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The Informational Content of Trades on the EuroMTS Platform

by

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ABSTRACT

This paper presents unambiguous evidence that trading European government securities on EuroMTS contributes to determine their (unobservable) efficient price. Using twenty-seven months of daily transaction prices data for 107 bonds issued by eleven European governments, the estimated EuroMTS market's contribution to price discovery is about 20 percent, on average. Further, the amount of price discovery turns out to be strongly related to trading activity and price volatility conditions even controlling for institutional factors and for the maturity of bonds. Overall, the empirical results suggest that trades conveying information occur on EuroMTS when the level of liquidity is sufficiently high.

Keywords: European bond markets, price discovery, MTS system.

JEL Classification: G10, C21, C32.

NON TECHNICAL SUMMARY

Over the past decade, the European financial system has progressively evolved towards a paradigm with greater reliance on capital markets as a source of funding and risk mitigation and away from all-encompassing bank intermediation. Signs of the move in direction of a more market-based system have arisen largely out of the bond markets for government and corporate securities. In particular, the development of pan-European inter-dealer electronic trading platforms has been a key factor favouring the integration process of secondary market for Treasury bonds, with the MTS (Mercato Telematico dei Titoli di Stato) system emerging as the most relevant trading venue for euro-denominated government bonds.

One of the most striking features of the MTS system concerns the parallel listing of benchmark government securities on a domestic and on a European (EuroMTS) platform. If trades for the same security occur on two distinct market places, some new information about domestic MTS prices should be reflected in EuroMTS prices first, and vice-versa. As a consequence, the price discovery mechanism, that is the timely incorporation into prices of heterogeneous private information or heterogeneous interpretation of public information through trading, should take place in either trading venues.

This is the first paper to directly measure the percentage of price discovery across domestic and European MTS platforms. To do this, we employ an original and extensive dataset as compared to that of the existing literature. Our sample is of independent interest because its construction involved tracking 107 pairs of bonds over a 27-month horizon, using daily observations for benchmark government bonds traded on the MTS system issued by all euro area member countries' governments over the period January 2004 - March 2006. The extensiveness of our data sample allows us to explore not only the dynamic interactions between prices of trades occurring on the domestic MTS and the European platform, but also the cross-sectional variation in price discovery measures.

The paper reaches two main findings. *First*, we document that the architecture of the MTS system is able to eliminate persistent price discrepancies for the same bond traded on the two markets. The determination of the efficient price appears to take place with the contribution of trading activity on both platforms, with EuroMTS market's contribution to price discovery to be about 20 percent, on average. *Second*, estimation results reveal a systematic linkage between trading activity and price volatility and cross-sectional variability of price discovery taking place on the European platform. Trade cost differentials, instead, seem to have a minor role in explaining market players' preferences in trading government fixed income instruments on a platform rather the other. When institutional factors are included as additional

explanatory variables, the strong relationship between observable market characteristics and EuroMTS market's contribution to price discovery remains unaffected. The robustness of these results is checked across a number alternative specifications. Aside from their scientific merit, these conclusions are of direct importance for investors trading government securities on the EuroMTS platform and have relevant implication for regulators attempting to identify conditions likely to promote further integration in the European financial system. In this respect, the proliferation of alternative platforms for trading European government securities may be harmful if potential benefits from competition do not counterweight costs due to the increased liquidity fragmentation across trading venues. Further, the empirical evidence suggests that a wider standardisation of longer-maturity issuances and of regulatory arrangements in the primary markets could be beneficial.

IL CONTENUTO INFORMATIVO DELLE TRANSAZIONI EFFETTUATE SULLA PIATTAFORMA EuroMTS

SINTESI

Questo lavoro mostra come l'attività di scambio dei titoli di Stato sulla piattaforma EuroMTS contribuisca alla determinazione del loro prezzo efficiente (*price discovery*). Utilizzando dati giornalieri su un orizzonte temporale di 27 mesi per i prezzi di 107 differenti titoli di Stato emessi da 11 Paesi europei, il contributo del mercato EuroMTS alla determinazione del prezzo efficiente è circa del 20 per cento. La capacità di *price discovering* sulla piattaforma EuroMTS risulta, inoltre, fortemente connessa al volume delle transazioni e alla volatilità dei prezzi. Dai risultati delle stime si conclude che le transazioni effettuate sulla piattaforma EuroMTS favoriscono il processo di *price discovery* quando il livello di liquidità su tale mercato è sufficientemente elevato.

Parole chiave: Mercato europeo dei titoli di Stato, *price discovery*, piattaforma MTS.

Classificazione JEL: G10, C21, C32.

CONTENTS

1 INTRODUCTION[1]

Over the past decade, the European financial system has progressively evolved towards a paradigm with greater reliance on capital markets as a source of funding and risk mitigation and away from all-encompassing bank intermediation. Signs of the move in direction of a more market-based system have arisen largely out of the bond markets for government and corporate securities. While in 1992, the European bond markets were about half the size of their US counterparts in terms of the value of debt outstanding relative to GDP, they have by now almost converged, growing from 84 percent of GDP in 1992 to 145 percent in 2004, whereas US markets grew from 150 to 175 percent (Paesani and Piga, 2007). In particular, the development of pan-European inter-dealer electronic trading platforms (MTS, Icap/BrokerTec Eurex Bonds, eSpeed) has been a key factor favouring the integration process of secondary market for Treasury bonds, with the MTS (Mercato Telematico dei Titoli di Stato) system emerging as the most relevant trading venue for euro-denominated government bonds. According to the computations in Persaud (2006), the MTS system records around 72 percent volume of electronic trading of European cash Treasury fixed income instruments.

One of the most striking features of the MTS system concerns the parallel listing of benchmark government securities (i.e. on-the-run bonds with an outstanding value of at least 5 billion euro that satisfy listing requirements such as number of dealers acting as market makers) on a domestic and on a European (EuroMTS) platform. If trades for the same security occur on two distinct market places, some new information about domestic MTS prices should be reflected in EuroMTS prices first, and vice-versa. As a consequence, the price discovery mechanism, that is the timely incorporation into prices of heterogeneous private information or heterogeneous interpretation of public information through trading, should take place in either trading venues.

Establishing that the price discovery process involves both markets has important practical implications for traders watching signals about future price movements. Further, assessing whether trading on EuroMTS convey information has a relevant institutional significance with respect to the recent debate on the restructuring of the bond segment of European financial system. At first sight, indeed, the European trading venue might seem redundant as all bonds being traded on that market are a fraction of the bucket of securities

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traded on the respective domestic trading system (Cheung et al., 2005). Moreover, the adoption of the Directive 2004/39/EC, disciplining the functioning of Markets in Financial Instruments in Europe (MiFID) has stimulated an intense debate among academics and practitioners on how and whether extending the MiFID regime to the government bond market (see Anolli and Petrella, 2007; Paesani and Piga, 2007). Thus, understanding what factors drive market participants' willingness to trade on EuroMTS could offer useful insight into the effects of market segmentation due to the possible proliferation of alternative trading venues.

Previous works on the European secondary bond markets have been focusing on the dynamic relationship between trading activity and price movements (Cheung et al., 2005) or between yield dynamics and order flow (Menkveld et al., 2004), on the determination of the benchmark status among government securities of similar maturity (Dufour and Nguyen, 2007; Dunne et al., 2007), on the analysis of yield differentials between sovereign bonds in the Euro area (Favero et al., 2005; Beber et al., 2008).

To the best of our knowledge, this is the first paper to directly measure the percentage of price discovery across domestic and European MTS platforms. To do this, we employ an original and extensive dataset as compared to that of the existing literature. Our sample is of independent interest because its construction involved tracking 107 pairs of bonds over a 27-month horizon, using daily observations for benchmark government bonds traded on the MTS system issued by all euro area member countries' governments, except for Luxemburg, over the period January 2004 - March 2006.

We contribute to the growing empirical literature on the European secondary government bond market on two dimensions. *First*, we use the methodology proposed by Harris et al. (1995) and Hasbrouck (1995) to provide an empirical assessment of the contribution to price discovery of trades on the EuroMTS platform. While these approaches have been applied to stock (Harris et al., 1995; Hasbrouck, 1995; Huang, 2002), credit derivatives (Blanco et al., 2005) and foreign exchange (Tse et al., 2006) markets, there is scant empirical evidence for the market of government fixed income securities. Noteworthy exceptions are the works by Upper and Werner (2002a, 2002b), Brandt et al. (2007) and Chung et al. (2007), where the dynamic interactions between spot and future prices are examined. Here, instead, we focus on two cash markets (the domestic MTS and EuroMTS platforms). We document that about 20 percent of price discovery occurs in the European trading platform, on average, with estimates for individual bonds ranging from less than 3 percent to 56 percent. *Second*, having found evidence of significant role for trades on EuroMTS in determining the (unobservable) efficient price, we focus on what factors contribute to price discovery on that market. We find that its contribution is greater when relative trading activity (proxied by the EuroMTS/domestic MTS ratio of trading volume or, alternatively, by the relative number of trades) is higher, and when relative price volatility (proxied by the EuroMTS/domestic MTS ratio of the standard deviation of the first differenced logarithms of transaction prices) is lower. We also find limited evidence that the level of contribution to price discovery on EuroMTS is negatively related to the effective bid/ask spreads differentials between domestic and European trading platforms. In addition, we investigate whether the level of EuroMTS market's contribution depends on institutional factors such as member states' degree of protection in auctioning securities in the primary market, differences in market making obligations between domestic and European MTS markets, maturity effects, differences between large and small borrowers and other country specific factors. Even though we find a non-negligible role for a vast majority of these variables, the strong positive (negative) linkage between relative trading activity (price volatility) remains unaffected, suggesting that market conditions may be the primary force driving price discovery in the European platform. These results are robust across a number of modifications and extensions of the baseline empirical design.

