Valuation, Liquidity and Risk in Government Bond Markets*

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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 present 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 be the "appetite for risk" of international investors, or – to use more familiar wording – the world price of risk, which can change sharply as a result of changes in the conditional volatility of

[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.

the world market portfolio or in the risk tolerance of the marginal investor. 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 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]

The first contribution of this paper is to investigate these issues in a simple, two-period general equilibrium model of bond pricing with liquidity and default risk. The model features four assets - a riskless asset, a benchmark bond, a risky bond, and a risky outside fixed-income investment - and thus allows us at the same time to determine the relevant yields endogenously and give a precise meaning to the notion of changing "world price of risk" that is current in the empirical literature. 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 our theoretical analysis is that liquidity matters for pricing, but that it interacts with fundamental risk. In particular, if a market becomes less liquid, this can either amplify or dampen the effect of increases in the world price of risk. The sign of this interaction term depends on the liquidity of the bond relative to the benchmark and is typically different from the direct effect of liquidity. 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.

[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).

Our second 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 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) when an increase in perceived risk induces investors to require an increased yield differential on a bond, the shadow price that they place on liquidity tends to decrease: the increased risk premium is associated with a reduced illiquidity premium.

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 after-market 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 The Model

In order to guide our investigation of the determinants of yield differentials, we consider a simple discrete-time general equilibrium model that combines the analysis of risk typical of the Consumption Capital Asset Model (CCAPM) such as Lucas (1978) with the modelling of liquidity risk of the Diamond-Dybvig (1983) model, where some consumers have to liquidate their assets prematurely.

The model has three dates, $\tau = 0, 1, 2$, a non-storable consumption good at each date, and four assets. We will determine the prices (or yields) of all assets endogenously and simultaneously. The consumption good is the numeraire at each date. The period from date 0 to date 1 corresponds to a typical short-term holding period, during which liquidity needs may arise, and the period from date 1 to 2 represents the "long run". The first asset is a safe short-term asset, called S, that pays out a fixed first-period (net) interest rate $r_0 > 0$ and $r_1 \geq 0$ from date 1 to 2. We will sometimes refer to holdings of this asset as "cash". The second and third assets, denoted A and B, respectively, are government bonds that are issued (or traded) at date 0, pay out nothing at date 1, and $\widetilde{V}^i$, $i = A, B$, at date 2. The fourth asset, denoted W, describes alternative world-wide long-term investment opportunities and pays out nothing at date 1 and $\widetilde{V}^W$ at date 2. All payouts are in terms of the consumption good.

We assume that bond A is safe and let $\widetilde{V}^A = V$ almost surely.[3] Although no Western European government bond in the postwar period has ever been defaulted upon, it seems that markets attribute a slight default risk to some bonds. To capture this perception (which may be purely psychological), we assume that $\widetilde{V}^B$ is indeed a random variable, with support $[0, \overline{V}]$ and a large mass on $\overline{V}$. Purely for convenience, we standardize the expected final payoff of the risky bond to equal that of the safe bond: $E\widetilde{V}^B = V$, so that the payoff of the risky bond is a mean-preserving spread of the latter. The final value of the world asset $\widetilde{V}^W$ is an arbitrary real-valued random variable. We denote $\sigma_B^2 =$var$(\widetilde{V}^B)$, $\sigma_W^2 =$var$(\widetilde{V}^W)$, $\sigma_{BW} =$cov$(\widetilde{V}^B, \widetilde{V}^W)$, and assume that the final payoffs of the risky bond and the world-wide asset opportunity are positively correlated: $\sigma_{BW} > 0$. This assumption is intended to capture the idea that default on the risky bond is more likely in states in which the world economy fares badly.

All assets are traded at dates 0 and 1. We denote date-τ prices by p_τ^i, $i = S, A, B, W$. By assumption, the S-asset trades without frictions. For the other assets, we model trading frictions explicitly by assuming that market makers set bid prices $(1 - t_0^i)p_0^i$ and ask prices $(1 + t_0^i)p_0^i$ in period 0, and similarly bid prices $(1-t_1^i)p_1^i$ and ask prices $(1+t_1^i)p_1^i$ at date 1. The t_0^i and t_1^i represent proportional transactions costs that can be due to order processing costs, asymmetric information or other motives discussed in the literature. We take them to be exogenous and symmetric around p_τ^i, the fundamental value of asset i at time τ. We also assume $t^W = 0$, since we want to focus on the impact of differential transactions costs on government bond trading.

