover, since the fractional-differencing parameter can be consistently estimated, the asymptotic properties of the estimation method are not affected by the conditioning to the initial estimate of the persistence parameter. In addition, the two-step iterated procedure is leading to maximum likelihood estimation of the model and therefore to full efficiency. Finally, the above model can also be estimated by relying on the surplus lag approach, requiring neither the computation of the fractional-differencing parameter nor the binomial expansion, as the $C(L)$ and $\Phi(L)$ matrices can be modeled as standard finite-order stationary polynomials in the lag operator, provided that appropriate accounting of the excess lag is carried out.

4. Data and Modeling Issues

The sample covered in the econometric analysis runs from June 27, 2000, until December 11, 2007. The last eighteen weeks in the sample

[6]Theoretical results also validate the use of PCA in the case of both weakly and strongly dependent processes. See, for instance, Bai (2003, 2004) and Bai and Ng (2004).

(5 percent of the sample) cover a particularly volatile period, after August 2007, related to the impact on the euro money market of the U.S. subprime credit crisis, herein referred to as the *turmoil* period.[7] The following one-week maturity interest rates series are included: the marginal MRO rate, w_t^{mar}; the weighted average MRO rate, w_t^{war}; the uncollateralized loan rate, w_t^{depo}; the collateralized loan rate, w_t^{repo}; and the EONIA swap rate, w_t^{swap}. The data is of weekly frequency, collected on the allotment day of the MRO of the ECB (Tuesday). The market rates were collected at 9:30 a.m. from selected brokers by the Front Office Division (ECB).

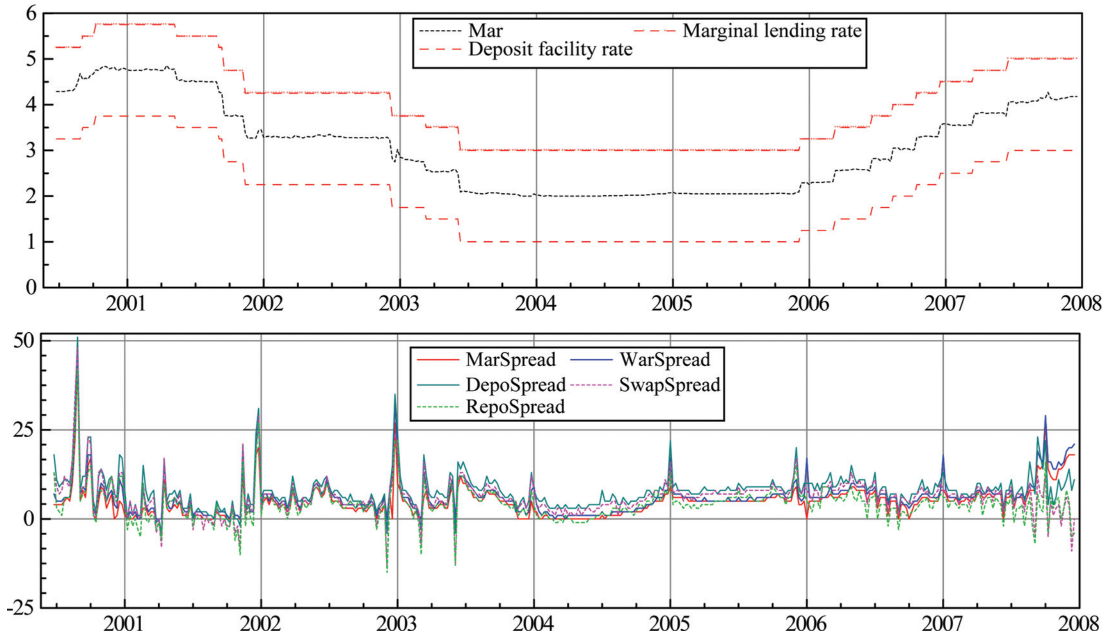
Figure 1 shows, in the upper panel, the time series of the marginal MRO rate and of the rates of the two standing facilities. The former moved smoothly and very close to the midpoint of the interest rate corridor defined by the latter. The lower panel of figure 1 shows the spreads of money-market rates against the MBR. Two facts are noteworthy: first, the relatively smooth behavior of the spreads except for a few spikes mainly around the end of the year; and second, the higher volatility and dispersion after August 2007. The (rounded) sample averages of the spreads, excluding the period between August 2007 and December 2007, are as follows: $w_t^{mar} - MBR = 4$ basis points; $w_t^{war} - MBR = 6$ basis points; $w_t^{depo} - MBR = 8$ basis points; $w_t^{repo} - MBR = 4$ basis points; and $w_t^{swap} - MBR = 6$ basis points.[8] The swap spread is small but different from zero, suggesting less-than-perfect control of the overnight interest rate by the ECB. Still, such a small deviation should not be considered as jeopardizing the monetary policy signaling function of the MBR.

Figure 2 plots the spreads among money-market rates. Credit risk (upper panel) shows some volatility around the end of the year (excluding the turmoil period, the sample average, $w_t^{depo} - w_t^{repo} =$

[7]See European Central Bank (2007) and Ferguson et al. (2007) for an early assessment of the U.S. subprime credit crisis.