The paper is structured as follows. In Section 2, we describe the most relevant institutional features of the MTS system along with the set of the research questions tackled throughout the rest of the paper. In Section 3, we present the statistical methods as well as data we use to compute estimates of price discovery taking place in the EuroMTS market. In Section 4, we seek to explain cross-sectional variation in price discovery measures of EuroMTS trades. In Section 5, we report some additional robustness tests and extensions of the baseline model. Section 6 summarizes our findings and contains suggestions for future research. Appendices containing the list of bonds involved in the empirical analysis and the construction of explanatory variables for the cross-sectional analysis conclude.

2 PRICE FORMATION IN THE MTS SYSTEM

2.1 The institutional architecture: A duplicated market setting for benchmark securities

Trading on the secondary Treasury market can occur *via* four channels: inter-dealer (B2B) platforms and dealer-to-customer (B2C) electronic trading platforms, either multi-dealer or single-dealer, OTC inter-dealer via voice brokers and OTC dealer-to-customer trading. B2B platforms serve essentially for the trading of Treasury bonds and generally operate via cross-matching methods. In the European case, MTS, Icap/BrokerTec Eurex Bonds and eSpeed are the most widespread B2B platforms.

In the MTS system, market makers' quotes are aggregated in a single order book to match best anonymous bids and offers automatically, subject to non-discretionary priority rules. The ability to bring together issuers, with long-term financing needs, and dealers, willing to place liquid funds in interest-bearing securities, and to induce them to a mutual commitment (the "liquidity pact") constitutes the key to its widespread success (Pagano and von Thadden, 2004).

Trades are anonymous and the identity of the counterpart is only revealed after an order is executed for clearing and settlement purposes, so as to avoid free-riding generated by the existence of less sophisticated traders and allowing for liquidity providers to reduce their exposure when trading (Albanesi and Rindi, 2000; Massa and Simonov, 2003)[2].

As far the type of market participants on the MTS system, we can categorize them either as market makers (primary dealers) or as market takers (dealers). In the light of the recent debate on extending the MIFiD regime to the Treasury bond market, it is worth recalling obligations faced by primary dealers. They include: *i)* stringent capital requirements and trading protocols, *ii*) obligation to continuously post firm two-way prices for a selected subset of securities; *iii*) price-posting for at least five hours per day and for a certain minimum quantity; *iv*) possibility to be subject to maximum spread obligations. In return, they are the market participants entitled to participate in supplementary auctions and may gain other privileges. By contrast, dealers cannot enter quotes into the system and are obliged to trade bonds on the basis of bid/ask quotes placed by the primary dealers. In the primary market, a subset of primary dealers is committed to subscribe to specified shares of auctions, thus establishing a possible interplay between practices on the primary market and trading strategies in the secondary market.

All government marketable securities, including benchmark bonds, are listed on their respective domestic MTS platforms. Only benchmark securities, or on-the-run bonds with an outstanding value of at least 5 billion euro that satisfy listing requirements such as number of dealers acting as market makers, are admitted, instead, to trading on the wholesale European market

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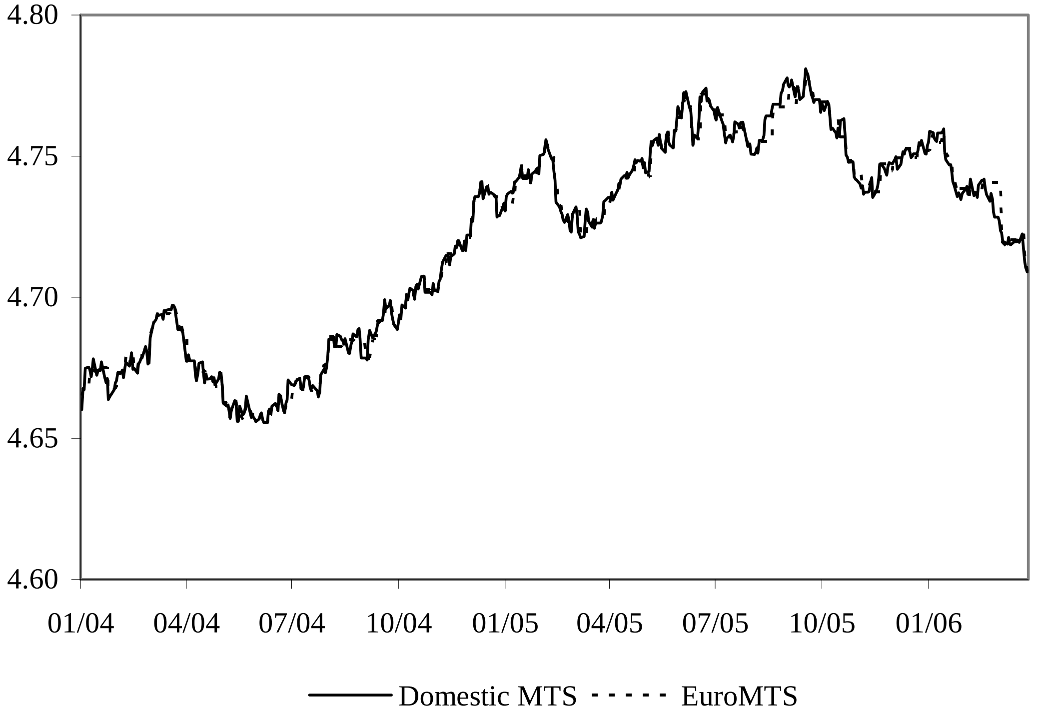
[2] Recently, the full anonymity has been reached by means the introduction of the central counterparty (CCP) system, which aims at eliminating any risk faced by participants in trading with other dealers. For a detailed discussion of MTS, see Scalia and Vacca (1999).

(EuroMTS[3]). For benchmark securities, thus, dealers are allowed to post their quotes on both market simultaneously (parallel quoting).

As an illustrative example of a typical security contemporaneously traded on EuroMTS and a domestic MTS market, we present in Figure 1 (the logarithm of) daily transaction prices for a 15-year bond (coded as IT0003242747) with fixed coupons paid at the annual rate of 5.25 percent issued by the Italian Treasury on February 1 2002 with maturity date on August 1 2017, over the period January 2004 - March 2006.

Fig. 1 Daily transaction prices on EuroMTS and on the domestic MTS



Note. Dashed (continuous) line indicates the logarithm of daily transaction prices recorded on the EuroMTS (domestic) platform. The plots refer to prices of a 15-year bond (coded as IT0003242747) issued by the Italian Treasury, over the period January 2004 - March 2006.

2.2 Prices dynamics of benchmark securities in the MTS system

The price of a bond at time t, P_t, can be expressed as the sum of its face value and all future coupon payments discounted by the yield-to-maturity. Being

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[3] Designed by the Italian MTS Group, the London-based EuroMTS was set up in 1999 as a trading venue for euro-denominated benchmark bonds. In April 2003, the first month of available data, the bond traded on EuroMTS are from 15 government and quasi-government issuers, namely all euro area member countries, except for Luxembourg, Depfa, the European Investment Bank, Freddie Mac and Kreditanstalt für Wiederaufbau (Dufour and Skinner, 2004).

P_t a forward-looking quantity in nature, only new information arrivals (due to macroeconomic releases and policy announcements and statements) should cause revisions to what is built into the current price of the bond (Andersson *et al*., 2006). We retain the assumption that information arrivals may affect the yield to maturity and thus the dynamics of government bond prices, which can be characterised as non-stationary processes, along the lines of Albanesi and Rindi (2000)[4].

Consider a government benchmark security traded on both EuroMTS (E) and the domestic MTS (D) platform. Its (log-) price on market $j = E, D$ at time t, p_t^j, can be represented as the sum of a permanent component, ϕ_t, and a market-specific transient part, υ_t^j

$$p_t^j = \phi_t + \upsilon_t^j \tag{1}$$

The law of motion of the stochastic trend, ϕ_t, is assumed to be $\phi_t = \phi_{t-1} + \mu_t^\phi = \phi_0 + \sum_{i=1}^{t} \mu_i^\phi$, where the ϕ_0 term captures initial conditions and μ_t^ϕ is an uncorrelated white noise process such that $E(\mu_t^\phi) = 0$, $E(\mu_t^\phi)^2 = \sigma_\phi^2$, $E(\mu_t^\phi \mu_s^\phi) = 0$ for $s \neq t$. This set of assumptions implies that ϕ_t behaves as a random walk. The transitory disturbance υ_t^j, instead, is modelled as a covariance stationary process, following an ARMA scheme $\upsilon_t^j = \sum_{i=1}^{\infty} \delta_i^j \xi_{t-i}^j = \delta^j(L)\xi_t^j$, where the elements of the polynomial in the lag operator L, $\delta^j(L)$, are market-specific parameters and ξ^j's are independently distributed with mean zero and constant variance[5]. Thus, the difference between a generic pair of bond prices recorded on the two trading venues is

$$p_t^E - p_t^D = \delta^E(L)\xi_t^E - \delta^D(L)\xi_t^D = \varepsilon_t \tag{2}$$

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[4] This is a standard practice used in the analysis of stock market prices (see, among others, Hasbrouck, 1995; Harris et al. 1995). As pointed out by Albanesi and Rindi (2000), in the case of bond prices, such a representation is correct as far as the series used do not include the whole life of the asset.

[5] Given only the observed transaction prices the decomposition in equation (1) is unidentified. This implies that, even with an infinite sample of past and future transaction prices, neither ϕ_t nor υ_t^i can be exactly determined. The literature on permanent and transitory decompositions offers several ways to split the price vector in permanent and transient components, depending on the conditions imposed on the relationships between ϕ_t and υ_t^i and on the stochastic properties of these two components. In this work we focus on the approaches proposed by Harris et al. (1995) and Hasbrouck (1995).

where the disturbance ε_t is a linear combination of stationary processes and thus stationary itself.

