Valuation, Liquidity and Risk in Government Bond Markets*

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Abstract

We explore the determinants of yield differentials between sovereign bonds in the Euro area. There is a common trend in yield differentials, which is correlated with a measure of the international risk factor. In contrast, liquidity differentials display sizeable heterogeneity and no common factor. We estimate a model that predicts that yield differentials should increase in both liquidity and risk, with an interaction term whose magnitude and sign depend on the size of the liquidity differential with respect to the reference country. Testing these predictions on daily data, we find that the international risk factor is consistently priced, while liquidity differentials are priced for a subset of countries and their interaction with the risk factor is crucial to detect their effect.

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1 Introduction

As soon as the European Monetary Union (EMU) took place in 1999, an integrated market for fixed-income securities came to life in the Euro-area. EMU eliminated currency risk within this area, and standardization of bond conventions by Euro-area sovereign issuers made public bonds more easily comparable. As a result, the public debt securities issued by different Euro-area governments became very close substitutes: yield spreads on Euro-area government bonds converged significantly, narrowing from highs in excess of 300 basis points, for certain maturities, to less than 30 basis points across the maturity spectrum over the course of 1997-98. Yet, despite such convergence, euro-zone government bonds are still not regarded as perfect substitutes by market participants: non-negligible differences in yield levels across countries have remained, to different extents for different issuers and maturities, and they fluctuate over time without a clearly discernible trend. Even the bonds issued by the highest-rated issuers are not regarded as perfect substitutes of each other, so that for example French bonds traded in the cash market are not considered as a perfect hedge for positions in Bund futures.[1] What is the reason for these persistent differentials? One possible explanation is persistent risk differences. Different sovereign issuers are perceived as having different solvency risks, in spite of the provisions of the Stability Pact. A second possible explanation is liquidity. This is indeed the explanation that is often advanced by the financial press. But a look at the time-series behavior of Euro-area yield differentials suggests that neither one of these two factors in isolation is likely to provide the full answer. First, as shown below, the yield differentials relative to the German Bund tend to fluctuate together, much more than measures of liquidity (bid-ask spreads) do. This suggests that liquidity alone cannot be the full answer, and that there must be another common factor driving the differentials' time-series behavior. But this factor can hardly be the solvency of individual issuers, which is unlikely to change sharply over time and to correlate strongly across issuers. It might instead the return to an international risk factor, which can change sharply over time. For instance, even if the default risk of the Italian and French governments relative to the German one were very stable over time, a changing world price for risk could induce the implied yield differentials to correlate over

[1] See Pagano and von Thadden (2004) for an account of the integration of European bond markets and for a survey of the relevant literature.

time. However, this cannot be the full story either. Sizable yield differentials have been observed for several years even within the group of AAA-rated euro-zone countries, even though they have generally narrowed considerably over time. Still, as late as 2002, 10-year AAA-rated Finnish debt yielded on average 20 basis points more than the 10-year German Bund. This suggests that indeed liquidity differences may play a role, as practitioners claim.[2]

As a possible explanation of this phenomenon we develop a simple asset-pricing model with risk neutral investors and exogenous transactions costs. In the context of this model, we show how the risk factors interact with the changing world price of risk, and thus affect the level and time-series behavior of yield differentials. The most important insight of the theoretical analysis is that liquidity matters for pricing, but that it interacts with fundamental risk. In particular the impact of the world price of risk on yields is dampened if a market becomes more liquid when the international price of risk is high This implies that a direct estimation of the impact of liquidity on prices, i.e. an estimation that ignores the indirect effect caused by the interaction with world-wide risk, is likely to underestimate the impact of liquidity. Our main contribution is then to bring these ideas to the data using two years of daily observations on yields and liquidity variables for Euro-area sovereign bonds at the 5-year and 10-year maturities. The results show that a proxy for the world price of risk – the difference between high-risk U.S. corporate bonds and U.S. government bonds at the corresponding maturity – is the single most important explanatory variable for Euro-area yield differentials. Liquidity differentials – as proxied by the difference between the local and the relevant reference bid-ask spread – play a role only in a subset of countries. Whenever it appears with a statistically significant coefficient, the bid-ask spread impacts positively the corresponding yield relative to that of the benchmark, and its interaction with the world risk factor is negative

[2]For instance, the increase of yield differentials relative to the Bund rate in late 1999 was explained as follows: "after having tested the waters of Europe's smaller bond markets, institutional investors are deciding they've had enough ... declining liquidity in the smaller debt markets is boosting the premiums these countries are having to pay investors compared with the core euro-zone nations" (Wall Street Journal Europe, November 3, 1999). Market practitioners clearly attribute remaining yield differentials to liquidity premia, which are held to be larger in thinner markets, irrespective of their credit rating: "'Peripheral issuers in Europe are in trouble: They're paying a huge liquidity premium' says Steven Mayor, chief bond strategist at ING Barings in London. He says that their problem comes down to the fact that some still only represent 1% to 2% of the euro-zone issuance" (ibidem).

and precisely estimated. In other words, (i) illiquidity appears to command a premium, as in most of the literature following Amihud and Mendelson (1986), and (ii) the size of such premium depends on the covaration between the cost of illiquidity and the world price of risk. The structure of the remainder of this paper is as follows. Section 2 sets the paper in the context of the relevant literature. Section 3 presents the data and describes the stylized facts that emerge from them. Section 4 lays out the model and its predictions. Section 5 presents and discusses the estimation results. Section 6 concludes.