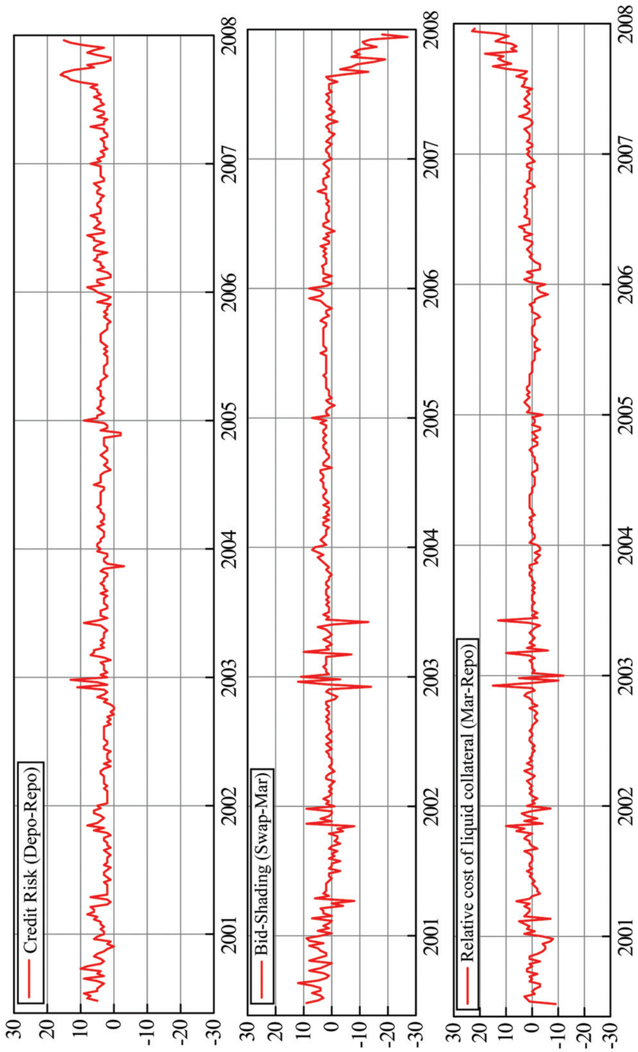
[8]As shown below, the spreads against the MBR follow a stationary long-memory process ($0 < d < 0.5$). Results of Beran (1994, ch. 8) indicate that the sample mean estimator is unbiased and efficient also in the case of long memory. In the Gaussian case the estimator is also the maximum likelihood estimator and therefore it is optimal relative to the class of both linear and nonlinear estimators. The sample mean estimator is also consistent, though the rate of convergence in the stationary long-memory case is slower than for the i.i.d. case: T^{-d} rather than $T^{-0.5}$.

Figure 1. Short-Term Money-Market Rates, Key ECB Policy Rates, and Spreads against the MRO Minimum Bid Rate



Note: Rates are in percent and spreads are in basis points.

Figure 2. Money-Market Rates Spreads (in basis points)



4 basis points). Bid shading (middle panel) is small and shows high volatility around the end of the year (excluding the turmoil period, the sample average, $w_t^{swap} - w_t^{mar} = 2$ basis points). After August 2007, bid shading became negative, which is somewhat puzzling. The relative cost of liquid collateral (lower panel) hovered around zero for most of the sample period, suggesting that scarcity of collateral has not been an issue in the euro area (excluding the turmoil period, the sample average, $w_t^{mar} - w_t^{repo} = 0$). This can be explained by the large pool of collateral accepted by the ECB in its operations.[9] However, during the turmoil period, the relative cost of liquid collateral increased sharply. In fact, the evolution of bid shading and of the costs of liquid collateral, after August 2007, are the two sides of the impact of the turmoil in the euro money market. These developments will be discussed further below.

4.1 Cobreaking Properties and Long-Term Linkages

The multivariate interest rate structure is investigated for cobreaking using the tests of Bierens (2000), with critical values computed by simulation, in order to allow for potential long memory in the break-free series.[10] Two different analyses have been carried out. The first analysis assesses cobreaking between each interest rate series and the minimum bid rate. The second analysis assesses cobreaking among the five money-market interest rate series, jointly considered, with the minimum bid rate. In both cases the cobreaking space has been estimated both unrestrictedly and under the homogeneity restrictions. Moreover, in order to assess the impact of the recent market turmoil, the analysis has been carried out considering two samples. The first sample ends on August 7, 2007, i.e., just before the beginning of the turmoil period (*pre-turmoil* sample, 372 observations); the second sample includes also data since August 14, 2007 (*full* sample, 390 observations). The results of the tests are reported in table 1, panels A and B. As shown in the table, for the pre-turmoil period, both the bivariate and the multivariate analyses point to a single nonlinear deterministic process driving the six

[9] For details on the collateral framework, see European Central Bank (2006b).

[10] See Hendry (1996) for the seminal work on cobreaking and Hendry and Massmann (2007) for a recent survey.

Table 1. Persistence and Copersistence Analysis

A. Bivariate Cobreaking Analysis

Pre-Turmoil Sample

#*cbr*	*mar*	*war*	*depo*	*swap*	*repo*	20%	10%	5%
1	0.213	0.179	0.155	0.202	0.087	0.489	0.665	0.803
2	0.951	0.942	0.935	0.935	0.938	0.698	0.876	1.033
1(H)	0.224	0.179	0.155	0.210	0.087	0.489	0.665	0.803

Full Sample

#*cbr*	*mar*	*war*	*depo*	*swap*	*repo*	20%	10%	5%
1	0.521	0.508	0.241	0.134	0.068	0.489	0.665	0.803
2	0.963	0.924	0.766	0.766	0.774	0.698	0.876	1.033
1(H)	0.660	0.594	0.242	0.134	0.070	0.489	0.665	0.803

B. Joint Cobreaking Analysis

Pre-Turmoil Sample

#*cbr*	U	H	20%	10%	5%
1	0.023	0.024	0.489	0.665	0.803
2	0.040	0.043	0.698	0.876	1.033
3	0.055	0.150	0.947	1.075	1.214
4	0.209	0.217	1.084	1.226	1.364
5	0.497	0.937	1.211	1.358	1.466
6	1.401		1.278	1.428	1.515

Full Sample

#*cbr*	U	H	20%	10%	5%
1	0.038	0.041	0.489	0.665	0.803
2	0.049	0.057	0.698	0.876	1.033
3	0.062	0.133	0.947	1.075	1.214
4	0.224	0.248	1.084	1.226	1.364
5	0.585	1.384	1.211	1.358	1.466
6	1.893		1.278	1.428	1.515

Notes: Panel A reports the value of the Bierens (2000) cobreaking test for each market interest rate (*mar*, *war*, *depo*, *repo*, *swap*) with the minimum bid rate, with (H) and without imposing the homogeneity restriction on the cobreaking space. #*cbr* denotes the number of cobreaking relationships, while 20%, 10%, and 5% are the corresponding critical values, computed by simulation, in order to account for long memory in the candidate break-free series. The null of the test is that the number of cobreaking relationships is less than or equal to n, $n = 1, 2, \ldots, 6$. Similarly, panel B reports the value of the joint Bierens (2000) cobreaking test for all the market interest rates with the minimum bid rate, with (H) and without imposing the homogeneity restriction on the cobreaking space.