According to the Law of One Price (LOP) in condition (2), the price of a government bond should reflect the same information arrival irrespective of which of the two trading platforms are considered. Thus, p_t^E and p_t^D are expected to be driven by a common factor, the $\sum \mu_i^\phi$ term, which represents the efficient price related to news cumulating over time, while the ε_t term should capture market-specific transient noises, affecting the speed at which market participants in a specific market process the information flows.

2.3 Empirical issues to be addressed

Figure 1 shows a close overlapping of the two log-price series, albeit some deviations occur. Understanding what factors originate those discrepancies constitutes the bulk of our empirical analysis. More specifically, the main purposes of this paper can be summarised as follows.

A - Price convergence. Transaction prices of the same bond recorded on different trading venues are not independent of one another. The LOP condition (2) dictates that these prices may exhibit individually a non-stationary behaviour, but they should be linked to one another by a stationary long-run equilibrium. Thus, discrepancies are expected to be temporary in nature. The empirical implication of the LOP can be suitably captured by specifying, for each pair (p_t^E, p_t^D), a dynamic system and testing whether p_t^E and p_t^D are cointegrated for the equilibrium condition (2) to hold.

B - Price discovery. With closely related securities traded in different market places (as in the case of the government bonds traded on the MTS system) the timely incorporation of heterogeneous private or heterogeneous interpretation of public information into market prices is split among trading venues (Lehmann, 2002). Thus, both the domestic and the European MTS markets are expected to contribute to price discovery, although to a different extent. Assessing whether trading euro-denominated government bonds on EuroMTS has an informative content is a relevant institutional issue, since, at first sight, the European trading might seem redundant ("redundancy hypothesis") as all bonds being traded on this market are also traded on their respective domestic trading counterparts (Cheung et al., 2005).

C - Price discovery and observable market characteristics. The speed at which information arrivals are processed and, thus, the contribution to price discovery may be influenced by market-specific characteristics. Consistently with previous empirical works on the determinants of price discovery measures

(Eun and Sabherwal, 2003; Chakravarty et al., 2004, among others), markets' contributions are likely to be systematically associated to trading activity, transaction costs and price volatility measures. Establishing the role of those possible determinants could be of interest to promote a regulatory framework aimed at achieving a more integrated secondary market for government securities in Europe.

D - Price discovery and institutional factors. Answering why dealers operating on the domestic MTS platforms would also be willing to operate on EuroMTS in the light of observable market-specific characteristics alone may be partial. Institutional arrangements may confound, indeed, the linkage between market liquidity and price discovery (Huang, 2002). How institutional factors (for instance, anonymous trades) may impinge the nexus between price discovery and observable market characteristics is less than clear and calls for a careful empirical investigation.

3 MEASURING PRICE DISCOVERY IN THE EUROMTS PLATFORM

3.1 Econometric framework

Adopting the same notation as introduced in Section 2.2, the following Vector Error Correction (VEC) model (Johansen, 1995) constitutes the basis of our investigation

$$\begin{bmatrix} \Delta p_t^E \\ \Delta p_t^D \end{bmatrix} = \Pi \cdot \begin{bmatrix} p_{t-1}^E \\ p_{t-1}^D \end{bmatrix} + \sum_{j=1}^{k-1} A_j \cdot \begin{bmatrix} \Delta p_{t-j}^E \\ \Delta p_{t-j}^D \end{bmatrix} + \begin{bmatrix} u_t^E \\ u_t^D \end{bmatrix}, \; E\left(u_t \cdot u_t'\right) = \Sigma = \begin{bmatrix} \sigma_E^2 & \rho\sigma_E\sigma_D \\ \rho\sigma_E\sigma_D & \sigma_D^2 \end{bmatrix} \quad (3)$$

where Δ is the first difference operator, A's are matrices of autoregressive coefficients up to the order $k-1$, u's are the residuals with variance-covariance matrix Σ, where ρ is the correlation coefficient and σ's are standard deviations. If the LOP condition (2) holds, we expect rank equal to 1 for matrix Π, i.e. the log-two price series sharing a common stochastic factor. In this case, the long-run matrix can be factored as

$$\Pi = \begin{bmatrix} \alpha_E \\ \alpha_D \end{bmatrix} \cdot \begin{bmatrix} 1 & -1 \end{bmatrix} \quad (4)$$

with $\alpha^E < 0$ and $\alpha^D > 0$. This implies that prices do not drift too far from one another, suggesting that pricing errors are corrected over time through the feedback parameters collected in α. These coefficients provide information on the rate at which the corresponding market processes information arrivals and constitute a fundamental ingredient for the computation of each market's contribution to the determination of the (unobservable) efficient price[6].

The common factor models proposed by Harris et al. (1995) and Hasbrouck (1995) are elegant ways to capture where price discovery occurs in closely linked securities traded in multiple markets. Even though both models build on a VEC framework as the one in (3), the two measures for efficient price determination reflect different definition of price discovery. Using the decomposition proposed by Gonzalo and Granger (1995), Harris et al. (1995) attribute superior price discovery to the market that adjusts the least to price movements in the other market by decomposing the common factor itself. More formally,

$$\gamma_E = \frac{\alpha_D}{\alpha_D - \alpha_E} \ , \ \gamma_D = \frac{\alpha_E}{\alpha_E - \alpha_D} \tag{5}$$

so that, EuroMTS (domestic MTS) market's contribution, γ_E (γ_D), is defined to be a function of both α's. Based on the Cholesky factorisation of matrix Σ, Hasbrouck's model assumes, instead, that market's contribution to price discovery should be (positively) related to market's contribution to the variance of the innovations to the common factor (market's information share). Since price innovations are generally correlated across markets, matrix Σ is likely to be non-diagonal. In such an occurrence, Hasbrouck's approach can only provide upper and lower bounds on the information shares of each market. For the EuroMTS market, these bounds are

$$S_E^{ub} = \frac{(\gamma_E\sigma_E + \rho\gamma_D\sigma_D)^2}{(\gamma_E\sigma_E + \rho\gamma_D\sigma_D)^2 + \gamma_D^2\sigma_D^2(1-\rho^2)} \ , \ S_E^{lb} = \frac{\gamma_E^2\sigma_E^2(1-\rho^2)}{\gamma_E^2\sigma_E^2(1-\rho^2) + (\rho\gamma_E\sigma_E + \gamma_D\sigma_D)^2}$$

respectively. However, Baillie et al. (2002) argue that the average of these bounds

[6] Ideally, either or both α^E and α^D are likely to respond to deviations from the long-run equilibrium relationship. On the one hand, as domestic MTS is the home-market of our sample of government bonds, we expect that EuroMTS prices adjust to some extent to departures from the LOP condition (2). On the other hand, as the EuroMTS is the wholesale European market for euro-denominated benchmark government bonds, we expect some feedback going the other way around.

$$\zeta_E = \frac{1}{2}(S_E^{ub} + S_E^{lb}) \qquad (6)$$

provides a sensible estimate of the markets' roles in the mechanism of determination of the efficient price. Both γ_E and ζ_E can range in the interval [0,1], where the higher their value is the most EuroMTS market contributes to price discovery. In what follows, we employ these two statistics to assess the informational content of trades on the EuroMTS platform[7].

3.2 Data description

Data are taken from MTS Time series database. Daily observations cover the period from January 2 2004 to March 31 2006; a total of 27 months comprising 586 trading days. For each trading day, we have a time stamp, the nominal value of trading volume, the average size of trades, the last transaction price recorded before the 17.30 Central European Time close, and the average best bid/ask spread throughout the trading day[8]. Furthermore, we use information on the issuer country, the issuing and maturity dates, the hours in a trading day that dealers must have an active quote, the maximum spread that is quoted and the minimum quantity that a dealer can bid or offer.

The government bond markets involved in the empirical analysis are those of Austria (ATS), Belgium (BEL), Germany (DEM), Spain (ESP), Finland (FIN), France (FRF), Greece (GGB), Ireland (IRL), Italy (MTS), the Netherlands (NLD) and Portugal (PTE)[9]. For each country, we select all benchmark government bonds traded in January 2004 maturing after the end of our estimation horizon. It translates into a collection of 107 securities. Table 1 summarises the chosen bonds, classified by issuer and maturity. Appendix A provides the entire list of bond codes.

[7] See Ballie et al. (2002), among others, for a detailed discussion and a formal derivation of the two price discovery measures.

[8] Previous studies on price discovery have used data of varying frequency, ranging from daily (Blanco et al., 2005) down to few seconds (Hasbrouck, 1995). Green and Joujon (2000) argue that daily resettlement creates a strong argument for using daily closing prices, since they determine the cash flows of traders. In general, frequent sampling is generally desirable, as it ensures low correlations among the innovations in prices. However, Shiller and Perron (1985) and Hakkio and Rush (1991) show that the power of many tests commonly used in financial market research does not increase with the number of observations, unless this translates into an extension of the data span.

[9] Luxembourg is not included since the lack of fixed income securities having a *government* status.