In this fashion, the model allows us to analyze separately the impact of changes in current and future liquidity: from the standpoint of an investor who is choosing her portfolio at date 0, an increase in t_0^i is a decrease in the current liquidity of asset i, whereas an increase in t_1^i is a decrease in its expected liquidity, which will be relevant to the investor if she trades in period 1. In this sense, t_0 describes current transactions costs, whereas t_1 is the long-term (expected) level of transactions costs.

In the spirit of the "tree model" by Lucas (1978), we assume that the four assets are in fixed supply Q^i, $k = S, A, B, W$. These asset supplies are purchased at the respective ask prices by a continuum of agents of mass 1 at date 0. (Asset payoffs are normalized to a per-capita basis.) Agents can be thought as purchasing these asset supplies from a previous generation of

[3] A may be thought of as Germany, at least until 2002.

investors, who are selling them inelastically.

Investors are identical *ex ante*, all having the same preferences and being endowed with 1 unit of the numeraire in period 0 and nothing thereafter. *Ex post*, however, individuals differ, because they are exposed to different liquidity needs. To capture this in a simple way, we adopt a Diamond-Dybvig setting. We assume that utility from consuming in $\tau = 1$ and $\tau = 2$ is given by

$$u(c_1, c_2) = \begin{cases} u(c_1) & \text{with probability } \pi \\ u(c_2) & \text{with probability } 1 - \pi \end{cases}$$

where u satisfies the usual boundary conditions to get interior optima, and π is the probability of early liquidation. For simplicity (and without loss of generality) we assume no discounting. We also assume the Law of Large Numbers such that at date 1 a fraction of exactly π agents liquidate prematurely and consume. We call agents who must consume at date 1 "impatient" and those who must consume at date 2 "patient".

4.1 Optimality and equilibrium conditions

Consider an agent at date 1 who holds a portfolio $(y_0^S, y_0^A, y_0^B, y_0^W)$ chosen at date 0. (In equilibrium, of course, $(y_0^S, y_0^A, y_0^B, y_0^W) = (Q^S, Q^A, Q^B, Q^W)$) The date 0 budget constraint is

$$p_0^S y_0^S + \sum_{i=A,B,W} (1 + t_0^i) p_0^i y_0^i = 1. \tag{1}$$

If the agent is impatient, at date 1 she sells her portfolio inelastically and consumes

$$\begin{aligned} c_1 &= (p_1^S + r_0) y_0^S + \sum_{i=A,B,W} (1 - t_1^i) p_1^i y_0^i \\ &= \frac{p_1^S + r_0}{p_0^S} + \sum_{i=A,B,W} \left((1 - t_1^i) p_1^i - \frac{p_1^S + r_0}{p_0^S} (1 + t_0^i) p_0^i \right) y_0^i \end{aligned} \tag{2}$$

where the last equality used (1). Note that $(p_1^S + r_0)/p_0^S - 1$ is the first-period rate of return on the safe asset (the date-0 risk-free rate).

If the agent is patient, she uses the return on her cash holdings $r_0 y_0^S$ to buy quantities $(y_1^S, y_1^A, y_1^B, y_1^W)$ at date 1 so as to

$$\text{maximize } Eu(c_2),$$

where

$$c_2 = (1 + r_1)(y_0^S + y_1^S) + \sum_{i=A,B,W} \left(y_0^i + y_1^i\right) \widetilde{V}^i,$$

subject to the budget constraint

$$r_0 y_0^S = \sum_{i=S}^{W} (1 + t_1^i) p_1^i y_1^i. \tag{3}$$

Here we have used the fact (which can easily be proved formally) that patient agents will not sell the two bonds because of their transactions costs (i.e. that optimally $y_1^i \geq 0$ for $i = A, B$). Using (1) and (3) to eliminate the y_τ^S, final consumption c_2 can be written as