(continued)

Table 1. (Continued)

C. Fractional-Differencing Parameter Estimation

Moulines and Soulier (1999) Broadband Log-Periodogram Estimater

Pre-Turmoil Sample

mar	0.311 (0.047)	*war*	0.320 (0.047)	*depo*	0.272 (0.047)	*swap*	0.308 (0.047)	*repo*	0.275 (0.047)

Full Sample

mar	0.295 (0.046)	*war*	0.338 (0.046)	*depo*	0.268 (0.046)	*swap*	0.285 (0.046)	*repo*	0.266 (0.046)

***ARFIMA*(4,*d*,0)**

Pre-Turmoil Sample

mar	0.246 (0.042)	*war*	0.287 (0.044)	*depo*	0.270 (0.044)	*swap*	0.226 (0.043)	*repo*	0.234 (0.043)

Post-Turmoil Sample

mar	0.316 (0.039)	*war*	0.352 (0.040)	*depo*	0.272 (0.042)	*swap*	0.223 (0.042)	*repo*	0.223 (0.043)

D. Fractional Cointegration Analysis (Robinson and Yajima 2002)

Pre-Turmoil Sample

$\#cr$	1	2	3	4	t_1	t_2	t_3	t_4	PV
1%	0.000	0.007	0.026	0.076	0.040	0.080	0.120	0.160	0.952
5%	0.000	0.006	0.022	0.064					
10%	0.000	0.005	0.020	0.058					
20%	0.000	0.004	0.018	0.050					

Full Sample

$\#cr$	1	2	3	4	t_1	t_2	t_3	t_4	PV
1%	0.002	0.008	0.022	0.120	0.040	0.080	0.120	0.160	0.942
5%	0.002	0.007	0.019	0.101					
10%	0.001	0.006	0.020	0.099					
20%	0.001	0.005	0.015	0.080					

Notes: In panel C the estimated fractional-differencing parameters obtained using the Moulines and Soulier (1999) broadband log-periodogram estimator and the ARFIMA(4,d,0) model are reported, with standard errors in parentheses. In panel D the results of the Robinson and Yajima (2002) fractional cointegrating rank test, at the 1%, 5%, 10%, and 20% significance levels, are reported for the selected bandwidth, i.e., four periodogram ordinates. $\#cr$ denotes the number of cointegrating relationships. The null of the test is that the number of cointegrating relationships is less than or equal to n, $n = 1, 2, \ldots, 4$, and t_n denotes the corresponding threshold values. Finally, PV denotes the proportion of total variance explained by the largest eigenvalue. The pre-turmoil period refers to the period June 27, 2000, through August 7, 2007, for a total of 372 observations, while the full sample period refers to the period June 27, 2000, through December 11, 2007, for a total of 390 observations.

series investigated. Concerning the bivariate case, the null of up to a single cobreaking vector (or a single common-break process) is not rejected at the 5 percent significance level, while the null of up to two cobreaking vectors (or structural stability) can be rejected at the 10 percent significance level. Similarly, for the multivariate case, the null of up to five cobreaking vectors (or a single common-break process) driving the six series is not rejected at the 5 percent level, while the null of up to six cobreaking vectors (or structural stability) can be rejected at a significance level slightly higher than 10 percent. In light of the dependence of the money-market interest rates on the MBR, then the single nonlinear deterministic trend detected in the data can be directly associated with the latter series. Since the homogeneity restriction is never rejected at the 5 percent significance level, given the identifying exclusion and normalization restrictions imposed, the cobreaking vectors can be interpreted in terms of spreads against the minimum bid rate in both periods and are irreducible, i.e., of the minimum possible dimension. If the MBR is the only source of structural change in the interest rate series, in light of the above results, the spreads from the minimum bid rate should be purely stochastic. The latter property is actually confirmed by the Dolado, Gonzalo, and Mayoral (2004) structural-break test (DGM test), modified to account for a general and unknown structural-break process (Morana 2007b), pointing to only residual long memory in the break-free series. Similar findings hold for the full sample, suggesting that, conditional on the evidence available so far, the market turmoil appears not to have affected the long-term structure of the euro-area money market.[11]

These results can be used to (partially) answer the first question: the announcement of an MBR is effective for steering short-term money-market interest rates. Moreover, our results suggest that no structural break occurred in the longer-term structure of the money-market spreads since 2000.

[11]For the *pre-turmoil* sample, the p-values of the DGM test, carried out on the spreads against the MBR, are 0.750 for w^{mar}, 0.600 for w^{war}, 0.325 for w^{depo}, 0.530 for w^{swap}, and 0.270 for w^{repo}. On the other hand, figures for the *full* sample are 0.565 for w^{mar}, 0.365 for w^{war}, 0.360 for w^{depo}, 0.495 for w^{swap}, and 0.290 for w^{repo}.