Tab. 1 Selected benchmark government bonds by maturity and issuer

	Short/medium maturity	Long maturity	Very long maturity	*Sum by country*	*Percentage by country*
ATS	0	8	2	*10*	*9.3*
BEL	2	5	2	*9*	*8.4*
ESP	3	5	2	*10*	*9.3*
FIN	3	3	0	*6*	*5.6*
FRF	6	5	3	*14*	*13.1*
GEM	6	5	2	*13*	*12.1*
GGB	3	6	3	*12*	*11.2*
IRL	2	1	1	*4*	*3.7*
MTS	7	6	4	*17*	*15.9*
NLD	2	3	1	*6*	*5.6*
PTE	2	3	1	*6*	*5.6*
Sum by maturity	*36*	*50*	*21*	*107*	.
Percentage by maturity	*33.6*	*46.7*	*19.6*	.	*100.0*

Note. According to the classification in Dunne et al. (2007), government bonds with small-medium maturity are those with maturity less than 6.5 years; long maturity and very long maturity securities refer to instruments with maturity ranging between 6.6 and 13.5 years and more than 13.5 years, respectively. The first and the second row (column) in italics present the sum and the percentage by maturity (issuer), respectively. Market codes are from Dufour and Skinner (2004). See Section 3.2 of the paper for details on the criteria for inclusion in the sample.

3.3 VEC models: Estimation results

Standard cointegration methods require equally spaced data without missing values. Following Upper and Werner (2002a), in the presence of missing observations we use the last available transaction price ("fill-in" method). The estimation horizon ranges from 557 to 585 observations, with an average value of 580 daily datapoints. As a preliminary exercise, we check for the presence of a unit root in each of 214 individual transaction price series expressed in logarithms. ADF tests (Dickey and Fuller, 1979) are performed on the series, both in levels and first differences[10]. In each case, we are unable to reject the null hypothesis of a unit root at conventional levels of significance. On the other hand, differencing the series appears to induce stationarity. The KPSS stationarity tests (Kwiatkowski et al., 1992) corroborate these results (not reported to save space).

[10] The optimal lag is determined on the basis of the AIC, with the maximum tested lag set equal to four. Critical values for these tests are provided by Davidson and MacKinnon (1993).

Having identified that all series involved in the analysis are $I(1)$ variables, cointegration techniques are used next to examine the existence of a (stationary) long-run relationship for all 107 pairs (p_t^E , p_t^D). This translates into the estimation of 107 VEC models, as the one in equation (3), testing whether the restriction (4) is not rejected by the data. Here is an overview of the estimation results. The order of autoregression k of the VEC models, formulated in their isomorphic Vector AutoRegression (VAR) representation, is chosen on the basis of the AIC, the BIC and LR-based system reduction tests, with the maximum lag length set equal to four. In the presence of discordant results among lag determination criteria, we prefer the AIC in order to ensure richer system specifications[11]. Overall, the order of autoregression is quite limited: k=1, k=2, k=3 and k=4 is chosen for 62, 25, 15 and 5 entities of reference, respectively.

The trace test (Johansen, 1995) suggests choosing rank 1 for the Π matrix in 104 entities of reference[12], giving compelling support to our *a priori* theoretical assumptions. The symmetry and proportionality condition implied by the LOP condition (2) is tested by a χ^2-distributed LR test with one degree of freedom. In 88 entities of reference, the over-identifying restriction is not rejected by the data (at least) at the 10 percent level of significance, while in 6 cases (at least) at the 5 percent level. For the remaining 10 models the evidence is less conclusive, even though the cointegration test developed by Horvath and Watson (1995) supports the validity of the LOP, as shown in Table 2[13]. In what follows, we focus on the dynamic properties of the 104 models satisfying the condition (4).

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[11] Both the AIC and the BIC suggest similar lag length in most of the cases. The Monte Carlo study by Cheung and Lai (1993) show that the AIC and the BIC indicate the correct lag order of a VAR used for testing for cointegration in 99.86 and 99.96 percent of cases, respectively.

[12] Johansen's maximum likelihood approach to cointegration allows for testing procedures which are fairly robust to the presence of non-normal (Cheung and Lai, 1993) and heteroskedastic disturbances (Lee and Tse, 1996). Notice that the VAR specification here considered is model $H_1^*(r)$ in Johansen's notation, where a linear deterministic trend is implicitly allowed for but this can be eliminated by the cointegrating relations and the process contains no trend stationary components. In three cases (FI0001005514, GR0110014165, IT0003522254), the rank of the long-run matrix turns out to be two. On the one hand, this finding is at odds with the conclusions from the unit root/stationarity tests; on the other hand, it confirms that the LOP holds in these three cases too. On the lack of power of unit root test see, among others, Sarno and Taylor (2002).

[13] The test of the null of no cointegration against the known alternative of rank one with $\beta' = [1 \ -1]$ corresponds to a Wald test for the inclusion of the error-correction term, i.e. the LOP condition, in a VAR in first differences with an unrestricted constant. The test statistics is computed as $2(\ln LL_{VECM} - \ln LL_{VAR})$, where LL denotes the value of the likelihood function under the respective model.

Tab. 2 Results of the Horvath and Watson (1995), HW, cointegration test

	χ^2	p-value	LL_{VECM}	LL_{VAR}	HW
BE0000296054	5.33	0.0209	5672.8	5612.6	120.48
BE0000297060	6.67	0.0098	6660.3	6627.2	66.21
BE0000302118	5.34	0.0209	6200.9	6139.0	123.77
ES0000012445	7.76	0.0053	6914.9	6898.1	33.75
ES0000012452	5.27	0.0216	5634.6	5606.1	56.96
FR0000187635	5.43	0.0198	4634.0	4611.8	44.53
FR0104446556	9.18	0.0025	6556.5	6516.3	80.41
GR0124011454	5.06	0.0245	5832.9	5746.1	173.65
NL0000102697	9.27	0.0023	7153.9	7105.3	97.34
PTOTEGOE0009	5.26	0.0218	5460.6	5378.4	164.47

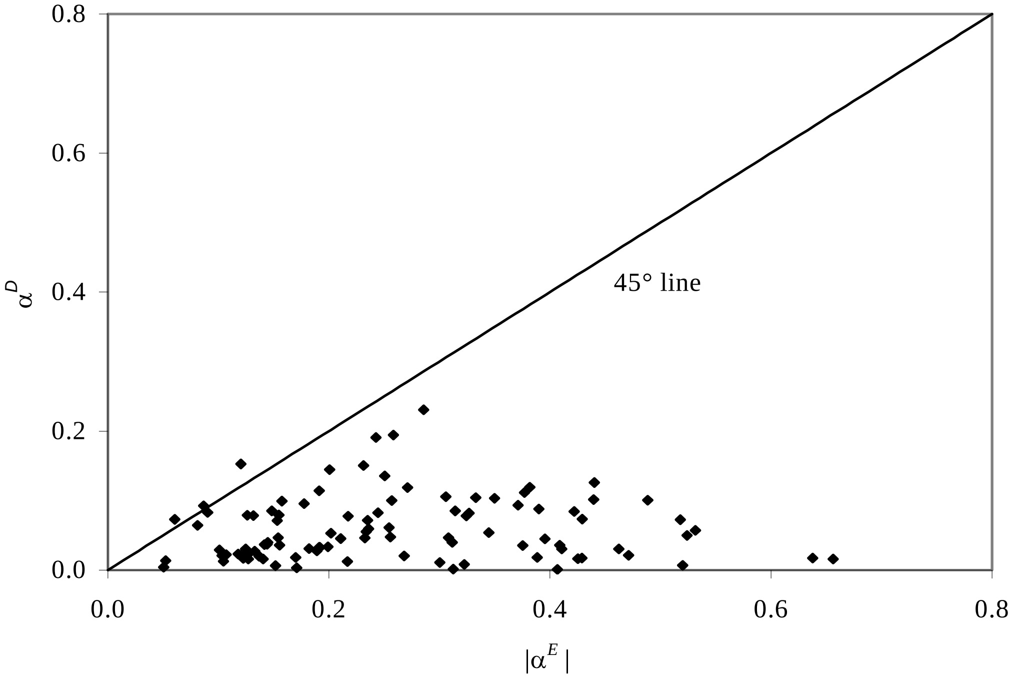
Note. The first numerical column reports the χ^2-distributed LR test statistics associated to the over-identifying restriction implied by the LOP (see equation (2) of the paper). The corresponding p-values are reported in the second column. The HW test (Horvath and Watson, 1995) of the null of no cointegration against the known alternative of rank one with cointegration space $[1 \; -1]$ (see equation (4) of the paper) corresponds to a Wald test for the inclusion of the error-correction term in a VAR in first differences with an unrestricted constant. Its test statistic (last column of the Table) is computed as $2(\ln LL_{VECM} - \ln LL_{VAR})$, where LL denotes the value of the likelihood function under the respective model. The 5 percent and 1 percent critical value is 10.18 and 13.73, respectively.

Estimation results reveal a number of interesting aspects with respect to the interactions between transaction prices on the domestic MTS and on the European MTS markets. *First*, the feedback coefficients associated to the Δp_t^E equation in (3) are statistically significant at the 1 percent level in all models; by contrast, only one half of the estimated α^D coefficients turn out to be statistically significant (at the 1 percent level in 24 entities of reference, at the 5 percent in 15, at the 10 percent in the 12 remaining cases). *Second,* both α^E and α^D are correctly signed, implying direct convergence to the long-run relationship in all but six models (where the estimated α_D's are negative). Figure 2 plots α^D versus $|\alpha^E|$ for our 98 entities of reference we retain after discarding those with wrongly signed α^D's[14]. Most observations are substantially off the horizontal and vertical axes, suggesting that domestic MTS prices react to those formed on the European trading venue, and viceversa. *Fourth*, greater adjustments seem to take place in the European platform, since

[14] In Section 5, we show that the results are not affected by such a choice.