2 Related literature

This paper adds to a considerable literature on the relation between returns and liquidity. At a theoretical level, two main views have been advanced to explain why liquidity should be priced by financial markets: illiquidity (i) creates trading costs, and (ii) can itself create additional risk. These views are not mutually exclusive, although they have emerged sequentially in the literature. This paper builds on the first view and develops it in a new direction, which is similar in spirit to that of the second view. The "trading cost view" holds that illiquid securities must provide investors with a higher expected return to compensate them for their larger transaction costs, controlling for fundamental risk. The prediction here is a cross-sectional one: risk-adjusted expected return must be higher for less liquid securities. This view, first proposed and tested by Amihud and Mendelson (1986), has been the basis of a vast empirical literature. Many subsequent studies of stock-market data have confirmed a significant cross-sectional association between liquidity (as measured by the tightness of the bid-ask spread or trading volume) and asset returns, controlling for risk: among these are Brennan and Subrahmanyam (1996), Chordia, Roll and Subrahmanyam (2000), Datar, Naik and Radcliffe (1998), and Eleswarapu (1997). Other studies have focussed on liquidity effects in fixed-income security markets. Here, too, the initiators were Amihud and Mendelson (1991), who showed that the yield to maturity of treasury notes with six months or less to maturity exceeds the yield to maturity on the more liquid treasury bills. Other studies on U.S. public debt securities by Warga (1992), Daves and Ehrhardt (1993), Kamara (1994) and Krishnamurthy (2000) confirmed these findings, although using more recent data Strebulaev (2001) found that the yield spread between bills and matched notes is much smaller than previously found, especially when bills are on-the-run. Recently, Goldreich, Hanke and Nath (2002) investigated the impact of expected liquidity on securities' prices. They analyze the prices of Treasury securities as their liquidity changes predictably, in the transition from on-the-run to the less liquid off-the-run status, and show that the liquidity premium depends on the expected future liquidity over their remaining lifetime rather than on their current liquidity. The "liquidity risk view" highlights that liquidity is priced not only because it creates trading costs, but also because it is itself a source of risk, since it changes unpredictably over time. Since investors care about returns net of trading costs, the variability of trading costs affects the risk of a security. Acharya and Pedersen (2004) show in a CAPM framework with overlapping generations of investors that liquidity risk should be priced to the extent that it is correlated across assets and with asset fundamentals, and uncover evidence consistent with this prediction. Similarly, Ellul and Pagano (2004) show that the initial underpricing of IPO shares should compensate investors also for the expected illiquidity and for the liquidity risk that investors face in aftermarket trading, and not only for the fundamental risk and adverse selection problems they are exposed to. Also Gallmeyer, Hollifield and Seppi (2004) propose a model of liquidity risk where traders have asymmetric knowledge about future liquidity, so that less informed investors try to learn from the amount of current trading volume how much liquidity there may be in the future. They show that current liquidity is a predictor of future liquidity risk, and therefore is priced. Our paper puts forward what may be labeled the "risk-liquidity interaction view" and points out that liquidity alters the impact of changes in risk on current prices and yields. So here the emphasis is not on liquidity risk (indeed in this approach future liquidity is perfectly anticipated), but rather on the interaction between liquidity and fundamental risk. In the model presented in this paper, changes in fundamental risk are shown to affect less the price of bonds that are *currently* less liquid, but more the prices of bonds that are *expected* to be less liquid in the future. This prediction is consistent with the empirical findings of Goldreich, Hanke and Nath (2002). The second result parallels that in the model by Vayanos (2004), where fund managers are subject to withdrawals when their performance falls below a given threshold, and therefore are more likely to liquidate at times of high volatility. This increases the liquidity premium at times of high volatility. So in both models increased risk generates a flight to liquidity. On technical grounds, our three-date analysis is much simpler than that in Vayanos' continuous-time dynamic equilibrium model with stochastic volatility, and in fact more akin to that by Gallmeyer, Hollifield and Seppi (2004), who, like us, rely on a Diamond-Dybvig framework to motivate liquidity trading. (Instead, we share with Vayanos (2004) the modeling of illiquidity as an exogenous transaction cost.) On the empirical front, our analysis adds to a small recent literature on Euro-area yield differentials. Codogno, Favero and Missale (2003) estimate models of Euro-area differentials with both monthly and daily data. Their estimates of monthly data show that for most countries only international risk factors, and not domestic ones, have explanatory power (the former being proxied by U.S. bond yield spreads and the latter by national debt/GDP ratios). In their estimates of daily data (that refer to 2002 only), macroeconomic variables are not included because they move too slowly to allow the estimation of the impact of the domestic risk factor. Again, the international factor is statistically significantly for most countries, while liquidity (as measured by trading volume) is significantly and positively correlated with spreads for France, Greece, the Netherlands and Spain. Geyer, Kossmeyer and Pichler (2004) estimate with weekly data a multi-issuer state-space version of the Cox-Ingersoll-Roll (1985) model of bond yield spreads (over Germany) for four EMU countries (Austria, Belgium, Italy, and Spain). They find that idiosyncratic country factors have almost no explanatory power, and yield-spread data reflect mainly a single ("global") factor, whose variation can, to a limited extent, be explained by EMU corporate bond risk (as measured by the spread of EMU corporate bonds over the Bund yield), but by nothing else – in particular not by measures of liquidity. Their measurement of liquidity variables is, however, at best indirect, as they do not use data on bid-ask spreads, but rather derived measures of liquidity, such as issue size and the yield differential between on-the-run and off-the-run bonds. Despite the considerable differences in the methodology and data used, these two studies agree on the finding that yield differentials under EMU are driven mainly by a common risk (default) factor, related to the spread of corporate debt over government debt, and suggest that liquidity differences have at best a minor role in the time-series behavior of yield spreads. As we shall see, our results, which rely on a more direct measure of liquidity (daily bid-ask spreads), confirm the former result but also highlight that the effect of liquidity cannot be properly gauged without taking into account its interaction with changes in the common risk factor. Interestingly, the interaction between liquidity and risk appears to be price-relevant also in the European treasury bill market:

Biais, Renucci and Saint-Paul (2004) document that, when volatility is high, yields are lower for bills with a larger outstanding supply, which are likely to be the most liquid.

3 Data and stylized facts

The data that we use in the empirical analysis concern benchmark bonds' prices and liquidity indicators for the Euro area, observed at daily frequencies for the period from 1 January 2002 to 23 December 2003. The data are collected from the Euro MTS Group's European Benchmark Market trading platform, and refer to a snapshot taken at 11 a.m. (Central European Time) in all market days for the Telematico cash markets. The database contains: (i) the best five bid and ask prices across all markets, (ii) the aggregate quantity of all the outstanding proposals made at the best bid and best ask prices, and (iii) the daily trading volume of each bond on the EBM. From these data we calculate redemption yields, maturities and a range of liquidity-related variables described in the Appendix. We consider Austria, Belgium, Finland, France, Germany, Italy, the Netherlands, Portugal and Spain. We do not include Greece and Ireland in the sample, because in 2002 the convergence process to EMU was still ongoing for Greece, while the Euro MTS data for Ireland become available only at a very late stage of our sample. Table 1 provides descriptive statistics for the yield differentials relative to Germany and the bid-ask spreads by country. For 10-year benchmark bonds, average yield differentials range from 4.16 and 6.94 basis points for France and the Netherlands to 14.47 and 15.50 basis points for Italy and Portugal respectively, while the range of variation is smaller for 5-year bonds. In both cases, the standard deviation indicates that yield differentials feature considerable time-series variability. The statistics reported in the lower panel indicate that bid-ask spreads are all very tight and stable over time. For 10-year benchmark bonds, average bid-ask spreads range from 2.52 and 2.86 ticks for Italy and France to 4.60 and 4.87 for Austria and Finland, respectively. German Bunds are the third most liquid bonds after Italian and French ones in the cash market, with a spread of 3.25 ticks. The situation is similar for 5-year bonds. Figure 1 illustrates the time variation of 10 year yield differentials between each country in our sample and Germany, taken as the reference country. For clarity, we report separately the data for the Netherlands, France and Austria in the upper panel, and for all the remaining countries in the lower panel of the Figure. Yield differentials have a clear tendency to comove. The presence of comovement is confirmed by Table 2, which reports the correlation between yield differentials over the sample period and presents a principal-components analysis. Correlations are very high both within and between groups, and the principal-components analysis shows that the first principal component explains above 90 percent of the variance of the series. Liquidity indicators behave differently. Figure 2 shows the difference in bid-ask spreads observed for benchmark bonds relative to German ones, for the same groupings of countries as those used in Figure 1. The figure reports five-days moving averages of the daily observations to smooth volatility. Clearly, liquidity indicators have a different time pattern from yield differentials. This is confirmed by the correlations and principal-components analysis shown in Table 2. The correlation between differentials in liquidity indicators is much lower than that between yields differentials. Moreover, the principal- components analysis reveals that for liquidity indicators at least six components are needed to explain the same proportion of the total variance as that explained by the first component in the case of yield differentials. The principal-components analysis of Table 2 shows clearly that there is a common international factor in yield differentials in Europe. In Figure 3 we display the behavior of a variable that is often proposed in the literature as a proxy for this factor: the spread between the yield on 10-year fixed interest rates on swaps and the yield on 10-year US government bonds. There is ample evidence of a common trend in international bond spreads (see, for example, Dungey et al.1997). The empirical literature on sovereign bond spreads in emerging markets shows that the yield of US government bonds, the slope of the US yield curve and risk indicators on the US bond markets, are the main determinants of sovereign spreads (see, for example, Eichengreen and Mody, 2000; Barnes and Cline, 1997, and Kamin and Von Kleist, 1999, Arora and Cerisola, 2001). Blanco (2001) and Codogno et al. (2003) use proxies for global credit risk derived from the US yield curve in their models of euro-zone government security yields. Consistent with these findings and with the results of Geyer, Kossmeier and Pichler (2004), Figure 3 shows that this international risk factor is strongly correlated with the first principal component of yield differentials in the Euro area.

4 Theoretical Framework

We consider a partial-equilibrium model with three dates $t = 0, 1, 2$. There are two traded bonds, denoted by A and B, and a riskless asset that yields a net return of r per period. A continuum of investors $h \in [0, 1]$ are ready to invest at date 0 in order to consume at date 2. Investors are risk neutral and maximize date-2 consumption.