4.2 Persistence Properties of Interest Rate Spreads

The persistence properties of the interest rate spreads are investigated by means of both semiparametric and parametric estimators of the fractional-differencing parameter. Since the data are not affected by observational noise, the broadband log-periodogram estimator of Moulines and Soulier (1999) has been employed. Relative to other semiparametric estimators, it has the advantage of avoiding bandwidth selection problems, being also asymptotically efficient. Moreover, ARFIMA(p,d,q) models also have been fitted to the spreads.[12] The estimates are reported in table 1, panel C, for both the pre-turmoil and full samples. As shown in the table, the degree of long memory is estimated with precision in all cases, pointing to a similar degree of persistence for all spreads. In fact, for the pre-turmoil sample, the estimated fractional-differencing parameter is in the range 0.27 through 0.32 when the broadband log-periodogram estimator is employed and in the range 0.23 through 0.29 when the ARFIMA model is employed. Since in none of the cases the null hypothesis of equality of the fractional-differencing parameter can be rejected at the 5 percent level, the average value of the estimated fractional-differencing parameter has been employed as a common estimate, yielding an overall value of 0.275 (0.045). Moreover, the fractional cointegrating rank test of Robinson and Yajima (2002) points to up to four cointegrating relationships, at the 5 percent significance level, relating the five interest rate spreads to a single common long-memory factor driving the five processes. At the selected bandwidth (four periodogram ordinates), the largest eigenvalue of the spectral matrix at the zero frequency accounts for about 95 percent of total variance, supporting the conclusion drawn in favor of a single common long-memory factor. Similar findings for both the persistence and copersistence analysis hold for the full period, only pointing to a nonsignificant increase in the degree of average persistence, since the estimated fractional-differencing parameter increases to 0.284 (0.044), while the evidence in favor of cointegration is slightly weaker than for the pre-turmoil sample, as the test values are closer to the critical values for the null of cointegration. However, the impact on

[12] An ARFIMA(4,d,0) was selected for all the spreads series on the basis of misspecification criteria. The results are available upon request from the authors.

the proportion of explained total variance is negligible, since for the full sample the largest eigenvalue of the spectral matrix accounts for about 94 percent of total variance.

Concerning the structure of the cointegration space, estimation has been carried out by means of the semiparametric approach of Beltratti and Morana (2006) and Morana (2004a, 2005). The estimated unrestricted vectors are reported in table 2, panel A. Since a single common factor drives the five series, there are four cointegrating vectors, and identification requires a bivariate structure in all cases. As suggested by the theoretical analysis expressed in equation (6) and equation (8), the cointegrating vectors are expressed relative to w_t^{repo}. As the $w_t^{swap} - w_t^{repo}$ spread can be written as the sum of bid shading with the relative cost of collateral, the bid-shading component can be retrieved from the identified cointegrating vectors.

The restricted estimates are reported in table 2, panel A. As shown in the table, in all cases near-homogeneous cointegrating vectors have been found for the pre-turmoil sample. Moreover, taking into account the estimated standard errors, the null of homogeneity cannot be rejected in all cases. The restricted structure is strongly supported by the zero-frequency squared-coherence analysis as well, showing values of the statistic very close to the reference unity value, both in the unrestricted and restricted cases.[13] Finally, support for the proposed identification scheme is also provided by the outcome of the correlation analysis carried out on the restricted and unrestricted factors, pointing to a virtually unitary correlation coefficient between the common factor estimated on the basis of the unrestricted cointegration space and the one obtained on the basis of the restricted cointegration space. The findings are robust to the inclusion of the turmoil data as well, although some differences can be noticed. While the null of homogeneity of the cointegration space cannot be rejected on the basis of the estimated standard errors, a reduction of the efficiency of the estimates, as revealed by the much larger estimated standard errors, can be noticed. Apart from the

[13]The existence of cointegration between $I(d)$ bivariate processes implies that the squared coherence at the zero frequency of the series in differences is equal to 1, while when more than two processes are involved, it is the multiple squared coherence to assume a unitary value. See Morana (2004b).

Table 2. Cointegration Analysis

A. Estimated Cointegrating Vectors									
Pre-Turmoil Sample									
	Unrestricted					**Restricted**			
	CV_1	CV_2	CV_3	CV_4		RCV_1	RCV_2	RCV_3	RCV_4
mar	1.000	0.875	−0.487	0.621	*mar*	1	—	—	—
war	1.034	1.000	0.591	−0.744	*war*	—	1	—	—
depo	−1.153	1.184	1.000	1.314	*depo*	—	—	1	—
swap	0.798	−0.809	0.713	1.000	*swap*	—	—	—	1
repo	0.308	−0.246	0.187	−0.191	*repo*	−0.936	−0.976	−1.081	−1.085
						(0.097)	(0.102)	(0.096)	(0.097)
C^2	0.990	0.993	0.997	0.994	C^2	0.974	0.894	0.985	0.902
Full Sample									
	Unrestricted					**Restricted**			
	CV_1	CV_2	CV_3	CV_4		RCV_1	RCV_2	RCV_3	RCV_4
mar	1.000	0.752	−0.243	0.062	*mar*	1	—	—	—
war	1.031	1.000	0.743	−0.561	*war*	—	1	—	—
depo	−0.267	0.597	1.000	0.999	*depo*	—	—	1	—
swap	0.038	−0.249	0.552	1.000	*swap*	—	—	—	1
repo	0.160	−0.084	−0.019	0.435	*repo*	−1.268	−1.224	−1.212	−0.967
						(0.267)	(0.298)	(0.184)	(0.100)
C^2	0.990	0.994	0.987	0.956	C^2	0.599	0.578	0.755	0.893

(continued)

Table 2. (Continued)

B. Causality Analysis: Bayesian Information Criterion						
			Pre-Turmoil Sample			
Predicted Variable	**Predictive Lagged Variables**					
factor BIC	*mar* 2.538	*war* 2.524	*depo* 2.524	*swap* 2.535	*repo* 2.567	*factor* 2.526
			Post-Turmoil Sample			
Predicted Variable	**Predictive Lagged Variables**					
factor BIC	*mar* 2.365	*war* 2.354	*depo* 2.373	*swap* 2.431	*repo* 2.449	*factor* 2.367

Notes: Panel A reports the estimated unrestricted and restricted cointegrating vectors, with bootstrap standard errors in parentheses. C^2 denotes the squared multiple coherence at the zero frequency. Panel B reports the Bayesian information criterion (BIC) for the predictive equations for the estimated common factor. The pre-turmoil period refers to the period June 27, 2000, through August 7, 2007, for a total of 372 observations, while the full sample period refers to the period June 27, 2000, through December 11, 2007, for a total of 390 observations.

fourth cointegrating vector, in all the other cases there is an increase in the absolute magnitude of the cointegrating parameter. While the increase does not seem to be statistically significant, numerically it is not negligible, i.e., in the range 20–30 percent. The increase in the cointegrating parameter is related to the peculiar behavior of the MRO rates (w_t^{mar} and w_t^{war}), as mean reversion for these rates seems to take longer than for other market rates—a feature that is, however, not inconsistent with the long-memory property detected in the spreads.