$|\alpha^E|$'s are larger than their corresponding α^D's in all but three models, with the average value for $|\alpha^E|$ equal to 0.26 as compared to 0.06 for α^D (Table 3). This conclusion is confirmed by testing the null $H_0 : |\alpha^E| = \alpha^D$: the LR test statistics turns out to be greater than 3.84 (the 95 percent critical value for a χ^2 distribution with one degree of freedom) for a majority of bonds (82 out of 98)[15]. *Finally*, the median of the overall adjustment process, $|\alpha^E| + \alpha^D$, is 31.95 percent, as reported in Table 3[16].

Fig. 2 **Scatter plot of feedback coefficients: α^E versus α^D**



Note. α^E and α^D are the adjustment coefficients to deviations from the long-run equilibrium condition implied by the LOP (see equation (2) of the paper) relative to the European price changes equation and the domestic price changes equation, respectively. See model (3) under the condition (4) of the paper. Points below (above) the 45° line indicate that $|\alpha^E|$ is greater (lower) than α^D. Computations based on 98 bivariate VEC models.

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15 A similar picture is obtained by comparing the R^2_{adj} for the two dynamic equations of the system (3) under condition (4). We find that, across all 98 bonds involved in the analysis, around 13.9 percent of the variation in Δp_t^E is explained by the model, a larger value with respect to 1.6 percent calculated for Δp_t^D. This suggests that most of the variability in Δp_t^D changes represents "news" arriving in the market and indicates that domestic MTS prices play a leading role in incorporating new information.

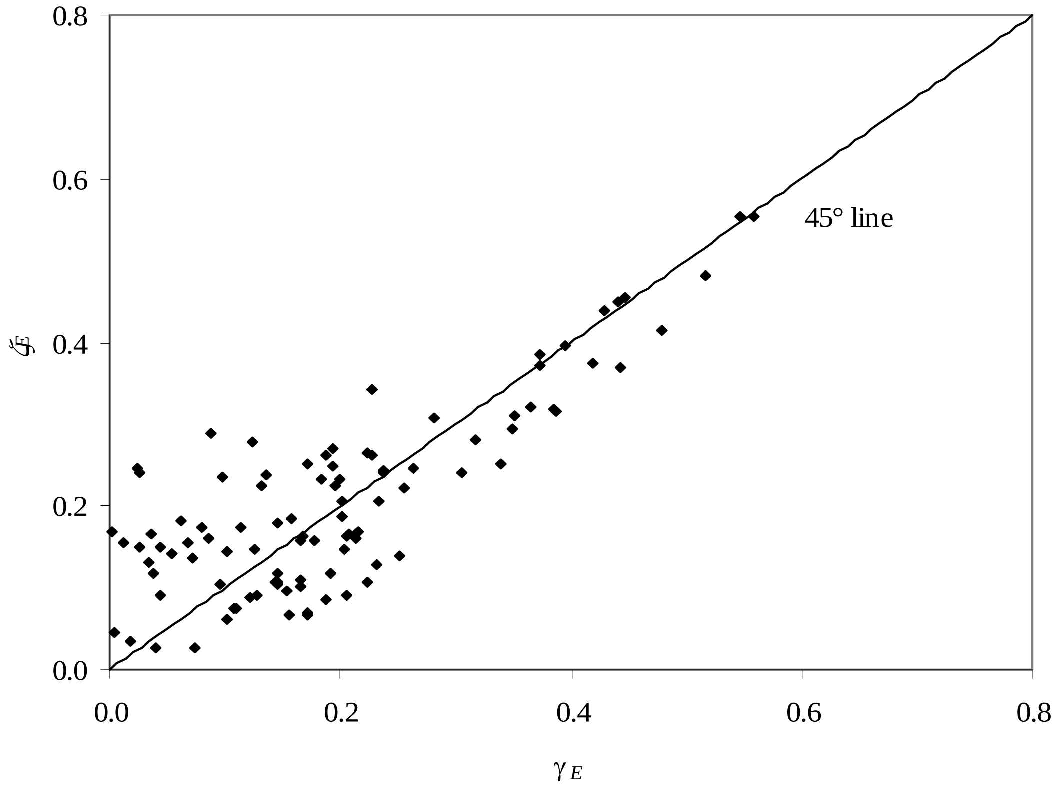
16 Our estimates suggest half-life deviations from the equilibrium condition, $n = \{\ln 0.5 / \ln[1 - (|\alpha^E| + \alpha^D)]\}$, lasting around two days, on average. As a result, the ratio between the sample length in terms of datapoints and the half-life is around 300. This adds confidence to our results, especially in the light of the Monte Carlo study by Hakkio and Rush (1991), who show that in cointegration analysis, the ratio of the length of the data set to the half-life is more relevant than the length of the data set alone.

Tab. 3 **Estimated values of the feedback coefficients**

	α^E	α^D	$\lvert\alpha^E\rvert+\alpha^D$
Mean	-0.2601	0.0594	0.3195
Minimum	-0.6561	0.0009	0.0547
Maximum	-0.0506	0.2304	0.6718
5th percentile	-0.5183	0.0066	0.1297
25th percentile	-0.3657	0.0222	0.1887
Median	-0.2345	0.0465	0.3042
75th percentile	-0.1457	0.0849	0.4395
95th percentile	-0.0898	0.1453	0.5761

Note. α^E and α^D are the adjustment coefficients to deviations from the long-run equilibrium condition implied by the LOP (see equation (2) of the paper) relative to the European price changes equation and the domestic price changes equation, respectively (see model (3) under the condition (4) of the paper). The last column refers to the overall speed of adjustment, $\lvert\alpha^E\rvert+\alpha^D$. Computations based on 98 bivariate VEC models.

Fig. 3 **Scatter plot of price discovery measures for the EuroMTS platform: γ_E versus ζ_E**



Note. The price discovery measures for the EuroMTS trading platform, γ_E and ζ_E, are defined by equation (5) and (6) of the paper, respectively. Points on the 45° line indicate equivalence between estimated values for γ_E and ζ_E. Computations based on 98 bivariate VEC models.

3.4 Estimated price discovery measures

Examining the price discovery measures given in (5) and (6) is a more direct way to assess whether the interplay of price dynamics on the two platforms is conducive to significant price discovery in the European platform. Estimated values of γ_E for individual entities of reference range from 0.2 percent (IT0003357982) to 55.9 percent (IE0031256328), while the ζ_E measure takes values from 2.7 percent (AT0000383864) to 55.5 percent (IE0031256328). Table 4 reports the results aggregated by issuing countries.

Tab. 4 Estimated price discovery measures for the EuroMTS trading platform

	Number of bonds	γ_E	ζ_E
ATS	10	0.1430	0.0988
BEL	8	0.1824	0.1803
ESP	10	0.2558	0.2446
FIN	5	0.1963	0.2149
FRF	14	0.1800	0.1623
GEM	12	0.2779	0.2664
GGB	10	0.2008	0.2624
IRL	4	0.5016	0.4773
MTS	13	0.0559	0.1799
NLD	6	0.2654	0.2114
PTE	6	0.1081	0.1114
Median	.	0.1742	0.1740
Mean	.	0.1966	0.2064
Std. error of mean	.	0.0132	0.0117

Note. The price discovery measures for the EuroMTS trading platform, γ_E and ζ_E, are defined by equation (5) and (6) of the paper, respectively. The values in the second and third numerical column are equally-weighted averages across bonds issued by the same country. Market codes are from Dufour and Skinner (2004). Computations based on 98 bivariate VEC models. See Section 3.3 of the paper for details on the selection of bonds chosen to compute EuroMTS market's contribution to price discovery.

Across the 98 bonds in our sample, the median of the two measures is the same (17.4 percent), with an average value slightly higher for ζ_E (20.6 percent) than the one for γ_E (19.7 percent). Based on the standard error of the mean values, these averages are significantly different from zero at the 1 percent level. This suggests that a non-negligible share of the efficient price determination occurs in the EuroMTS platform. Furthermore, a standard t-test for the equivalence of the mean (γ_E minus ζ_E) produces a test statistics equal to -0.56 with a p-value of 0.58, thus confirming that the estimated contribution is equivalent irrespective of which of the two price discovery measures is taken into account. As Figure 3 shows, most observations are substantially on the main diagonal of the scatter-plot (ζ_E versus γ_E) for the 98 bonds under investigation, implying that the two price discovery measures lead to non-conflicting conclusions. Finally, the correlation coefficient between γ_E and ζ_E turns out to be very high (0.81) and statistically significant at the 1 percent level.

3.5 Summary and discussion of results

The evidence here reported refers to items *A* (price convergence) and *B* (price discovery) as outlined in Section 2.3 above. We document that the LOP condition (2) holds for all pairs of bonds in the sample, suggesting that, despite the lack of explicit information linkage between European and domestic platforms, the architecture of the MTS system allows to eliminate persistent discrepancies between prices (Dufour and Skinner, 2004). Further, the formation of efficient prices seems to take place not only on domestic markets but also on the European platform, with trading on domestic MTS markets which dominates in an informative sense the orders executed on EuroMTS. This conclusion holds for all individual bonds (except for three bonds issued by the Irish government) and for all national markets (except for Ireland where γ_E turns out to be equal to 50.16 percent, as shown in Table 4). However, trades taking place on EuroMTS have a sizable informational content: we estimate its contribution to price discovery to be about 20 percent, on average. These findings are of practical interest for traders monitoring price developments in the European secondary market for benchmark Treasury bonds and have relevant institutional implications, since they can be interpreted as strong evidence against the “redundancy hypothesis”, in a way consistent with the conclusions in Cheung et al. (2005).

4 DETERMINANTS OF PRICE DISCOVERY IN THE EUROMTS

4.1 Observable market characteristics: Trading activity, price volatility and trading costs measures

In our multiple market setting, traders can access either market to obtain liquidity wherever and whenever it is cheapest. In keeping with previous works (Eun and Sabherwal, 2003; Chakravarty et al., 2004), markets' contributions to price discovery are likely to be systematically related to indicators of market functioning. It is generally understood that a well-functioning market should be characterised by *i*) high trading volumes, *ii*) low price volatility and *iii*) tiny bid/ask spreads.