Bond $i = A, B$ pays its face value V with probability q_i and 0 with probability $1 - q_i$ at date 2 and nothing at date 1. Without loss of generality we assume that $q_B \geq q_A$ and interpret bond A as the benchmark. Its purchase price at date zero is p_{0i}. The bonds can be re-traded at date 1 at a bid price of $(1 - t_i)p_{1i}$, and ask price of p_{1i}, where t_i is a proportional transactions cost. The reason why investors may want to liquidate their bonds at this interim date is to catch an alternative investment oppportunity at date 1. This investment pays zero with probability $a^h\rho$ and a gross return z with probability $1 - a^h\rho$, so that its expected return is

$$z(1 - a^h\rho),$$

where $a^h \in [0, A]$ is an investor-specific parameter that summarizes investment properties such as its exposure to aggregate risk, and ρ is an aggregate risk factor that impacts all private investment. Since this affects all investors alike and since we will apply the model to international investors, we also call this the international risk factor, in line with the empirical literature. As the risk sensitivity a^h increases over its support, the expected return of h's investment opportunity falls from z to 0. We assume that a^h satisfies the law of large numbers, so that uncertainty washes out in the aggregate and individual probabilities are the same as aggregate frequencies. Let us denote by G the differentiable cumulative distribution function of a^h and by $g = G'$ its density function.

At date 1, an investor will liquidate his holdings of bond i to invest in his outside investment opportunity if and only if the latter's expected profitability exceeds the expected return of the bond over its residual life, net of transactions costs:

$$z(1 - a^h\rho) > \frac{(1 - q_i)V}{(1 - t_i)p_{1i}}. \tag{1}$$

Hence, investor h will sell his holdings of bond i with probability

$$\pi_i = \text{prob}\left(a^h < \frac{1}{\rho} - \frac{(1-q_i)V}{\rho z(1-t_i)p_{1i}}\right) \tag{2}$$

$$= G\left(\frac{1}{\rho} - \frac{(1-q_i)V}{\rho z(1-t_i)p_{1i}}\right) \tag{3}$$

Note that here investors liquidate their assets not in response to an exogenous consumption shock as in models of the type of Diamond and Dybvig (1983), but in response to their changing investment opportunities, and therefore as a function of market prices and yields. In other words, their demand for liquidity is price-elastic, in contrast with what happens in the Diamond-Dybvig setting. At the interim date, investors will hold (or buy) bond i only if (1) does not hold and if the bond does not yield less than the safe investment, i.e. if its price is sufficiently low:

$$p_{1i} \leq \frac{(1-q_i)V}{1+r} \equiv R_i \tag{4}$$

Hence, $p_{1i} = R_i$ is the maximum possible price for an equilibrium at date 1 to exist. Necessary for the existence of an equilibrium is therefore that aggregate demand at this price is positive, i.e. that:

$$G\left(\frac{1}{\rho} - \frac{1+r}{\rho z(1-t_i)}\right) > 0. \tag{5}$$

As a result, equilibrium requires also that the argument of (5) be positive, that is, $z(1-t_i) > 1+r$. This is equivalent to requiring that at least the investor with the best investment opportunity will be willing to liquidate bond i, since the proceeds of the investment z, net of the proportional liquidation cost t_i, exceeds the opportunity cost of capital $1+r$. If this condition were not met, there would be no demand for (and supply of) liquidity at the interim stage. The bond market would be inactive at date 1.

Under this assumption, there are two possible cases, each with a unique equilibrium: (4) holds either with equality or with strict inequality. In the latter case, the amount of cash available for buying bonds at date 1 is so small that arbitrage fails and the bond trades at a discount. We will concentrate on the former, more realistic case, in which there is sufficient demand for bonds such that the expected yields of all assets are equalized. In this case we have

$$\begin{aligned} p_{1i} &= R_i,\ i = A, B &(6)\\ \pi_i &= G\left(\frac{1}{\rho} - \frac{1+r}{\rho z(1-t_i)}\right) &(7)\end{aligned}$$

In particular, the discount rate between dates 1 and 2 is given by r as between 0 and 1. The expected payoff of bond i at date 0 therefore is

$$p_{0i} = \frac{\pi_i(1-t_i)p_{1i}}{1+r} + \frac{(1-\pi_i)(1-q_i)V}{(1+r)^2} = \frac{(1-\pi_i t_i)(1-q_i)V}{(1+r)^2}$$

So the bond's pledged yield to maturity is

$$1 + Y_i = \frac{V}{p_{0i}} = \frac{(1+r)^2}{(1-\pi_i t_i)(1-q_i)}$$

The yield ratio between the two bonds is simpler to calculate than the yield differential and given by

$$\frac{1+Y_B}{1+Y_A} = \frac{(1-\pi_A t_A)(1-q_A)}{(1-\pi_B t_B)(1-q_B)}$$

By using the approximation $\ln(1+x) \approx x$ the yield differential at date 0 can be approximated simply by:

$$\Delta Y = Y_B - Y_A \approx \pi_B t_B + q_B - \pi_A t_A - q_A \qquad (8)$$

This expression for the yield differential is intuitive. First, there is a direct positive impact of transactions costs. If π_i were constant, i.e. the probability of liquidation were not endogenously affected by market data, then transaction costs would, as usual, drive up yields – or, equivalently, greater bond liquidity would be associated with lower required yields. This is clear from (8): if π_i is constant and the cost t_B of trading bond B increases, its yield increases relative to that of bond A. The reason is that the buyer of the asset anticipates trading costs that must be compensated. As these costs only materialize if the holder trades at date 1, t_B is weighed with the probability of this event, which is π_B. The second standard feature of (8) is that the yield differential increases in fundamental risk: the higher the risk of default of, say, bond B, the higher its required yield compared to bond A.

Given the absence of risk aversion in our model, this is not a risk premium but simply reflects the reduction in the discounted expected future value of the bond.

However, the response of bond yields to changes in liquidity and risk become more interesting once one takes into account that the probability of liquidation is endogenous in this model. This probability itself reacts to changes in both transaction costs (the terms at which the "supply of liquidity" is available) and outside investment opportunities (that determine the "demand for liquidity" by market participants). Indeed, as we shall see, it also depends non-trivially on the interaction between the two.

As one would expect, both higher transaction costs (t_i) and riskier outside investment opportunities (ρ) reduce a given bond's probability of liquidation (π_i):

$$\frac{\partial \pi_i}{\partial t_i} = -\frac{1+r}{\rho z(1-t_i)^2} g\left(\frac{1}{\rho} - \frac{1+r}{\rho z(1-t_i)}\right) < 0, \tag{9}$$

$$\frac{\partial \pi_i}{\partial \rho} = -\frac{1}{\rho^2}\left(1 - \frac{1+r}{z(1-t_i)}\right) g\left(\frac{1}{\rho} - \frac{1+r}{\rho z(1-t_i)}\right) < 0. \tag{10}$$

Intuitively, if transactions costs increase, i.e. the bond becomes less liquid, then liquidating it becomes less attractive, hence the probability of selling decreases; similarly, if the aggregate risk factor increases, the investors' market investment opportunities become less attractive, which again decreases their probability of selling the bond.

The joint effect of an increase in transaction costs and in aggregate risk, however, is less easy to sign. The relevant cross-derivative measures how the trading probability's sensitivity to international risk depends on the bond's liquidity:

$$\frac{\partial^2 \pi_i}{\partial \rho \partial t_i} = \frac{1}{z^2 \rho^3} \frac{1+r}{(1-t_i)^2}\left(\rho z g(\frac{1}{\rho} - \frac{1+r}{\rho z(1-t_i)}) - (\frac{1+r}{1-t_i} - z) g'(\frac{1}{\rho} - \frac{1+r}{\rho z(1-t_i)})\right). \tag{11}$$

From equation (10), we know that π_i decreases with aggregate risk. What equation (11) tells us is whether this effect is larger or smaller depending on the bond's transaction costs (or depending on whether the bond is simultaneously become more expensive to trade). If the derivative is positive, there is dampening: the effect of greater risk is weaker for bonds with higher transactions costs. If it is negative, there is amplification: illiquidity increases the effect of changes in international risk. The cross derivative cannot be signed in general, but it is positive in several simple examples (for example, it is trivially positive when the distribution of the a^h is uniform).