An additional difference between the pre-turmoil and full samples can also be noticed in the estimated squared coherence for the restricted case. Yet it would not be correct to argue against fractional cointegration on the basis of the non-negligible reduction in the statistics associated with the exclusion restrictions. For instance, if the bivariate structure was expressed as $w_t^{war} - w_t^{mar}$, $w_t^{depo} - w_t^{war}$, $w_t^{depo} - w_t^{swap}$, $w_t^{swap} - w_t^{repo}$, rather than as spreads against w_t^{repo}, for all rates, figures for the squared coherence would be 0.987, 0.921, 0.794, and 0.893, respectively. Yet a spread structure against w_t^{repo} can be derived from the alternative specification by substitution which is exactly the same as the one estimated directly. Additional supporting evidence can be found by comparing the estimated common long-memory factor in the two cases: the processes are virtually indistinguishable, showing a correlation coefficient close to 0.94, with mean spread equal to -0.03 and 0.08 standard deviation.

Overall, it can then be concluded that, so far, the turmoil has not led to significant changes in the medium-term structure of the euro-area money market. Still, there is some perturbation in the short-term dynamics that will be discussed in the next section. These results can be used to complete the answer to the first question: the persistence found in the spreads suggests less-than-perfect control of short-term interest rates by the ECB (e.g., ε_t^{swap} is not a white-noise process).

4.3 Interpretation of the Factor

Concerning the interpretation of the factor, a Granger causality analysis has been carried out in order to assess whether a stronger predictive power for the common factor could be singled out across the interest rates spreads. In order to control for multicollinearity, the estimated factor has been regressed on its own lags and on the lagged values of each spread, one at a time. As is shown in table 2, panel B, the comparisons of the Bayesian information criteria allows for singling out w_t^{mar} as the only rate whose predictive power is always stronger than of the factor itself. The same result applies to the full sample. These findings suggest that bidding behavior and tender results are the most important determinant of the dynamics of the spreads. These results support the economic interpretation suggested by equations (7) and (8).

Our results highlight the role of the two steps in the implementation of the monetary policy stance by the ECB: the first step sets the level of market rates and consists of the announcement of the MBR (and of the interest rate corridor); the second step consists of steering the spreads against the MBR, by conducting weekly refinancing operations whose results, ultimately, depend on the allotment policy of the ECB.

5. The FI-F-VAR Model

In light of the results of the persistence and copersistence analysis, pointing to a single-break process and a single common long-memory factor, the dimension of the FI-F-VAR model is six equations, corresponding to the five money-market interest rates plus the single common long-memory factor. Given that the common-break process is known—minimum bid rate—the estimation of the model can be performed following a simplified strategy, requiring the iterative procedure only for the joint estimation of the common long-memory factor and the short-term dynamics. To capture the potential economic determinants of credit risk, bidding behavior, and the cost of collateral, five weakly exogenous variables have been included in the specification. The following weakly exogenous variables (z_t) were included:[14]

- F_t is the first difference in the spread between the one-month EONIA swap rate three months forward against the minimum

[14]For all five variables, standard tests do not reject the null of weak exogeneity at the five percent significance level. The p-value of the test ranges between 0.18 and 0.98. Detailed results are available upon request from the authors.

bid rate. This variable is a proxy for changes in interest rate expectations and/or in the slope of the money-market yield curve. Given that it refers to a longer forecast horizon (three months), it should capture the pure effect of expectations beyond the effect of the forward transmission of the very short-term movements in the one-week EONIA swap rate spread. An increase in this variable is expected to put an upward pressure on spreads through the credit-risk component. Before the reform of March 2004, ECB policy rates were implemented within the reserve maintenance period; thus, interest rate expectations also affected the spreads, as they changed the expected levels of the MBR and of the interest rate corridor for the (crucial) last day of the maintenance period, relative to other days. After the March 2004 reform, the impact of interest rate expectations on the spreads should have been much more muted, given that under the new arrangements policy rate changes are implemented only at the start of the reserve maintenance period.[15]

- A_t is the first difference of the residual of an OLS regression of the logarithm of allotment volumes on a constant and a linear trend. Detrending and first differencing is needed because the (log) allotment volume is a nonstationary variable. This variable compares weekly rates of growth in allotment volumes with the trend growth rate. Econometric results are robust to different detrending techniques. This variable captures the effect of shocks to the allotment volumes, and an acceleration of its growth path is expected to put an upward pressure on all spreads against the minimum bid rate (see Ewerhart, Cassola, and Valla 2006, and Välimäki 2006).
- E_t is the squared first difference (Tuesday on Tuesday) in the overnight interest rate (not swap rate). This is a measure of short-term interest rate volatility capturing end-of-maintenance-period conditions and some seasonal factors

[15]In practice, our results should not be affected by these changes because the first half of the sample is dominated by policy rate cuts, and the minimum bid rate put an effective floor on the downward adjustment of tender rates within the maintenance period. This is clearly not the case when policy rates are expected to increase. Nevertheless, testing for a March 2004 structural break in the dynamics of the spreads seems justified.

(end-of-year). It performs better than direct measures of liquidity imbalances as a measure of rate and liquidity uncertainty. An increase in this variable is expected to put an upward pressure on all spreads against the minimum bid rate.