$$\begin{aligned} c_2 \quad = \quad & \frac{1 + r_1}{p_1^S} \frac{p_1^S + r_0}{p_0^S} + \sum_{i=A,B,W} y_0^i \left(\widetilde{V}^i - \frac{1 + r_1}{p_1^S} \frac{p_1^S + r_0}{p_0^S} (1 + t_0^i) p_0^i \right) \\ & + \sum_{i=A,B,W} y_1^i \left(\widetilde{V}^i - \frac{1 + r_1}{p_1^S} (1 + t_1^i) p_1^i \right). \end{aligned} \tag{4}$$

The first term in (4) is the rate of return on the safe asset between dates 0 and 2, the second the two-period excess return on date 0 investments, and the third the one-period excess return on date-1 investments.

The first-order conditions for the patient investors' problem at date 1 are

$$E \left(\widetilde{V}^i - \frac{1 + r_1}{p_1^S} (1 + t_1^i) p_1^i \right) u'(c_2) = 0 \tag{5}$$

for $i = A, B, W$.

At date 0, agents anticipate date-1 equilibrium prices, taking them as given, and maximize

$$\pi u(c_1) + (1 - \pi) E u(c_2)$$

over (y_0^A, y_0^B, y_0^W), where c_1 and c_2 are given by (2) and (4), respectively, and the y_1^i by the first-order conditions (5). Using the Envelope Theorem, the date-0 first-order conditions are

$$\begin{aligned} & \pi u'(c_1) \left((1 - t_1^i) p_1^i - \frac{p_1^S + r_0}{p_0^S} (1 + t_0^i) p_0^i \right) \\ & + (1 - \pi) E u'(c_2) \left(\widetilde{V}^i - \frac{1 + r_1}{p_1^S} \frac{p_1^S + r_0}{p_0^S} (1 + t_0^i) p_0^i \right) = 0 \end{aligned} \tag{6}$$

for $i = A, B, W$.

Combining (5) and (6) for $i = W$ and recalling the assumption $t_0^W = t_1^W = 0$, we obtain the standard arbitrage relationship

$$\frac{p_1^W}{p_0^W} = \frac{p_1^S + r_0}{p_0^S}, \tag{7}$$

which simply states that the rates of return of the two frictionless assets between dates 0 and 1 must be equal. To close the model, the optimality conditions must be complemented with the market clearing conditions at date 0,

$$y_0^i = Q^i$$

and at date 1,

$$\pi y_0^i = (1-\pi) y_1^i \tag{8}$$

for $i = S, A, B, W$. Inserting these two sets of conditions into the respective budget constraints yields

$$p_0^S Q^S + \sum_{i=A,B,W} (1+t_0^i) p_0^i Q^i = 1 \tag{9}$$

$$p_1^S Q^S + \sum_{i=A,B,W} (1+t_1^i) p_1^i Q^i = \frac{1-\pi}{\pi} r_0 Q^S \tag{10}$$

Finally, in equilibrium total consumption at each date must equal the total quantity of the consumption good available. Expressed in per-capita terms, this means:

$$c_1 = \frac{1}{\pi} r_0 Q^S \tag{11}$$

$$c_2 = \frac{1}{1-\pi}\left[(1+r_1)Q^S + VQ^A + \widetilde{V}^B Q^B + \widetilde{V}^W Q^W\right] \tag{12}$$

Equations (5), (6), (9), and (10) are 8 equations in the 8 unknowns $p_\tau^i, \tau = 0, 1, i = S, A, B, W$. Due to the simple intertemporal structure of the model, this system of equations can be solved fairly easily. In particular, the date-1 problem can be solved independently from the date-0 problem.