As for point *i*), the measures of relative trading activity (ratio between EuroMTS and domestic MTS quantities) we consider are total trading volumes, $tvol$, and the number of trades, $ntra$. According to point *ii*), more liquid markets, with a continuous trade flow, are commonly characterized by small price variations. In contrast, less liquid markets, with extensive non-trading intervals, are likely to exhibit a higher return volatility. This suggests an inverse link between (relative) standard deviations of price changes, $rsig$, and the degree of contribution to price discovery. Finally, as to point *iii*), it is reasonable to expect an inverse relationship between γ_E and ζ_E and bid/ask spreads, since they constitute the largest part of trading costs. On the other hand, market makers may set wider spreads to shield against orders executed by more informed agents, thus inducing a positive relation. We construct two measures of spreads: the (relative) quoted bid/ask spreads associated with transactions, $qspr$, and effective spreads, $espr$, that is the (relative) difference between transaction prices and relative mid-points of the prevailing bid/ask quotes.

Appendix B illustrates how we extract trading activity, price volatility and transaction costs measures from (equally-weighted) daily averages over the sample span of reference as well as the other controls described below.

4.2 Institutional determinants: Market makers' obligations, maturity effects and regulatory practices in the primary market

As pointed out by Huang (2002), institutional arrangements may confound the linkage between observable market characteristics and price discovery. We control for differences in market making activity by specifying three dummy variables aimed at capturing the *a*) the number of continuous quoting hours, $hour$, *b*) the maximum spread that can be quoted, $mspr$, and *c*) the minimum quantity that dealers have to bid or to offer, $mqty$, on EuroMTS relative to their domestic MTS markets' counterparts. Since continuous quoting is costly for primary dealers, it reasonable to expect a negative (positive) relation between $hour$ ($mspr$) and γ_E or ζ_E. On the other hand, the "large trader's blessing" hypothesis (Scalia and Vacca, 1999), according to which the introduction of the anonymity in the MTS system has favoured the category of informed and/or large traders and, thus, the occurrence of larger transactions in size, suggests a positive relationship between $mqty$ and degree of price discovery on EuroMTS.

Other possible institutional factors influencing EuroMTS market's contribution to price discovery are closely tied to the effects of the ongoing integration in the primary and secondary markets for Treasury fixed income securities in Europe. A number of studies (Adam et al., 2002; Pagano and Von Thadden, 2004, among others) emphasise that the degree of financial integration in Europe appears to be inversely related to the level of risk-taking market participants are ready to assume and directly proportional to the degree of standardisation of the different financial instruments. Following Dunne et al. (2007), we distinguish short/medium term bonds (with maturity less than 6.5 years) from bonds with longer maturity (more than 6.6 years) through a dummy variable, $smty$, which is expected to be positively related to γ_E and ζ_E.

An additional effect of financial integration relates to the increased investors' interest on the characteristics of bond issues rather than on the nationality of issuers. Favero et al. (2000) point out that such a process induces euro area governments to compete each other for the same pool of funding, which translates into competition to obtain the services of primary dealers. In turn, auctioning government securities may involve risks for the issuer such as market squeezes, price manipulations, speculative behaviours, bidders' collusion or technical mistakes. Further, despite their similar architecture, domestic MTS and EuroMTS platforms may reflect different scopes of functioning, with the former aiming at satisfying issuer' liquidity needs within a regulated and efficient setting and the latter serving as a pure inter-dealer market (Girardi and Piga, 2007). So as to evaluate to what extent the euro area member states cover themselves from potential risks involved in auctioning government securities in the primary market, Bagella et al. (2006) indicate a group of five countries (Belgium, France, Germany, Ireland and the Netherlands) with a high protection against such risks, while Finland and Greece show a slightly lower degree of risks covering and the remaining countries (Austria, Italy, Portugal and Spain) a quite weak framework of rules. In view of that, a higher government's degree of protection against those risks, *high*, should lessen the need for the concentration of trading activity in the domestic MTS platform, suggesting a positive relation with γ_E and ζ_E.

4.3 Cross-sectional analysis: Tobit estimates

Since our dependent variables, γ_E and ζ_E, are restricted to lie between 0 and 1 by construction, we use a tobit estimator for censored variables. Table 5 provides the results for benchmark specifications (Panel [A]), which include only observable market characteristics in the set of regressors, and for specifications

Tab. 5 Determinants of price discovery on EuroMTS: Tobit estimates

	Panel [A]				Panel [B]			
	Model [1]		Model [2]		Model [3]		Model [4]	
	γ_E	ζ_E	γ_E	ζ_E	γ_E	ζ_E	γ_E	ζ_E
tvol	**0.4397**	**0.4943**	**0.5313**	**0.5771**	**0.4520**	**0.4902**	**0.5352**	**0.5622**
	(0.1353)	(0.1188)	(0.1549)	(0.1391)	(0.1473)	(0.1401)	(0.1713)	(0.1576)
rsig	**-0.8556**	**-0.7593**	**-0.7756**	**-0.7184**	**-0.9210**	**-0.7841**	**-0.8630**	**-0.7720**
	(0.2850)	(0.2630)	(0.2735)	(0.2597)	(0.2866)	(0.2519)	(0.2962)	(0.2394)
espr	**-1.2744**	-0.5533	.	.	-0.8461	-0.00748	.	.
	(0.6288)	(0.6452)	.	.	(0.6618)	(0.6906)	.	.
qspr	.	.	0.2723	0.3133	.	.	0.2759	0.3291
	.	.	(0.3842)	(0.3482)	.	.	(0.4331)	(0.3409)
smty	.	.	.	.	0.0324	**0.0455**	0.0369	**0.0465**
	.	.	.	.	(0.0238)	(0.0227)	(0.0246)	(0.0228)
high	.	.	.	.	**0.0893**	**0.1106**	**0.0936**	**0.1105**
	.	.	.	.	(0.0351)	(0.0298)	(0.0389)	(0.0340)
hour	.	.	.	.	**-0.0931**	**-0.0916**	**-0.1087**	**-0.0889**
	.	.	.	.	(0.0437)	(0.0354)	(0.0486)	(0.0336)
mqty	.	.	.	.	0.0139	0.0343	0.0227	0.0329
	.	.	.	.	(0.0256)	(0.0267)	(0.0265)	(0.0273)
Country dummies	YES	YES	YES	YES	YES	YES	YES	YES
σ	**0.0798**	**0.0745**	**0.0811**	**0.0743**	**0.0758**	**0.0687**	**0.0761**	**0.0681**
	(0.0058)	(0.0050)	(0.0060)	(0.0056)	(0.0059)	(0.0048)	(0.0059)	(0.0048)
LL	108.73	115.47	107.18	115.72	113.70	123.33	113.33	124.23
AIC	-189.46	-202.94	-186.37	-203.44	-191.41	-210.66	-190.65	-212.46
Pseudo-R^2	0.6247	0.5811	0.6127	0.5833	0.6610	0.6432	0.6583	0.6497

Note. The dependent variables are γ_E and ζ_E, alternatively. They are defined by equation (5) and (6) of the paper, respectively. The intercept term, albeit included among the regressors, is omitted for ease of exposition. Statistically significant coefficients according to the 95 percent level confidence interval, calculated using the bootstrap method with 500 replications, are in bold. Bootstrapped standard errors are reported in parenthesis. σ is the ancillary parameter of the tobit regression. LL and AIC indicate the value of the likelihood function and the Akaike Information Criteria, respectively. $Pseudo-R^2$ is the McKelvey-Zavoina (1975) measure of the goodness of fit of the regression. The number of observations is 98. Definitions of the regressors are provided in Appendix B.

augmented by other regressors described in Section 4.2 (Panel [B]), separately for γ_E and ζ_E. Model [1] (Model [3]) differs from Model [2] (Model [4]) with respect to the bid/ask spread used as explanatory variable. In all specifications, country dummies are included in order to control for other possible country-specific effects. The intercept term, albeit included among the regressors, is omitted for ease of exposition. Statistically significant coefficients at the 95 percent level confidence interval, calculated using the bootstrap method with 500 replications, are reported in bold. Bootstrapped standard errors are in parenthesis[17].

Consider the estimation results from Model [1] and Model [2], first. The $tvol$ coefficient is positive and statistically significant, implying that relatively larger trading volumes on EuroMTS are conducive to greater contribution to price discovery on the European trading platform. The $rsig$ coefficient is negative and statistically significant, in a way consistent with our economic priors. Finally, the relative spread term is not statistically significant in three out of four specifications. Thus, trading costs differentials between the EuroMTS and domestic MTS cannot be accounted as a major factor for choosing a trading platform rather the other, corroborating the conclusions in Cheung et al. (2005).

Following the recommendations in Veall and Zimmermann (1994), we use the McKelvey-Zavoina-Pseudo-R^2 as a measure of the goodness of fit for our regressions. The results are impressive: roughly 60 percent of the cross-sectional variation in γ_E and ζ_E can be explained by observable market characteristics alone. Comparing the above-discussed results to the estimates of specifications in Panel [B] several considerations emerge. *First*, EuroMTS market's contribution to price discovery is positively associated with trading volume and negatively related to volatility measures, even when controlling for institutional factors; by contrast, transaction costs are correctly signed but statistically not significant. *Second*, the magnitude of the estimated coefficients for observable market characteristics are very close to those obtained in the benchmark specifications. *Third*, institutional variables ($high$, $hour$, $smty$ and $mqty$) have the expected sign and are jointly significant at the 5 percent level according to a simple χ^2-distributed likelihood ratio test. *Fourth*, $high$ and $hour$

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[17] All estimation results refer to regression where the dummy $mspr$ is not included among the regressors, since that variable turns out to be not statistically significant in all specification. On the other hand, the magnitude and the statistical significance of coefficients for observable market characteristics remain unaffected by the inclusion of that variable, with the only exception of the coefficient of $hour$ which is estimated with lower precision, probably due to collinearity between $mspr$ and $hour$.

are strongly significant in all specifications, while the "maturity effect" can be detected only when ζ_E is the dependent variable. *Fifth*, goodness of fit statistics show that the augmented specifications are able to capture two-third of the overall cross-sectional variation of both price discovery measures[18].