Equipped with these results, we can now explore how risk and liquidity affect bond yields with endogenous liquidation probabilities. For simplicity, we perform the comparative statics of the yield differential (8) with respect to t_B only (the effect of t_A is analogous):

$$\frac{\partial \Delta Y}{\partial t_B} = \pi_B + \frac{\partial \pi_B}{\partial t_B} t_B, \tag{12}$$

$$\frac{\partial \Delta Y}{\partial \rho} = \frac{\partial \pi_B}{\partial \rho} t_B - \frac{\partial \pi_A}{\partial \rho} t_A + \frac{\partial (q_B - q_A)}{\partial \rho}, \tag{13}$$

$$\frac{\partial^2 \Delta Y}{\partial t_B \partial \rho} = \frac{\partial \pi_B}{\partial \rho} + \frac{\partial^2 \pi_B}{\partial t_B \partial \rho} t_B. \tag{14}$$

The first term in (12) is positive, the second negative. Hence, the sign is a priori ambiguous. This ambiguity reflects two opposing effects – a direct and an indirect one. The direct effect of higher transactions costs is to drive up the price by a factor π_B: with exogenous liquidation probability, this would be the only effect, as observed above in discussing (8). But when liquidation is endogenously determined, this effect is at least partly counteracted by a reduction of the trading probability, as described in (9). As this probability decreases, the relative price of the bond increases, because liquidation costs are incurred less often. Hence, this effect is proportional to t_B. To the extent that empirically bid-ask spreads in the bond market are very small, we expect the latter effect to be small in absolute terms and the overall effect to be positive.

The impact of risk on yield differentials in (13) has first two terms of similar size and opposite sign, so that one would expect its overall sign to depend mainly on the sign of the third term $\partial(q_B - q_A)/\partial\rho$, which measures the impact of international risk on the fundamental risk of the two bonds. Two possibilities exist in principle. First, the effect of aggregate risk is the same for both bonds, in which case the term is zero. Alternatively, the riskier bond is more sensitive to international risk than the safer one, in which case we have $\partial(q_B - q_A)/\partial\rho > 0$. If this effect is sufficiently large to make the overall sign of $\partial\Delta Y/\partial\rho$ in (13) positive, we have the well-known "flight to quality": an increase in international risk makes the safer bond more attractive than the riskier one and hence drives up the yield differential. This is the typical effect stressed by practitioners.

Finally, the joint effect of changes in risk and in liquidity is in general ambiguous. This is not surprising, since the cross derivative (14) depends entirely on the derivatives of π_B. As seen above, while the direct effect of international risk on the probability of liquidation – the first term in (14) – is always negative, the second term is ambiguous. However, as the second term is proportional to t_B, which is small, we expect the negative first term to dominate.

To illustrate these considerations, we calculate the above derivatives for the case in which the individual characteristics a^h are uniformly distributed on the interval $[0, A]$. In this case we have

$$\pi_i = \frac{1}{A\rho z}\left(z - \frac{1+r}{1-t_i}\right)$$

and

$$\frac{\partial \Delta Y}{\partial t_B} = -\frac{1+r}{z(1-t_B)^2}t_B + 1 - \frac{1+r}{z(1-t_B)} = \frac{1}{A\rho z}\left(z - \frac{1+r}{(1-t_B)^2}\right),$$

which is positive if t_B is small (remember that $z > (1+r)/(1-t_B)$ by assumption (5). Furthermore, (13) becomes

$$\begin{aligned}\frac{\partial \Delta Y}{\partial \rho} &= \frac{1}{\rho}\left(\pi_A t_A - \pi_B t_B\right) + \frac{\partial (q_B - q_A)}{\partial \rho} \\ &= \frac{t_B - t_A}{A\rho^2 z}\left(\frac{1+r}{(1-t_A)(1-t_B)} - z\right) + \frac{\partial (q_B - q_A)}{\partial \rho}\end{aligned}$$

Since $t_B - t_A \approx 0$ this expression is dominated by the impact of aggregate risk on the relative default risks of the two bonds. Finally, the cross effect is

$$\frac{\partial^2 \Delta Y}{\partial t_B \partial \rho} = \frac{1}{A\rho^2 z}\left(\frac{1+r}{(1-t_B)^2} - z\right)$$

which has the opposite sign of $\partial \Delta Y / \partial t_B$, hence is negative if t_B is small.

If we remember that our variations of t_B are a mirror image of those of t_A and therefore describe those of the transactions cost (illiquidity) differential Δt, our general discussion can be summarized as follows.

Proposition: *The yield differential of the two bonds depends positively on their transactions cost differential. It is insensitive to aggregate risk, if aggregate risk affects bonds of different fundamental value identically; it depends* *positively on aggregate risk, if the fundamentals of riskier bonds react more strongly to aggregate risk than those of less risky ones. The positive effect of transactions costs on the yield differential is dampened by aggregate risk. In short,*

$$\frac{\partial \Delta Y}{\partial \Delta t} > 0 \tag{15}$$

$$\frac{\partial \Delta Y}{\partial \rho} \geq 0 \tag{16}$$

$$\frac{\partial^2 \Delta Y}{\partial \Delta t \partial \rho} < 0 \tag{17}$$

The proposition is clear in our simple model, but is not obvious a priori. Imagine, for example, that bond trades are driven by consumption or endowment shocks, as in many more conventional asset pricing models or the theoretical liquidity literature building on Diamond-Dybvig (1983). If, for example, trade was driven by pure consumption shocks, the π_i would be constant and (8) shows that yield differentials would only depend on liquidity and aggregate risk directly. The cross-effect identified in (17) would be zero in this case. If agents were subject to consumption risk rather than investment risk, the π_i would depend positively on ρ rather than negatively, as in our model, and the cross-effect (17) would be positive. Then aggregate risk would amplify rather than dampen liquidity effects.

5 Empirical evidence

The empirical strategy used to test the predictions of Section 4 is based on the estimation of a simultaneous equation model for yield differentials in the Euro area at different maturities.

We measure the international risk factor as the spread between j-year fixed interest rates on U.S. swaps,$R^j_{SWUS,t}$, and the yield on j-year U.S. government bonds,$R^j_{US,t}$. We opt for this measure because of its high correlation with all U.S.-based measures of risk and because of its availability at different maturities. In the next section we report the results of estimations using alternative measures of risk and show that our results are robust to the choice of risk measure. We measure the liquidity factor by the bid-ask spread of each bond. We have considered a range of alternative liquidity indicators and selected the bid-ask spread as the most significant measure.

In taking the model to the data, we take the following specification strategy

First, we chose as benchmarks German bonds for the ten-year maturity and French bonds for the five-year maturity. Our choice is supported by the econometric evidence provided by Dunne, Moore and Portes (2002) and by the fact that traders regard French OATs as the 5-year Euro-area benchmark in the same way as they regard the 10-year Bunds as the 10-year Euro-area benchmark, because French bonds are considered as particularly liquid for the 5-year maturity bucket.

Second as yield spread in the euro area are very persistent and as our predictions are derived within a static framework, we posit the following dynamic partial adjustment model for yield differentials:

$$(R_{i,t} - R_{b,t}) = \rho_i (R_{i,t-1} - R_{b,t-1}) + (1 - \rho_i) (R_{i,t} - R_{b,t})^* + u_{i,t},$$

where $u_{i,t}$ are independently identically distributed shocks and $(R_{i,t} - R_{b,t})^*$ is the theory-consistent long-run equilibrium value vor yield differentials.. Third, we augment the specification with the differentials in the residual maturity of the benchmark bonds in country i and the benchmark bonds are included to filter out of the data the effect introduced by the different maturity of benchmark bonds and the effect of changes in benchmarks occurring at different dates for different countries in the sample period.[3] .

5.1 The baseline model

To sum up , we estimate as a baseline model the following eight-equation model, where the dependent variables are the yield differentials relative to a benchmark government bond for the other eight countries listed in Section

[3]We also tried different methods of dealing with these problems such as omitting form the sample dates in which benchmarks are changed or constructing constant maturity yields. We favour the use of the maturity differentials in that it is a natural way of correcting the differentials and it allows our liquidity indicator to operate during episodes in which liquidity might highly matter, such as at dates when benchmarks are changed.

An alternative to the maturity differential is the duration differential, however the difference between these two measures is not very relevant in our case given that they both act as dummies to model the same jump in duration and maturity occurring in occasion of benchmark changes.