4.2 Date-1 equilibrium prices

First, because bond A is riskless, condition (5) for $i = A$ yields a standard riskless no-arbitrage relation:

$$p_1^A = \frac{1}{1+t_1^A}\frac{p_1^S}{1+r_1}V, \tag{13}$$

stating that the price of the safe asset A is its final value V multiplied by the discount factor $p_1^S/(1+r_1)$ and by a "transaction cost" factor. Transaction costs operate like a tax: the larger the cost t_1^A charged to buy asset A at date 1, the lower its "net price" p_1^A.

The two first-order conditions (5) for assets B and W yield analogous expressions for p_1^B and p_1^W (remember that $t^W = 0$):

$$p_1^B = \frac{1}{1+t_1^B}\frac{p_1^S}{1+r_1}\frac{E\widetilde{V}^B u'(c_2)}{Eu'(c_2)} \tag{14}$$

$$p_1^W = \frac{p_1^S}{1+r_1}\frac{E\widetilde{V}^W u'(c_2)}{Eu'(c_2)} \tag{15}$$

Using (13), (14), (15),and (11), (12) and the market clearing condition (10) then yield the equilibrium discount factor as

$$\frac{p_1^S}{1+r_1} = \frac{Eu'(c_2)}{Ec_2u'(c_2)}c_1 = \left[1 - \frac{\mathrm{cov}(c_2, u'(c_2))}{Ec_2u'(c_2)}\right]\frac{c_1}{Ec_2} \tag{16}$$

The first factor on the right-hand side is larger than 1, because the concavity of $u(\cdot)$ implies that $\mathrm{cov}(c_2, u'(c_2))$ is negative and increasing in the variability of date-2 consumption.[4] Intuitively, an increase in risk makes the short-term asset more appealing to investors and tends to raise its price. The second term reflects the demand for interim liquidity c_1 by impatient consumers relative to the expected consumption of the patient ones.

Replacing the discount factor (16) in equations (13), (14) and (15) yields the date-1 equilibrium prices of the three long-term assets, p_1^A, p_1^B and p_1^W. These expressions allow us to compare the price responses of the various assets to changes in risk and contemporaneous transaction costs.[5] By the

[4] Indeed it equals 1 in the limiting cases where c_2 is certain or where $u(c_2)$ is linear.

[5] This static analysis ignores the role of expected (long-term) transactions costs, which only the full three-period model can exhibit. But this preliminary analysis is already informative and helps to better understand the full analysis in the next subsection.

multiplicative nature of the pricing kernel, it is more natural to undertake such a comparison with price ratios than with differences. Thus, the ratio between the price of the risky bond and that of the safe one is

$$\frac{p_1^B}{p_1^A} = \frac{1+t_1^A}{1+t_1^B}\frac{E\widetilde{V}^B u'(c_2)}{EVu'(c_2)} = \frac{1+t_1^A}{1+t_1^B}\left[1+\frac{\text{cov}(\widetilde{V}^B, u'(c_2))}{EVu'(c_2)}\right] \tag{17}$$

Equation (17) shows very clearly two elements that drive a wedge between the valuations of bond A and B: differences in transaction costs, expressed by the first fraction, and differences in risk, expressed by the second one. The first term is smaller than 1 if the risky bond is less liquid than the safe one ($t_1^A < t_1^B$), and shows that the date-1 relative price of the risky bond is inversely related to its transaction cost t_1^B. The second term is smaller than 1 because, as before, the covariance in this term is negative. By the same token it is decreasing in the covariance of $\widetilde{V}^B$with c_2, which can be computed from (12).

An important implication of (17) is that, because the trading cost term and the risk term enter multiplicatively, an increase in the risk of bond B has a smaller effect on its relative price if its transaction cost is relatively large. In other words, while an increase in the risk of the B-bond, either through its variance or its covariance with world-wide returns, affects its relative price negatively, the magnitude of this negative effect is dampened by a relatively large transaction cost t_1^B. To summarize these results:

Remark 1: *The date-1 price ratio between the risky and the safe bond depends negatively on their relative transaction costs and on consumption risk. The effect of an increase in risk is smaller the larger is the relative transaction cost of the risky bond.*