4.4 Summary and discussion of results

The relationship between price discovery measures and indicators of market functioning (such as levels of trading activity, price volatility and trading costs) has been discussed extensively to address points *C* (price discovery and observable market characteristics) and *D* (price discovery and institutional factors) as presented in Section 2.3 above. Estimation results show that trades conveying information (in terms of contribution to price discovery) occur on EuroMTS when the level of trading activity is sufficiently high and the level of price volatility is sufficiently low. By contrast, trading cost differentials have a minor role in explaining market players' preferences in trading on the domestic platform rather the EuroMTS. When institutional factors are included as additional explanatory variables, the strong relationship between observable market characteristics and contribution to price discovery remains unaffected. The percentage of the cross-sectional variation in price discovery measures explained by the different specifications ranges from 58 to 66 percent.

In accordance with European authorities' principles behind the MiFID regime, favouring transparency is an essential mean to achieve an adequate price formation process. In turn, the relationship between transparency and price discovery is a complex one. On the one hand, the exposure of quotes forces market makers to be competitive, making it easier to find the best prices, especially for market takers, who are likely to be less sophisticated than larger market participants. On the other hand, order visibility may reduce the readiness of dealers, willing to keep large transactions confidential, to participate in the market. This may erode liquidity and impact the efficiency of price formation. Our results suggest that a proliferation of alternative trading platforms may be harmful in fostering integration of the European government bond market if the *potential* gains in competition across trading venues do not counterweight *certain* costs due to increased fragmentation in market liquidity.

[18] In order to extract as much information as possible from the tobit coefficients, we also consider their marginal effects, calculated at the mean, which provide a direct measure of the effect of the regressors on the dependent variable. Notice that the absence of censoring problems in our sample allows for an almost direct interpretation of the estimated coefficients as marginal effects. This is confirmed by a comparison of the coefficients of observable market characteristics from the tobit model in Table 5 and the marginal effects for the unconditional expected value of the dependent variable. For the sake of brevity, these regressions (available on request) are not reported.

Furthermore, a wider standardisation of longer-maturity issuances and of regulatory arrangements in the primary markets could be beneficial in improving the functiong of the European Treasury bond market.

5 ROBUSTNESS AND EXTENSIONS

We discuss the sensitiveness of our findings to modifications and extensions of the underlying empirical design with respect to *i*) the way of handling missing data for transaction prices series, *ii*) the computation of price discovery measures in the presence of wrongly signed feedback coefficients, *iii*) the econometric framework employed in explaining cross-sectional variability of price discovery measures, and *iv*) the inclusion of other possible regressors affecting price discovery variability across bonds.

Missing data. As discussed in Section 3.3 above, we use the “fill-in” method to overcome the problem of missing values. Such an approach does not affect the estimates of the long-run relationship equilibrium, but may influence the short-term information flow, since non-trading may produce a lower information share for the less frequent trading market even if the trades that take place do contain information (Lehmann, 2002). Since trades on EuroMTS are *fewer* than those occurring on the domestic trading venue for every pair of bonds involved in the analysis, the problem is less severe than it could appear. Our statistically significant estimates of EuroMTS market’s contribution to price discovery, indeed, can be interpreted as *lower* bounds. On the other hand, moving from the calendar-time (five days per week) to the transaction-time by synchronizing pairs of price series with respect to the less frequently observed variable can systematically discard important information if data are not missing completely at random (Little and Rubin, 1987).

Wrongly signed feedback parameters. In Section 3.3 we document that in six entities of references the γ_E statistics becomes negative and, thus, difficult to interpret. Following Blanco et al. (2005), we replace those negative numbers by zero. Summary statistics for γ_E and ζ_E computed for the entire sample (104 models) are quite similar to those reported in Table 4. The average values of γ_E and ζ_E (0.1853 and 0.2031, respectively) are statistically not different according to a standard t-test (p-value 0.31), with standard deviations of the mean substantially identical with respect to the values in Table 4, albeit the correlation between γ_E and ζ_E (0.78) turns out to be slightly lower than previously reported. Further, by comparing the mean value of γ_E (ζ_E) for the sub-sample of 98 bonds examined in Section 3.4 above and the entire sample of 104 entities, the t-test for the equivalence of the mean suggests not rejecting the null with a p-value of 0.54 (0.84).

Alternative specifications of the empirical framework. Three variants of the empirical setup discussed in Section 4.3 above are considered. First, the statistical significance of the estimated parameters is assessed by using standard errors calculated with the Huber-White sandwich estimator of variance (Huber, 1967; White, 1980) in place of those obtained from bootstrap techniques. The results are qualitatively similar, with bootstrapped confidence intervals slightly wider. Second, we take into account possible asymmetries by adding interaction terms between indicators of market functioning and $smty$ or $high$, alternatively. In none of regressions, we are able to detect statistically significant asymmetric effects. Third, as an additional check of robustness we replicate the estimation exercise presented in Table 5 using a standard linear regression model for logit tranformations of price discovery measures, ${}^{*}\gamma_E = \ln[\gamma_E/(1-\gamma_E)]$ and ${}^{*}\zeta_E = \ln[\zeta_E/(1-\zeta_E)]$, respectively. The results are presented in Table 6. Notice that the strong positive (negative) link between trading activity (price volatility) and price discovery measures is confirmed, giving support to our previous conclusions. Further, the explanatory power of the regressions is substantially similar, ranging from 49 to 61 percent. There are two main differences with respect to the estimates in Table 5: first, the maturity effect is statistically significant in all regressions; second, the $mqty$ variable turns out to be statistically significant in Model [3] and Model [4]. Particularly, the second finding seems to suggest the existence of possible informational asymmetries between uninformed dealers and traders who behave like informed investors (Fleming and Remolona, 1999) with their trades based on superior inventory and order flow information (Huang et al., 2002).

The interplay between primary and secondary government bond market: a reassessment. We turn on the above-discussed empirical evidence documenting a robust linkage between regulatory practices on primary market and the amount of price discovery taking place on EuroMTS (Table 5 and Table 6). Following the argumentations in Favero et al. (2000), it is reasonable to expect at least two other possible channels by which central government security issuances may influence the developments on the secondary bond market: namely, national gross issuances and the amount of outstanding public debt. In view of that line of reasoning, higher governments' levels of issuance should increase the need for liquidity in the domestic secondary Treasury bond market, leading to a negative relation with γ_E and ζ_E. Similar implications hold when outstanding debt levels are considered. Considering the national gross issuances during 2005, we distinguish large (more than 100 billion euro) from medium-small issuers, through a dummy variable $liss$. According to this criterion, France, Italy and Germany are to be considered large issuers. As far as the outstanding public debt level is concerned, we indicate Belgium, France, Italy and Germany as large debtors and the remaining countries as small debtors through the dummy variable $debt$.

Tab. 6 Determinants of price discovery on EuroMTS: OLS estimates

	Panel [A]				Panel [B]			
	Model [1]		Model [2]		Model [3]		Model [4]	
	$^*\gamma_E$	$^*\zeta_E$	$^*\gamma_E$	$^*\zeta_E$	$^*\gamma_E$	$^*\zeta_E$	$^*\gamma_E$	$^*\zeta_E$
tvol	**2.6722**	**3.0393**	**3.5955**	**3.5923**	**2.5310**	**2.9618**	**3.3169**	**3.4028**
	(1.0961)	(0.7499)	(1.4753)	(0.8875)	(0.9806)	(0.8161)	(1.2806)	(0.9046)
rsig	**-6.8009**	**-6.1305**	**-5.7250**	**-5.8945**	**-7.4250**	**-6.4518**	**-6.6761**	**-6.4774**
	(2.4410)	(1.8755)	(2.2597)	(1.8174)	(2.3402)	(1.7342)	(2.1377)	(1.6911)
espr	**-18.0010**	-2.9919	.	.	-11.8074	1.8428	.	.
	(8.1859)	(4.5802)	.	.	(6.6005)	(4.8557)	.	.
qspr	.	.	2.1652	2.1720	.	.	2.1297	2.2515
	.	.	(3.5485)	(2.1984)	.	.	(3.9118)	(1.8475)
smty	.	.	.	.	**0.3063**	**0.3306**	**0.3637**	**0.3291**
	.	.	.	.	(0.1594)	(0.1343)	(0.1784)	(0.1280)
high	.	.	.	.	**1.6794**	**1.1304**	**1.7403**	**1.1195**
	.	.	.	.	(0.6410)	(0.2749)	(0.6625)	(0.2851)
hour	.	.	.	.	**-1.3658**	**-0.9168**	**-1.5982**	**-0.8587**
	.	.	.	.	(0.6083)	(0.3077)	(0.6360)	(0.2957)
mqty	.	.	.	.	0.6918	**0.4164**	0.8233	**0.3838**
	.	.	.	.	(0.5015)	(0.1888)	(0.5673)	(0.2071)
Country dummies	YES	YES	YES	YES	YES	YES	YES	YES
LL	-109.97	-71.69	-114.11	-71.24	-99.12	-57.49	-100.94	-56.57
AIC	245.94	169.38	254.23	168.49	232.24	148.98	235.89	147.15
$AdjR^2$	0.4946	0.4998	0.4500	0.5043	0.5750	0.6071	0.5588	0.6144

Note. The dependent variables are the logistic transformation of γ_E and ζ_E, alternatively. They are constructed as $^*\gamma_E = \ln[\gamma_E/(1-\gamma_E)]$ and $^*\zeta_E = \ln[\zeta_E/(1-\zeta_E)]$. γ_E and ζ_E are defined by equation (5) and (6) of the paper, respectively. The intercept term, albeit included among the regressors, is omitted for ease of exposition. Statistically significant coefficients according to the 95 percent level confidence interval, calculated using the bootstrap method with 500 replications, are in bold. Bootstrapped standard errors are reported in parenthesis. LL and AIC indicate the value of the likelihood function and the Akaike Information Criteria, respectively. $AdjR^2$ is the square of the correlation between the predictor and the dependent variable, adjusted by the degrees of freedom. The number of observations is 98. Definitions of the regressors are provided in Appendix B.