3:

$$\begin{aligned}\left(R_{i,t}^{j}-R_{b,t}^{j}\right) &= \rho_{i}\left(R_{i,t-1}^{j}-R_{b,t-1}^{j}\right)+\left(1-\rho_{i}\right)\left(R_{i,t}^{j}-R_{b,t}^{j}\right)^{*}+ \\ &\quad +\delta_{i}\left(M_{i,t}^{j}-M_{b,t}^{j}\right)+u_{i,t} \\ \left(R_{i,t}-R_{b,t}\right)^{*} &= \beta_{1,i}\left(c_{i,t}-c_{b,t}\right)+\beta_{2,i}\left(R_{SWUS,t}^{j}-R_{US,t}^{j}\right)-\beta_{3,i}\left(c_{i,t}-c_{b,t}\right)\left(R_{SWUS,t}^{j}-R_{US,t}^{j}\right).\end{aligned}$$

The index i varies across countries and the index j varies across maturities (five and ten years). The estimation is performed by Seemingly Unrelated Regression (SUR), and the empirical results are shown in Tables 4.1 and 4.2. The estimates for the 10-year maturity yield differential are presented in Table 4.1.

The coefficient of the lagged dependent variable is always significant and close to unity, which indicates strong persistence of yield differentials. Also the coefficient of the maturity differential variable is uniformly positive and significant, confirming the importance of this correction. The corresponding results for the 5-year maturity are shown in Table 4.2. Again, the coefficient of the lagged dependent variable is positive and significant, but it is smaller for all eight countries, indicating lower persistence in the time-series behavior of 5-year yield differentials. Also the maturity correction coefficient stays positive and significant for all eight countries.

More importantly, the coefficient of the international risk factor is positive and significantly different from zero for all eight countries in both maturities. It ranges between 0.3 and 0.6. for the 10-year bonds and between 0.23 and 0.68 for the 5-year bonds (except in the latter case for Germany, where the coefficient is virtually zero). Interpreting these positive coeffcients from the perspective of the theory summarized in Proposition 1 leads us to conclude that the fundamental risk of non-benchamrk bonds is perceived to be more strongly affected by international risk than that of benchmark bonds. So, higher international risk – as proxied by our U.S. swap yield differential – is correlated with wider Euro-area yield differentials relative to the Bund, resp. the OAT.

Turning to the liquidity variable we note that at the 5-year maturity liquidity variables are significant in the case of Austria, Belgium, Spain, Italy, Netherlands and Portugal and, in line with the prediction of the models, the effect of the interaction between liquidity and the international risk factor has always a negative impact on the yield differentials. This evidence is substantially confirmed at the 10-year maturity, although the results are weakened in the case of Spain and Italy. This evidence illustrates the importance of non-linearities in the effect of liquidity indicators on yield differentials. Interestingly, the coefficient of the liquidity differential variable becomes significant only when the interaction between liquidity indicators and the international risk factor is also included in the regression. If the coefficient of the interaction is constrained to zero, then also the level of the liquidity indicator becomes insignificant.

This evidence does not simply reflect the fact that for less liquid bonds prices take more time to absorb the change in risks. In fact, we control for different dynamic effects across countries of the variables included in our model by having potentially different coefficients on the lagged dependent variable. Moreover a simple check, run by adding further lags of the included variables, delivers non-significant parameters for higher order dynamics. It could be observed that our SUR estimation is inefficient when valid cross-equation restrictions can be imposed on our model. This argument is is strengthened n the context of our theretical model where the cross-equation restrictions on the coefficients on the measure of liquidity and on the interaction between the international risk factor and this measure are indeed implied by the theoretical model. In Table 5, we explore this possibility by imposing cross-equation restrictions on our estimated models for 5-year and 10-year differentials. We test for the validity of cross equation restrictions on each coefficient separately and on the full set of coefficients. Interestingly, the Wald statistics reported in Table 5 illustrate that the panel restrictions can only be validly imposed on the liquidity indicators at the 10-year maturity. When these restrictions are imposed, the effect of the liquidity variables is significant and in line with the prediction of the theory. Unfortunately this result does not carry over to the 5-year maturity, where the panel restrictions cannot be imposed on the liquidity variables. On the positive side for our adopted theoretical framework, the panel restrictions on the effect of the international risk factor are always rejected, in line with the predictions of the model where the impact of international risk on fundamental risk is heterogenous across different bonds. Summing up our empirical results are generally supportive for the implications of the adopted thoeretical framework, in particular that the international risk factor is always significant and that there is an important interaction between liquidity and market returns.

5.2 Robustness

Swap spreads can be considered a good measure of risk, for a number of reasons. First, being differentials between bonds of the same maturity, they are not affected by the path of expected future risk-free rates and, differently from term spreads, they reflect only risk premia, as they are unaffected by expected monetary policy. Second, they are available at the different maturities relevant to our study, thus enabling us to account for a non-flat term structure of risk premia. Third, U.S. swap spreads provide a non-European measure of risk and therefore are much more likely to be an exogenous variable for the estimation of parameters of interest than any measure based on European yields. Fourth, as a spread between homogenous types of bonds, they are a superior measure of risk to the spread between Treasury bonds and corporate bonds.[4]. However, it must be recognized that swap spreads are a special measure of risk, in that they include the counterparty risk of swap dealers[5] and on some occasions they might reflect factors not related to international risk. A close examination of Figure 3 reveals that the positive and strong comovement between the first principal component of yield differentials in the Euro area and our measure of risk has a clear exception in late July 2003. At this date, swap spreads suddenly increased for reasons related to the hedging of mortgage-backed securities and hence little related to international factors. It is therefore important to assess the robustness of our results. We provide the relevant evidence in Table 6, where we report the results of estimating our model for the 10-year yields differentials on a shorter sample, which excludes the July 2003 episode. The table also reports the evidence obtained by augmenting the baseline regression with two alternative measures of risk. The first is the yield spread between BBB long-term corporate bonds and AAA long-term corporate bonds, the second is an indicator based on the European equity market: the implied volatility from options on the Eurostoxx 50. The results show that our estimates are robust both to the choice of the sample size and to the use of different measures of risk. In particular, the results on the shorter sample fully confirm the evidence from

[4]Duffee(1998) noted that the spread between Treasury bonds and corporate bonds is a spread between callable bonds and a mixture of callable and non-callable bonds. Given that the response of callable and non-callable bonds to shocks in the level of the term structure is different, the government-corporate spread is sensitive to the level of the term structure.

[5]Although, in practice this effect is minimal (see, for example, Duffie and Huang(1996))

our full sample with some slight modification of the original point estimates. Augmentation of the model with alternative measures of risk shows that, although all alternative measures of risk are significant, their inclusion does not affect the significance of all variables included in the original model. Overall the significance of the risk factors is more robust than that of the liquidity factor and of the interaction term. We performed similar robustness checks for the 5-year differentials, but for brevity we do not report the corresponding results, which confirm those obtained for 10-year spreads. In the case of the 5-year bonds we also re-estimated the model with the German Bund as a benchmark instead of the French OAT. This modification led to much less precise estimates of all relevant parameters and to a set of results that were much less consistent with those obtained for the 10-year differentials. We take this as confirmatory evidence of the econometric analysis of Dunne, Moore and Portes (2002) that clearly indicates the OAT as the preferred choice of benchmark for the 5-year maturity.

6 Conclusions

This paper aims to explore the determinants of observed yield differentials between long-term sovereign bonds in the Euro area. Daily data for the EMU period show that there is a strong comovement in yield differentials of benchmark bonds, and that their first principal components explains about ninety per cent of the variance. This common trend appears to be highly correlated with measures of international risk. In contrast, liquidity differentials – proxied, for example, by bid-ask spread differentials – display sizeable heterogeneity and no common factor. To generate predictions about the relation between yield differentials, fundamental risk, and liquidity, we develop a simple model of the interplay of aggregate risk and transactions costs. The model predicts that yield differentials should increase in both liquidity and risk, with an interaction term that captures the effect on required returns due to the covariotion between liquidity and market return. We test these predictions on a sample of daily data for the Euro-area sovereign yield differentials in 2002 and 2003. The econometric results show that the international risk factor is consistently priced, while liquidity differentials are priced only for a subset of nine country/maturity pairs (out of a total of 16), and that the interaction of liquidity differentials with the risk factor is always negative when significant and in line with the prediction of the model.