- P_{t-1} is the lagged deviation from benchmark allotment. This variable measures the deliberate, policy-induced deviation from a smooth accumulation of reserve requirements over the reserve maintenance period. An increase in this variable, particularly if persistent and sizable, forces banks to accumulate reserves "too early," thus raising the risk of an early fulfillment of the reserve requirement, which would force banks to park the surpluses at the deposit facility of the ECB, which is costly (100 basis points below the MBR). Thus, an increase in this variable puts downward pressure on all spreads against the MBR.
- CR_t is the iTraxx Financials, which is a credit-risk measure (see Blanco, Brennan, and Marsh 2005); it measures the credit default swap premium on a basket of major European financial firms. It refers to senior debt and has a maturity of five years. Thus, it can be interpreted as the compensation that market participants require in order to bear the default risk of a set of financial firms. Given that it is a market price, it contains the statistical expectation of default risk as well as a risk premium, which is affected by changes in traders' risk appetite. An increase in this variable is expected to put an upward pressure on the spreads that price credit risk.

Coherent with the results of the fractional cointegration analysis, the findings for the pre-turmoil sample point to a single principal component explaining about 94 percent of total variance, as well as between 91 percent and 97 percent of the variance of each of the various interest rate series. Interestingly, the final estimate of the common long-memory factor obtained by means of principal components analysis and the one obtained on the basis of the Gauss/Kasa decomposition, following the approach of Beltratti and Morana (2006) and Morana (2004a, 2005), are strongly correlated (the correlation coefficient is about 0.91) and numerically very close, as the spread between the two factors has zero mean and standard deviation equal to 0.10. A similar finding holds for the full sample, albeit with slightly smaller figures. In fact, the first principal component of the variance-covariance matrix of the spreads explains about 88 percent of total variance, as well as between 84 percent and 96 percent of the variance of each of the various interest rate series. The comparison with the long-memory factor estimated by means of the Gauss/Kasa decomposition also shows a slightly larger deviation, with zero-mean spread, but standard deviation equal to 0.24. Overall, the findings provide further support to the proposed estimation methodology.

Thick estimation (Granger and Jeon 2004) of the FI-F-VAR model has been implemented by allowing up to four lags in the short-memory autoregressive specification ($C(L)$) and twenty-five lags in the long-memory autoregressive specification ($\Phi(L)$),[16] setting Monte Carlo replications to 1,000 for each case, and considering two different orderings of the variables. Hence, the interest rate spread series have been ordered as $w_t^{mar}, w_t^{war}, w_t^{depo}, w_t^{swap}, w_t^{repo}$ in the first case, with the order inverted in the second case. The median estimates have been obtained from cross-sectional distributions counting 8,000 units.[17] While inverting the order of the variables has no consequences for the estimation of the model, it is useful to make the impulse-response analysis and forecast-error-variance decomposition, carried out conditional to a double Choleski identification procedure (see Morana 2007b for details), robust to variable ordering.

5.1 Forecast-Error-Variance Decomposition and Impulse-Response Analysis

As shown in table 3, the results of the forecast-error-variance decomposition are clear-cut. Firstly, the bulk of forecast-error variance for the interest rate spreads is explained by the common shock—interpreted as shocks to MRO outcomes induced by changes in bidding behavior—with the proportion of explained variance increasing

[16]The latter is long enough to describe the long-range dependence in the series. Given the value of the fractional-differencing parameter and the length of the short-term autoregressive specification, the value of the $\Phi(L)$ parameters is negligible beyond the selected truncation order for the binomial expansion.

[17]Detailed results are not reported for reasons of space. A full set of results is available upon request from the authors.

Table 3. Median Forecast-Error-Variance Decomposition

		Pre-Turmoil Sample				**Full Sample**			
		Idiosyncratic			**Common**	**Idiosyncratic**			**Common**
	Horizon (Weeks)	Own	Other	All	—	Own	Other	All	—
mar	1	23.20	0.00	23.20	76.80	28.47	0.00	28.47	71.53
	2	15.71	0.60	16.31	83.69	19.59	0.46	20.05	79.95
	3	15.37	0.72	16.09	83.91	18.90	0.61	19.50	80.50
	4	15.35	0.83	16.18	83.82	18.59	0.66	19.25	80.75
war	1	16.84	1.95	18.79	81.21	22.68	3.87	26.55	73.45
	2	11.35	1.54	12.89	87.11	15.31	3.30	18.761	81.39
	3	10.68	1.67	12.35	87.65	14.62	3.47	18.09	81.91
	4	10.53	1.78	12.31	87.69	14.38	3.60	17.98	82.02
depo	1	11.79	0.31	12.11	87.89	18.54	0.59	18.54	81.46
	2	7.70	0.48	8.18	91.82	12.46	0.62	12.46	87.54
	3	7.25	0.61	7.86	92.14	12.12	0.87	12.12	87.88
	4	7.16	0.66	7.82	92.18	12.01	1.00	12.01	87.99

(continued)

Table 3. (Continued)