We replicate the regressions in Table 5, with *i*) $liss$ and $debt$ as additional explanatory variable alternatively; *ii*) $liss$ or $debt$ in place of the indicator of governments' protection against risks in auctioning securities in the primary market, $high$. Overall, the relationship between observable market characteristics (trading activity levels, price volatility and trading costs) and price discovery measures holds even when controlling for these institutional factors. Particularly, when $liss$ and $debt$ enter as a further control, there is no additional contribution in explaining cross-sectional variability of price discovery measure (according to bootstrapped and asymptotically robust confidence intervals), with the other estimated coefficients substantially unchanged with respect to those reported in Table 5. This result seems to put forward no differenced patterns in price discovery revelation across markets between large and small issuers or between large and small debtors. By contrast, when $liss$ or $debt$ replaces $high$ as a regressor, we find weak statistical significance (at the 10 percent) only for the $liss$ coefficient in specification corresponding to Model [4] of Table 5 and when asymptotically robust standard errors are considered. On the one hand, these results reinforce the evidence of a possible influence of the primary market on EuroMTS market's contribution to price discovery mainly through regulatory practices in auctioning government securities. On the other hand, country-specific dummies in the regressions presented in Table 5 are likely to capture non-modelled institutional factors such as national gross issuances and the amount of outstanding public debt.

6 CONCLUDING REMARKS

Focusing on transaction data from the dominant trading platform for euro-denominated fixed income instruments, the MTS system, we provide a quantitative assessment of price discovery occurred across the domestic and the EuroMTS markets. We employ daily observations of benchmark government bonds issued by all euro area member countries' governments; a total of 107 bonds over the period January 2004 - March 2006. The extensiveness of our data sample allows us to explore not only the dynamic interactions between prices of trades occurring on the domestic MTS and the European platform, but also the cross-sectional variation in price discovery measures.

The paper reaches two main findings. *First*, we document that the architecture of the MTS system is able to eliminate persistent price discrepancies for the same bond traded on the two markets. The determination of the efficient price appears to take place with the contribution of trading activity on both platforms, with EuroMTS market's contribution to price discovery to be about 20 percent, on average. *Second*, estimation results reveal a systematic linkage between trading activity and price volatility and cross-sectional variability of price discovery taking place on the European platform. Trade cost differentials, instead, seem to have a minor role in explaining market players' preferences in trading government fixed income instruments on a platform rather the other. When institutional factors are included as additional explanatory variables, the strong relationship between observable market characteristics and EuroMTS market's contribution to price discovery remains unaffected. The robustness of these results is checked across a number alternative specifications. Aside from their scientific merit, these conclusions are of direct importance for investors trading government securities on the EuroMTS platform and have relevant implication for regulators attempting to identify conditions likely to promote further integration in the European financial system. In this respect, the proliferation of alternative platforms for trading European government securities may be harmful if potential benefits from competition do not counterweight costs due to the increased liquidity fragmentation across trading venues. A wider standardisation of longer-maturity issuances and of regulatory arrangements in the primary markets could be also beneficial in promoting a better functioning of the secondary Treasury bond market in Europe.

A fuller understanding of the relative importance of liquidity conditions and institutional factors is an empirical question that calls for further analysis. Possible improvements of the research agenda may include updating the sample span to monitor the effects of the ongoing harmonisation process in the European financial system on the developments of the secondary Treasury securities market. A second venue for further advances may take account a richer specification of the relationship between price discovery measures and their determinants across securities *and* over time, along the lines in Dufour and Nguyen (2007). In this respect, a closer scrutiny on the dynamics of market liquidity and trading activity indicators could be fruitful to increase market participants' confidence on trading securities on EuroMTS. These issues are left for future research.

APPENDIX A. LIST OF SELECTED GOVERNMENT BONDS.

The government bond markets covered in our dataset are those of Austria (ATS), Belgium (BEL), Spain (ESP), Finland (FIN), France (FRF), Germany (GEM), Greece (GGB), Ireland (IRL), Italy (MTS), the Netherlands (NLD) and Portugal (PTE). Market codes are from Dufour and Skinner (2004). For each country, we select all benchmark government securities traded in January 2004 with maturity date subsequent the end of our estimation horizon (March 2006). 107 bonds satisfy such a requirement. Bond codes are reported below.

ATS	BEL	ESP	FIN	FRF	GEM	GGB	IRL	MTS	NLD	PTM
AT0000383518	BE0000286923	ES0000012239	FI0001004822	FR0000187361	DE0001135176	GR0110014165	IE0006857530	IT0001448619	NL0000102101	PTOTECOE0011
AT0000383864	BE0000291972	ES0000012387	FI0001005167	FR0000187635	DE0001135192	GR0114012371	IE0031256211	IT0003080402	NL0000102317	PTOTEGOE0009
AT0000384227	BE0000296054	ES0000012411	FI0001005332	FR0000187874	DE0001135200	GR0114015408	IE0031256328	IT0003171946	NL0000102606	PTOTEJOE0006
AT0000384821	BE0000297060	ES0000012445	FI0001005407	FR0000188328	DE0001135218	GR0124006405	IE0032584868	IT0003190912	NL0000102671	PTOTEKOE0003
AT0000384938	BE0000298076	ES0000012452	FI0001005514	FR0000188690	DE0001135226	GR0124011454	.	IT0003242747	NL0000102689	PTOTEWOE0009
AT0000384953	BE0000300096	ES0000012783	FI0001005522	FR0000188989	DE0001135234	GR0124015497	.	IT0003256820	NL0000102697	PTOTEXOE0016
AT0000385067	BE0000301102	ES0000012791	.	FR0000189151	DE0001135242	GR0124018525	.	IT0003271019	.	.
AT0000385356	BE0000302118	ES0000012825	.	FR0010011130	DE0001141380	GR0124021552	.	IT0003357982	.	.
AT0000385745	BE0000303124	ES0000012866	.	FR0103230423	DE0001141398	GR0124024580	.	IT0003413892	.	.
AT0000385992	.	ES0000012882	.	FR0103840098	DE0001141406	GR0128002590	.	IT0003472336	.	.
.	.	.	.	FR0104446556	DE0001141414	GR0133001140	.	IT0003477111	.	.
.	.	.	.	FR0105427795	DE0001141422	GR0133002155	.	IT0003493258	.	.
.	.	.	.	FR0105760112	DE0001141430	.	.	IT0003522254	.	.
.	.	.	.	FR0106589437	.	.	.	IT0003532097	.	.
.	.	.	.	.	.	.	.	IT0003535157	.	.
.	.	.	.	.	.	.	.	IT0003611156	.	.
.	.	.	.	.	.	.	.	IT0003618383	.	.

APPENDIX B. CONSTRUCTION OF VARIABLES.

B.1 Observable market characteristics

Let x^j be the (equally-weighted) daily average of a variable x over the sample span of reference, where $j = E, D$ indexes the EuroMTS (E) or the domestic MTS (D) platform, respectively. The following log-transformations are performed: $tvol = \ln[1 + (vol^E / vol^D)]$, where vol is the nominal amount of trades in million euro; $ntra = \ln[1 + (tra^E / tra^D)]$, where tra is the ratio between vol and the average size of trades in million euro; $rsig = \ln[1 + (sig^E / sig^D)]$, where sig is the standard deviation of the first differenced logarithms of transaction prices (Δp^j); $qspr = \ln[1 + (qsp^E - qsp^D)]$, where qsp is the quoted bid/ask spread associated with the transaction; $espr = \ln[1 + (esp^E - esp^D)]$, where esp is the difference between transaction prices and the mid-point of the prevailing bid/ask quote.

B.2 Institutional variables

These regressors are binary variables: $smty$ is a dummy taking value 1, if bonds have a maturity (in terms of the difference between the maturity date and the issue date) less than 6.5 years, and 0, otherwise; $high$ is a dummy taking value 0, if countries have a high overall auction risks covering degree, and 1, otherwise; $hour$ is a dummy taking value 1, if the number of quoting hours for a bond on EuroMTS is higher than on the domestic MTS, and 0, otherwise; $mspr$ is a dummy taking value 1, if the maximum bid/ask spread for a bond on EuroMTS is lower than the one on the domestic MTS, and 0, otherwise; $mqty$ is a dummy taking value 1, if the minimum quantity for a bond on EuroMTS is higher than the one on the domestic MTS and 0, otherwise.

B.3 Other controls

The third class of regressors includes eleven country dummies (taking value 1, when the bond is issued by the Treasury of that country, and 0, otherwise); the distinction between large and small issuers (borrowers) by means of $liss$ ($debt$), a dummy taking value 1, if bonds are from Italy, Germany or France (or Belgium), and 0, otherwise.

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