It is useful to stress that our data do not allow us to draw cross-sectional macroeconomic conclusions. For example, simple cross-country regressions show that on average over the sample period, (average) yield spreads were positively correlated with (average) government debt/GDP ratios, which in turn were negatively correlated with (average) bid-ask spreads. But such regressions with 9 data points have little econometric value, and in our time-series analysis we do not have sufficient variation of macroeconomic variables such as debt/GDP ratios in order to obtain conclusions about possible macroeconomic determinants of the variables that we observe. The implications of our analysis for policy-makers and for portfolio managers are rather more subtle. From a policy-making standpoint, the empirical estimates highlight the importance of the international risk factor in determining bond yield spreads, and thus underscore the importance of good macroeconomic fundamentals to minimize exposure to the international risk factor – so as to minimize not only the spreads on benchmark bonds, but also their dependence on sudden changes in the world price of risk. On the other hand, there seems to be little need for further action on the liquidity side, because bid-ask spreads are already rather uniform and very small across European bond markets, at least for benchmark bonds. Instead, the lesson for portfolio management is that liquidity can affect the risk sensitivity of the assets being held, and that this interaction depends on the covariance of illliquidity costs with international risk. In this sense, the lesson of our model – in spite of its simplicity – is considerably more general than our specific application to Euro-area bond markets.

Appendix: Description of Data

The data for 5-year and 10-year maturities for the time from 1/1/2002 to 23/12/2003 are collected from Euro MTS Group's European Benchmark Market (EBM) trading platform, at 11 a.m. CET during all market days in the Telematico cash markets. The database contains the best bid or ask prices across all markets, the aggregate quantity of all of the outstanding proposals on basis of the best bid and best ask prices, and the daily trading volume of each bond on the EBM. From these data we calculate redemption yields, maturities and a set of liquidity variables for time series consisting of the benchmark bonds for each country in our sample. The countries are Austria, Belgium, Finland, France, Germany, Italy, the Netherlands, Portugal and Spain. We constructed the following liquidity variables (in all cases as the difference between the relevant country's value and the value observed for the benchmark, which was Germany for the 10-year bucket and France for the 5-year one):

- 5-day-moving-average of the bid-ask spread (in ticks);
- trading volume for the benchmark bond, in million of Euros;
- bid-side market depth, defined as the difference between bid and mid price, divided by the bid quantity;
- ask-side market depth, defined as the difference between mid price and ask price, divided by the ask quantity;
- maximum quantity available at the best 5 prices.

After experimentation, we selected the bid-ask spread as the most significant liquidity indicator, and reported the results of estimating our non-linear empirical model only for this liquidity indicator.

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Table 1. Descriptive statistics by country *Panel A. Euro-area yield differentials relative to Germany (10y), resp. France (5y), in basis points*

	10-year benchmark bonds			5-year benchmark bonds		
Country	Mean	Median	St. dev.	Mean	Median	St. dev.
Austria	10.05	9.46	7.19	3.35	0.74	9.22
France	4.16	5.62	4.36	3.57	2.37	4.70
Netherlands	6.94	6.92	4.48	6.07	5.60	6.87
Belgium	13.45	11.79	6.80	4.78	4.40	8.09
Spain	9.72	8.06	7.44	-2.16	-0.42	10.13
Finland	10.88	9.34	8.30	6.48	582	11.18
Italy	14.47	15.70	4.88	7.97	8.34	8.01
Portugal	15.50	14.48	7.73	6.46	12.03	16.76

Panel B. Bid-ask spreads in ticks

	10-year benchmark bonds			5-year benchmark bonds		
Country	Mean	Median	St. dev.	Mean	Median	St. dev.
Austria	4.60	4.4	1.10	4.11	4.00	0.64
France	2.86	2.80	0.46	2.52	2.60	0.34
Netherlands	3.55	3.60	0.50	3.75	3.80	0.45
Belgium	3.47	3.40	0.53	2.71	2.60	0.31
Spain	3.47	3.20	0.80	2.94	2.60	0.78
Finland	4.87	4.60	1.09	4.07	3.80	0.81
Italy	2.52	2.40	1.37	2.12	2.00	0.43
Portugal	4.33	4.40	0.69	3.16	3.00	0.51
Germany	3.25	3.00	0.67	3.20	3.20	0.45

Table 2. Correlation and principal components of Euro-area yield differentials, 10 year bonds, relative to Germany *Panel A.*

Correlation matrix

Country	AT	FR	NL	BE	ES	FI	IT	PT
Austria	1	-	-	-	-	-	-	-
France	0.65	1	-	-	-	-	-	-
Netherlands	0.51	0.48	1	-	-	-	-	-
Belgium	0.88	0.72	0.61	1	-	-	-	-
Spain	0.88	0.67	0.58	0.94	1	-	-	-
Finland	0.84	0.81	0.73	0.93	0.90	1	-	-
Italy	0.75	0.84	0.52	0.82	0.80	0.89	1	-
Portugal	0.92	0.69	0.61	0.87	0.89	0.88	0.78	1

Panel B. Principal components

Component	1	2	3	4	5	6	7	8
Eigenvalue	7.28	0.26	0.16	0.13	0.06	0.05	0.03	0.01
Proportion of variance	0.91	0.03	0.02	0.016	0.008	0.006	0.004	0.001
Cumulative proportion	0.91	0.94	0.96	0.98	0.988	0.99	0.998	1

Table 3. Correlation and principal components of Euro-area bid-ask spread differentials relative to Germany *Panel A.*

Correlation matrix

Country	AT	FR	NL	BE	ES	FI	IT	PT
Austria	1.00	-	-	-	-	-	-	-
France	0.22	1.00	-	-	-	-	-	-
Netherlands	0.49	0.51	1.00	-	-	-	-	-
Belgium	0.39	0.46	0.44	1.00	-	-	-	-
Spain	0.58	0.26	0.60	0.35	1.00	-	–	-
Finland	0.48	0.21	0.54	0.26	0.76	1.00	–	-
Italy	0.09	0.37	0.19	0.26	-0.08	-0.11	1.00	-
Portugal	0.22	0.50	0.29	0.56	0.20	0.19	0.24	1.00

Panel B. Principal components

Component	1	2	3	4	5	6	7	8
Eigenvalue	3.46	1.76	0.80	0.61	0.48	0.38	0.29	0.19
Proportion of variance	0.43	0.22	0.10	0.08	0.06	0.05	0.04	0.03
Cumulative proportion	0.43	0.65	0.75	0.83	0.89	0.93	0.97	1

Table 4.1 Estimation of a system of simultaneous equations for Euro-area 10-year yield differentials

The equations are estimated by SURE, on a sample of daily observations from 1/1/2002 to 23/12/2003. The Panel shows the coefficient estimates for the 10-year maturity, spreads are on German bonds. Standard errors are reported within brackets below the coefficient estimates. An asterisk (*) and a cross (†) indicate that the corresponding coefficient is significantly different from zero at the 5 and 10 percent level, respectively.

Variable	Constant	Own lag	Maturity	Risk factor	B-A spread	Interaction
Austria	-0.167*	0.857*	0.280*	0.546*	0.043*	-0.077*
	(0.026)	(0.016)	(0.034)	(0.060)	(0.014)	(0.026)
Belgium	-0.129*	0.936*	0.357*	0.497*	0.052*	-0.099*
	(0.021)	(0.007)	(0.040)	(0.043)	(0.022)	(0.048)
Spain	-0.135*	0.867*	0.349*	0.485*	0.007	-0.009
	(0.034)	(0.018)	(0.061)	(0.077)	(0.024)	(0.047)
Finland	-0.159*	0.956*	0.207*	0.467*	-0.01	-0.025
	(0.049)	(0.006)	(0.045)	(0.118)	(0.024)	(0.079)
France	-0.119*	0.945*	0.184*	0.321*	0.016	-0.025
	(0.038)	(0.01)	(0.077)	(0.072)	(0.038)	(0.079)
Italy	-0.077*	0.912*	0.288*	0.290*	0.017	-0.042
	(0.021)	(0.01)	(0.037)	(0.047)	(0.018)	(0.043)
Netherlands	-0.076*	0.891*	0.314*	0.305*	0.034*	-0.052†
	(0.019)	(0.012)	(0.029)	(0.042)	(0.016)	(0.032)
Portugal	-0.150*	0.920*	0.384*	0.633*	0.080*	-0.139*
	(0.044)	(0.010)	(0.052)	(0.099)	(0.033)	(0.07)

Table 4.2 Estimation of a system of simultaneous equations for Euro-area 5-year yield differentials

The equations are estimated by SURE, on a sample of daily observations from 1/1/2002 to 23/12/2003. The Panel shows coefficients estimates results for the 5-year maturity, spreads are on French Bonds. Standard errors are reported within brackets below the coefficient estimates. An asterisk (*) and a cross (†) indicate that the corresponding coefficient is significantly different from zero at the 5 and 10 percent level, respectively.