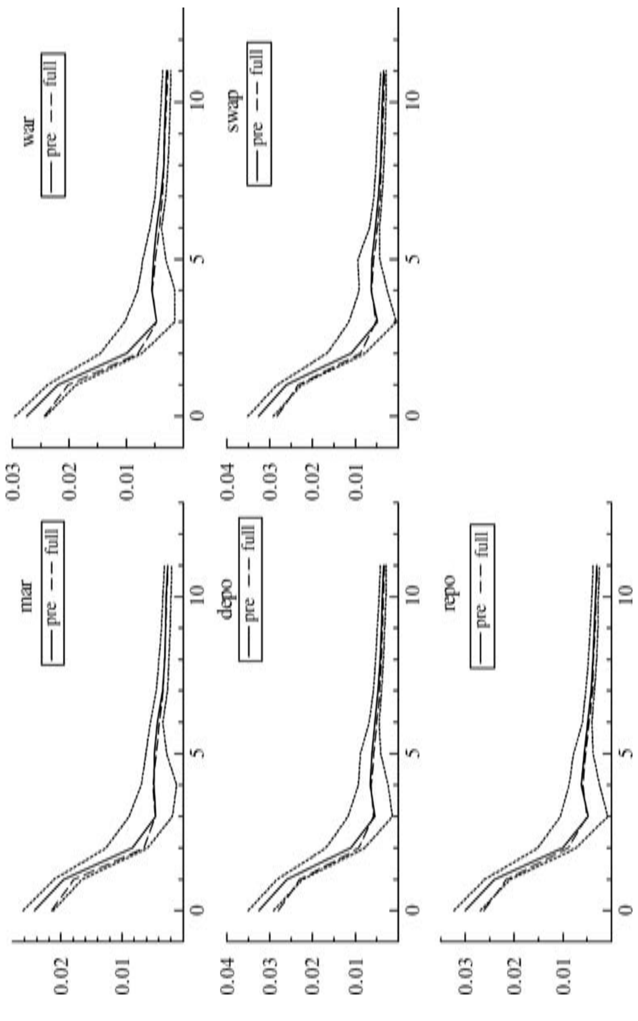
		Pre-Turmoil Sample				**Full Sample**			
		Idiosyncratic			**Common**	**Idiosyncratic**			**Common**
	Horizon (Weeks)	Own	Other	All	—	Own	Other	All	—
swap	1	13.00	0.33	13.32	86.68	18.60	1.33	19.93	80.07
	2	8.41	0.32	8.72	91.28	12.17	1.26	13.42	86.58
	3	7.90	0.45	8.35	91.65	11.59	1.46	13.06	86.94
	4	7.82	0.54	8.36	91.64	11.42	1.55	12.97	87.03
repo	1	15.84	0.90	16.75	83.25	21.30	3.22	24.52	75.48
	2	10.91	0.91	11.82	88.18	14.62	2.25	16.87	83.13
	3	10.46	0.99	11.44	88.56	13.95	2.25	16.20	83.80
	4	10.46	1.00	11.45	88.55	13.74	2.25	15.89	84.01

Notes: The table reports for each interest rate spread—i.e., the marginal rate (*mar*), the rate (*war*), (*depo*), (*swap*), and (*repo*)—the median forecast-error-variance decomposition, from the one-week to the one-month horizon, obtained from the structural VMA representation of the FI-F-VAR model, following the thick modeling estimation strategy. For each series the table reports the percentage of forecast-error variance attributable to the common and idiosyncratic shocks. For the latter the contribution of the own shock is distinguished from the cumulated effect of the other idiosyncratic shocks. The pre-turmoil period refers to the period June 27, 2000, through August 7, 2007, for a total of 372 observations, while the full sample period refers to the period June 27, 2000, through December 11, 2007, for a total of 390 observations.

with the forecast horizon. In fact, for the pre-turmoil sample, while at the one-week horizon the proportion of variance explained by the common shock is in the range 76 percent to 88 percent, at the one-month horizon the range is 84 percent to 92 percent. Moreover, full stabilization of the effects of the common shock is achieved within one month in all the cases. Secondly, as far as the idiosyncratic shocks are concerned (interpreted as money-market noise), the only non-negligible contribution to the explanation of variance is provided by the own idiosyncratic shock, explaining almost all of the remaining residual variability (in the range 8 percent to 15 percent at the one-month horizon and 12 percent to 23 percent at the one-week horizon), as the other non-own idiosyncratic shocks never explain more than 2 percent jointly. Similar findings can be noted for the full sample, although figures reveal a slightly larger role for the own idiosyncratic shock. In fact, while at the one-week (one-month) horizon the proportion of variance explained by the common shock is in the range 72 percent to 82 percent (81 percent to 88 percent), the proportion of variance explained by the own idiosyncratic shock at the one-week (one-month) horizon is in the range 19 percent to 29 percent (11 percent to 19 percent). The stronger role for the own idiosyncratic shock found for the full sample suggests that the market turmoil has affected the short-term structure of the euro-area money market. This finding is consistent with the cobreaking and cointegration analyses, which on the other hand point to stability in the long- and medium-term structure of the euro-area money market.

Concerning the impulse-response analysis, a similar dynamic reaction to the shocks can be detected for all interest rate series, with only the reaction to the common and the own idiosyncratic shocks being statistically significant. As is shown in figure 3, the dynamic reaction to the common shock points to a hyperbolic decay in all cases, consistent with the long-memory property of the spreads. Yet, the magnitude of the impact of the shock is sizable only within the first two weeks, with a unitary shock leading to a median contemporaneous increase in the spreads close to 3 basis points. Similarly, the decay of the response to the own idiosyncratic shock is also monotonic, though much quicker, and smaller in magnitude, than for the common shock: while the median contemporaneous impact of a unitary own shock is close to 1 basis point, its effects are already halved in magnitude after one week. These findings are consistent

Figure 3. Impulse Responses of Interest Rate Spreads to Common Shock



with the economic interpretation of the common shock (tender outcomes) and idiosyncratic shocks (money-market noise). Findings are very similar, and not statistically different, for both sample periods.[18]