Variable	Constant	Own lag	Maturity	Risk factor	B-A spread	Interaction
Austria	-0.251*	0.833*	0.170*	0.679*	0.079*	-0.184*
	(0.039)	(0.017)	(0.011)	(0.09)	(0.023)	(0.048)
Belgium	-0.082*	0.774*	0.214*	0.297*	-0.022	-0.033
	(0.015)	(0.016)	(0.008)	(0.034)	(0.018)	(0.042)
Spain	-0.143*	0.693*	0.210*	0.337*	0.048*	-0.095*
	(0.013)	(0.021)	(0.007)	(0.03)	(0.020)	(0.041)
Finland	-0.106*	0.606*	0.205*	0.258*	-0.018	-0.025
	(0.017)	(0.022)	(0.005)	(0.041)	(0.012)	(0.024)
Germany	-0.017	0.742*	0.168*	0.01	-0.007	0.004
	(0.015)	(0.018)	(0.007)	(0.03)	(0.012)	(0.026)
Italy	-0.043*	0.584*	0.172*	0.231*	0.107*	-0.208*
	(0.016)	(0.03)	(0.006)	(0.028)	(0.017)	(0.032)
Netherlands	-0.123*	0.563*	0.191*	0.317*	0.017*	-0.045*
	(0.016)	(0.021)	(0.04)	(0.036)	(0.009)	(0.020)
Portugal	-0.122*	0.853*	0.240*	0.458*	0.052*	-0.125*
	(0.022)	(0.010)	(0.006)	(0.05)	(0.018)	(0.04)

Table 5. Testing panel restrictions

The table is based on a fixed-effects panel estimates for the 10-year and 5-year yield differentials. The p-value of the Wald test of the identity restriction of individual coefficients for all eight countries is shown on the right of the relevant coefficient. The p-value of the Wald test of the identity restriction of all the coefficients for all eight countries is shown in the bottom row. Standard errors are reported within brackets below the coefficient estimates. An asterisk (*) and a cross (†) indicate that the corresponding coefficient is significantly different from zero at the 5 and 10 percent level, respectively.

	10-year yield differentials		5-year yield differentials	
Variable	Coefficient and S.E.	Wald p-value	Coefficient and S.E.	Wald p-value
Own Lag	0.956*	0.000	0.853*	0.000
	(0.006)		(0.006)	
Maturity	0.269*	0.000	0.232*	0.000
	(0.041)		(0.003)	
Risk factor	0.172*	0.000	0.372*	0.000
	(0.063)		(0.032)	
Bid-ask spread	0.047*	0.207	0.039*	0.000
	(0.021)		(0.008)	
Interaction	-0.064*	0.192	-0.087*	0.000
	(0.033)		(0.018)	
Panel restriction		0.000		0.000

Table 6. Robustness Analysis

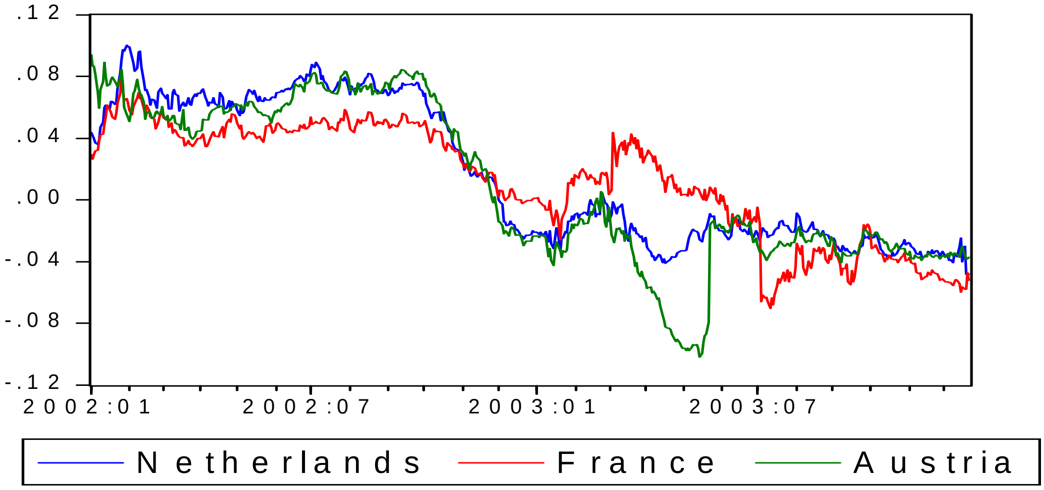
The table reports robustness analysis for the SURE system on 10-year yield differentials. We consider three alternative Risk Factors. R F 1 is the swap spread, R F 2 is the differential between yields on seasoned US BAA bonds and seasoned US AAA bonds (the source for these data is the FRED database), R F 3 is the implied volatility in options on the EUROstoxx 50. The source for these data is Datastream.

Variable	Sample	Constant	Own lag	Maturity	R F 1	R F 2	R F 3	Bid-ask	Interaction
Austria	02-01-03:06	-0.135*	0.775*	0.285*	0.503*			0.05*	-0.089*
		(0.021)	(0.022)	(0.027)	(0.046)			(0.011)	(0.020)
	02-01-03:12	-0.324*	0.811*	0.237*	0.388*	0.193*		0.035*	-0.062*
		(0.034)	(0.018)	(0.026)	(0.052)	(0.034)		(0.010)	(0.019)
	02-01-03:12	-0.199*	0.813*	0.308*	0.492*		0.178*	0.031*	-0.051*
		(0.021)	(0.018)	(0.027)	(0.047)		(0.032)	(0.010)	(0.020)
Belgium	02-01-03:06	-0.081*	0.906*	0.317*	0.433*			0.081*	-0.14*
		(0.018)	(0.011)	(0.034)	(0.035)			(0.019)	(0.039)
	02-01-03:12	-0.252*	0.925*	0.327*	0.384*	0.148*		0.042*	-0.08*
		(0.039)	(0.009)	(0.043)	(0.053)	(0.041)		(0.019)	(0.039)
	02-01-03:12	-0.177*	0.917*	0.343*	0.488*		0.174*	0.042*	-0.077*
		(0.019)	(0.008)	(0.031)	(0.033)		(0.003)	(0.019)	(0.037)
Spain	02-01-03:06	-0.10*	0.77*	0.324*	0.465*			$0.034^{\dagger}$	$-0.06^{\dagger}$
		(0.017)	(0.02)	(0.045)	(0.059)			(0.019)	(0.034)
	02-01-03:12	-0.33*	0.82*	0.239*	0.284*	0.242*		0.02	0.006
		(0.05)	(0.02)	(0.05)	(0.073)	(0.053)		(0.02)	(0.035)
	02-01-03:12	-0.17*	0.83*	0.343*	0.417*		0.203*	-0.002	0.025
		(0.027)	(0.048)	(0.048)	(0.059)		(0.055)	(0.019)	(0.04)
Finland	02-01-03:06	-0.110*	0.953*	$0.127^{\dagger}$	0.433*			-0.02	-0.017
		(0.069)	(0.009)	(0.07)	(0.15)			(0.03)	(0.06)
	02-01-03:12	-0.377*	0.944*	0.166*	0.312*	0.250*		-0.006	0.02
		(0.07)	(0.01)	(0.043)	(0.101)	(0.07)		(0.02)	(0.03)
	02-01-03:12	-0.190*	0.950*	0.282*	0.527*		0.01	-0.01	0.03
		(0.045)	(0.006)	(0.05)	(0.107)		(0.08)	(0.02)	(0.04)

Table 6. continued

Variable	Sample	Constant	Own lag	Maturity	R F 1	R F 2	R F 3	Bid-ask	Interaction
France	02-01-03:06	-0.035	0.945*	0.053	0.184*			0.042	-0.062
		(0.039)	(0.02)	(0.073)	(0.069)			(0.034)	(0.067)
	02-01-03:12	-0.176*	0.944*	0.212*	0.293*	0.057		0.01	-0.02
		(0.069)	(0.01)	(0.08)	(0.09)	(0.07)		(0.04)	(0.07)
	02-01-03:12	-0.169*	0.930*	0.163*	0.307*		0.188*	0.01	-0.02
		(0.033)	(0.012)	(0.06)	(0.056)		(0.056)	(0.03)	(0.06)
Italy	02-01-03:06	-0.027	0.88*	0.263*	0.242*			0.02	-0.042
		(0.021)	(0.02)	(0.040)	(0.043)			(0.016)	(0.039)
	02-01-03:12	-0.108*	0.89*	0.270*	0.215*	0.114*		0.01	-0.03
		(0.038)	(0.02)	(0.040)	(0.043)	(0.03)		(0.01)	(0.03)
	02-01-03:12	-0.06*	0.86*	0.237*	0.292*		0.187*	0.014	-0.036
		(0.015)	(0.01)	(0.02)	(0.03)		(0.023)	(0.01)	(0.028)
Netherl.	02-01-03:06	-0.076*	0.88*	-0.07*	0.329*			0.046*	-0.072*
		(0.02)	(0.012)	(0.026)	(0.037)			(0.016)	(0.032)
	02-01-03:12	-0.183*	0.86*	0.28*	0.180*	0.130*		0.028*	-0.042†
		(0.03)	(0.013)	(0.025)	(0.04)	(0.03)		(0.013)	(0.026)
	02-01-03:12	-0.097*	0.88*	0.33*	0.338*		0.333*	0.028	-0.031*
		(0.02)	(0.012)	(0.031)	(0.033)		(0.043)	(0.034)	(0.014)
Portugal	02-01-03:06	-0.110*	0.890*	0.406*	0.598*			0.098*	-0.173*
		(0.013)	(0.038)	(0.044)	(0.083)			(0.029)	(0.06)
	02-01-03:12	-0.327*	0.90*	0.329*	0.436*	0.224*		0.06*	-0.10†
		(0.057)	(0.012)	(0.046)	(0.095)	(0.06)		(0.027)	(0.057)
	02-01-03:12	-0.215*	0.870*	0.386*	0.589*		0.293*	0.049*	-0.082*
		(0.029)	(0.013)	(0.032)	(0.061)		(0.038)	(0.020)	(0.04)

Figure 1. 10-year yield differentials in the euro area, relative to German bonds



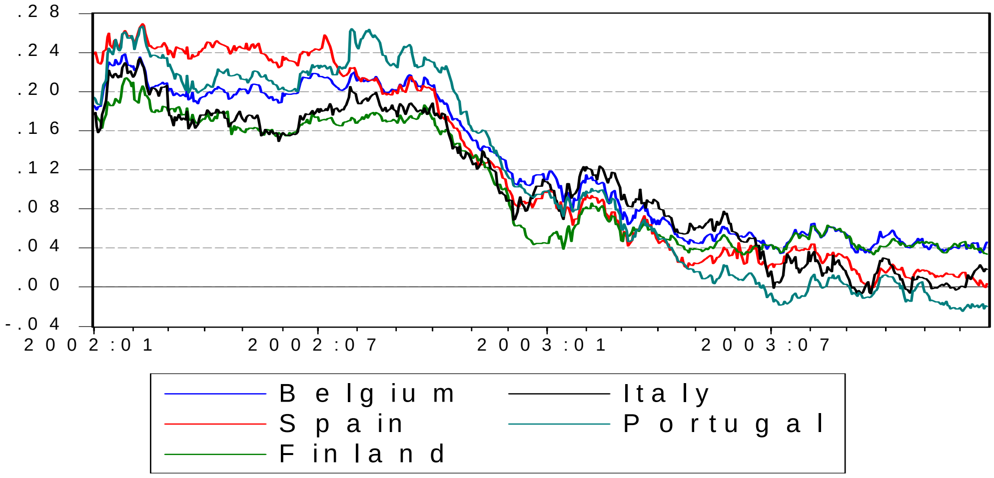
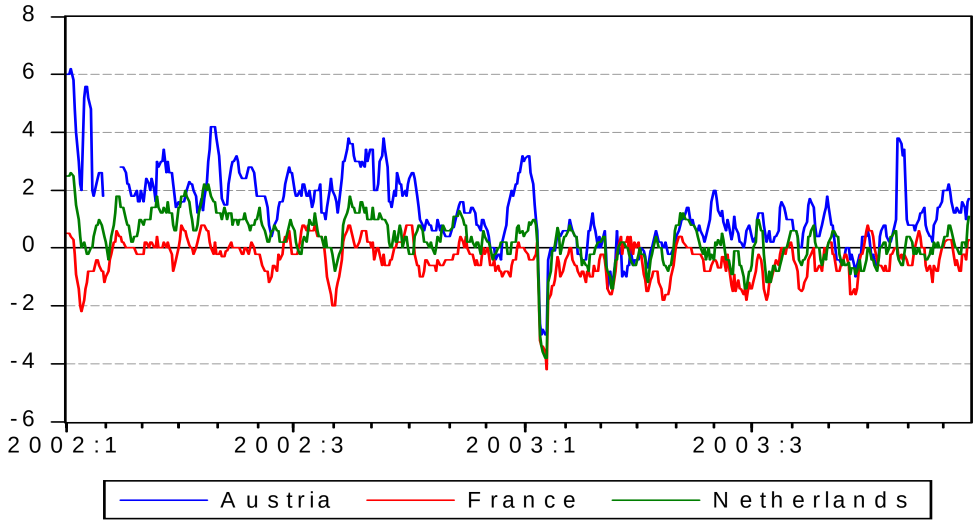


Figure 2. Bid-ask spread differentials in the Euro area, relative to German bonds (10-year benchmark bonds)



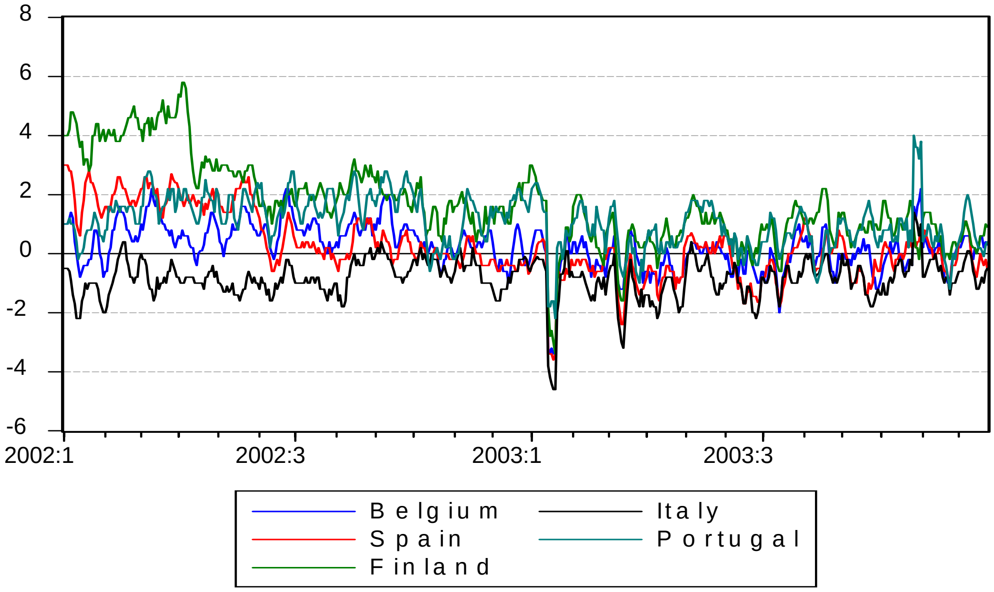


Figure 3. First principal components of Euro-area yield differentials and the spread between the 10-year fixed interest rate on swaps and US government bond yield