5.2 Impact of Exogenous Variables

Table 4 reports the estimated median contemporaneous impact of the weakly exogenous variables included in the FI-FVAR model, with 95 percent confidence intervals. The upper panel shows the impact before the turmoil. The lower panel shows the impact including the turmoil period. We can use the results to answer the second set of questions asked in the introduction. For the pre-turmoil sample the median contemporaneous impact of the exogenous variables are all statistically significant and have the expected signs. Positive shocks to allotment volumes (A_t), interest rate expectations (F_t), and overnight rate volatility (E_t) put an upward pressure on spreads against the policy rate (MBR). Positive shocks to the policy liquidity variable have a negative impact on the spreads, acting as a counteracting factor (P_{t-1}). As predicted, the increasing liquidity deficit seems to exert a positive impact on bid shading, $w_t^{swap} - w_t^{mar}$, while the effect of volatility on bid shading seems negligible. When the turmoil period is included, a few changes can be noticed, illustrated by the slightly lower estimated impacts for P_{t-1} for all the series, while for A_t and F_t only *mar* and *war* show a smaller impact. Differently, a stronger impact can always be found for E_t. Still, all signs are as predicted. As expected, the impact of credit risk (CR_t) is not statistically significant for w_t^{repo} and w_t^{swap}, because these rates do not price credit risk (see Feldhütter and Lando 2007). The somewhat puzzling finding is that the credit-risk shock had a significant impact on ECB tender rates. In fact, this shock generates two deviations from equilibrium that are two sides of the same coin: an increase in the cost of liquid collateral ($w_t^{mar} - w_t^{repo}$) and a reversion in the sign of bid shading ($w_t^{swap} - w_t^{mar}$). The former is a movement that illustrates the dramatic and sudden disruption in securitization and

[18] Only the impulse-response functions for the common shocks are reported in the plots for reasons of space. A full set of results is available upon request from the authors.

Table 4. Median Impact of Exogenous Variables with 95% Confidence Interval

Pre-Turmoil Sample					
	mar	*war*	*depo*	*swap*	*repo*
P_{t-1}	−0.10	−0.16	−0.24	−0.20	−0.20
	(−0.12 −0.06)	(−0.19 −0.12)	(−0.28 −0.20)	(−0.24 −0.17)	(−0.24 −0.17)
A_t	2.84	3.54	4.85	5.52	4.84
	(2.34 3.29)	(3.04 4.12)	(4.31 5.48)	(4.92 6.44)	(4.35 6.05)
F_t	8.46	9.05	12.70	14.46	11.70
	(6.92 9.57)	(7.51 10.20)	(10.95 14.13)	(12.21 16.71)	(10.17 13.18)
E_t	1.90	3.03	2.35	1.80	1.96
	(0.86 2.80)	(1.58 4.04)	(1.07 3.54)	(0.72 3.17)	(0.79 3.31)
Full Sample					
	mar	*war*	*depo*	*swap*	*repo*
P_{t-1}	−0.04	−0.05	−0.11	−0.11	−0.10
	(−0.08 −0.02)	(−0.12 −0.04)	(−0.24 −0.09)	(−0.21 −0.09)	(−0.19 −0.09)
A_t	2.61	3.38	4.95	5.65	5.02
	(2.00 2.91)	(2.72 3.69)	(4.22 5.29)	(4.90 6.00)	(4.35 5.33)
F_t	7.32	8.24	12.89	14.68	12.08
	(6.30 8.71)	(6.99 9.80)	(11.10 14.62)	(12.65 16.25)	(10.28 13.51)
E_t	2.58	4.16	4.04	2.63	3.10
	(0.15 3.91)	(1.34 5.48)	(0.83 5.48)	(−0.25 4.08)	(0.28 4.41)
CR_t	0.16	0.17	0.13	−0.01	−0.01
	(0.14 0.18)	(0.15 0.19)	(0.12 0.14)	(−0.02 0.01)	(−0.02 0.01)

Notes: The table reports for each interest rate spread—i.e., the marginal rate (*mar*), the rate (*war*), (*depo*), (*swap*), and (*repo*)—the median impact of the exogenous variables (policy (P_t), allotment (A_t), forward (F_t), EONIA volatility (E_t), and credit (CR_t)), with 95% confidence interval. Figures are multiplied by 100. The pre-turmoil period refers to the period June 27, 2000, through August 7, 2007, for a total of 372 observations, while the full sample period refers to the period June 27, 2000, through December 11, 2007, for a total of 390 observations.

the resulting drying up of the asset-backed securities (ABS) market and is interpreted, within our framework, as a short-term disequilibrium movement. The latter illustrates the liquidity premium paid by banks to participate in the ECB tenders, which can be explained by the fact that ABSs are included in the collateral accepted for ECB operations, and have been increasingly used as collateral, while ABSs are not accepted in the private repo market. Moreover, the fact that toward the end of the sample period unsecured money-market rates were below ECB tender rates, $w_t^{depo} - w_t^{war} < 0$, is a clear sign of asymmetric information in the interbank market, pointing to the prevalence of credit rationing during the turmoil.

6. Conclusions

The main conclusions of the paper and the answers to the questions asked in the introduction are as follows. First, one-week interest rates in the euro area are cobreaking and the policy rate is the common-break process; this provides evidence on the effective steering of short-term interest rates by the ECB via the announcement of a minimum bid rate. Second, there is evidence of one common long-memory factor driving interest rate spreads against the policy rate, which is mainly related to shocks to the marginal ECB tender rate spread; this points to bidding behavior and tender outcomes as the driving force behind developments in the money-market spreads against the policy rate. These spreads are conditional on the liquidity policy followed by the ECB over the sample period and, therefore, the mean-reversion properties reveal the preference of the ECB for smoothing short-term interest rate spreads. The detected persistence in the spreads, and their statistical significance, shows that the ECB does not exercise perfect control of short-term money-market rates. This may be related to the low frequency of money-market interventions by the central bank and may characterize operational frameworks like those of the ECB and of the Bank of England which rely on the averaging mechanism of reserve requirements rather than daily open-market operations to stabilize short-term money-market rates (see Nautz and Scheithauer 2008 for a similar conclusion). The results point to another, related source of persistency in the spreads, which is the bidding behavior of counterparties. We found that the conditional means of the spreads are influenced by a number of shocks capturing the impact of changes in interest rate expectations, changes in allotment volumes, and interest rate uncertainty. All these variables put an upward pressure on the spreads. Allotments above benchmark had a downward and stabilizing impact on the spreads, being a counteracting force against the other factors. The market turmoil after August 2007 does not seem to have changed the long- to medium-term structure of the euro money-market spreads. However, credit risk and the associated funding risks have conditioned the short-term dynamics of money-market rates.

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