The first part of this remark is straightforward: if the trading costs of bond B or its risk increase relative to bond A, its relative price decreases because patient investors will attempt to substitute away from this asset towards more attractive ones. This is the well-known point made by the "transaction cost view" described in Section 2: in equilibrium, higher transaction costs must be compensated by lower prices. The second finding, concerning the interaction between risk and transaction costs is less obvious. It is a key insight about the importance of interaction effects for the modelling of asset prices that we will discuss more broadly in the following section. But already this simpler context of period -1 prices provides some interesting intuition. This intuition is similar to the logic of trading distortions resulting from taxation in public economics. Suppose that on an asset the buyer and seller must pay a proportional transaction tax. Then the larger is the tax, the lower the after-tax price faced by either one. Suppose that the asset becomes riskier, so that its price tends to fall. The effect on the price will be smaller the larger is the tax, since the initial after-tax price will be correspondingly lower. The tax effectively reduces the variance of the price arising from news about the future.

4.3 Date-0 equilibrium prices

We now turn to our main objective, the analysis of date-0 prices. As of period 0, there are two sources of risk, liquidity risk and fundamental risk. The former arises from the fact that date-1 consumption is random and subject to liquidity costs, the second from randomness in the returns of the local and the international assets Different from the analysis of the last subsection, the date-0 problem allows us to study the impact of changes in contemporaneous transactions costs on prices (the impact of t_0 on p_0) holding the average (expected) level of transactions costs (t_1) fixed, and viceversa.

First, we study the relation between the initial price of the risky bond p_0^B and that of the safe bond p_0^A. Combining the first-order condition at date 0, (6), with that of date 1, (5), we obtain

$$\left(\pi \frac{1-t_1^i}{1+t_1^i} p_1^S u'(c_1) + (1-\pi)(1+r_1)Eu'(c_2)\right) \frac{E\widetilde{V}^i u'(c_2)}{Eu'(c_2)}$$
$$= \frac{p_1^S + r_0}{p_0^S}(1+r_1)\left(\pi u'(c_1) + (1-\pi)\frac{1+r_1}{p_1^S}Eu'(c_2)\right)(1+t_0^i)p_0^i$$

for $i = A, B, W$. Dividing the equation for A by the one for B yields

$$\frac{p_0^B}{p_0^A} = \frac{1+t_0^A}{1+t_0^B} \cdot \frac{E\widetilde{V}^B u'(c_2)}{EVu'(c_2)} \cdot \frac{\pi \frac{1-t_1^B}{1+t_1^B} p_1^S u'(c_1) + (1-\pi)(1+r_1)Eu'(c_2)}{\pi \frac{1-t_1^A}{1+t_1^A} p_1^S u'(c_1) + (1-\pi)(1+r_1)Eu'(c_2)}, \quad (18)$$

which upon substituting p_1^S from (16) becomes

$$\frac{p_0^B}{p_0^A} = \frac{1+t_0^A}{1+t_0^B} \cdot \frac{E\widetilde{V}^B u'(c_2)}{EVu'(c_2)} \cdot \frac{\pi \frac{1-t_1^B}{1+t_1^B} c_1 u'(c_1) + (1-\pi)Eu'(c_2)c_2}{\pi \frac{1-t_1^A}{1+t_1^A} c_1 u'(c_1) + (1-\pi)Eu'(c_2)c_2} \quad (19)$$

Condition (19) generalizes the static relation (17) and shows that the full wedge between the valuations of bond A and bond B is driven by three components that enter multiplicatively: (i) the difference in their current transaction costs, t_0^A and t_0^B; (ii) the consumption risk associated with bond B and (iii) an interaction term of consumption risk $(Eu'(c_2)c_2)$ and future transactions costs (t_1^i). The first and the second terms are the same as those found in the date-1 price ratio (17) and already discussed above, except of course for the time index of the current transaction costs. In fact, if the (average) future liquidity of the two bonds is the same ($t_1^A = t_1^B$), the third term equals 1, leaving only the first and the second term in (19): in this special case, the date-0 comparative statics are exactly the same as those of period 1 and described by Remark 1.

In particular, by focussing only on the first and the second fraction on the right-hand side of (19), we again find that the price ratio of the two bonds depends negatively on their relative current transactions costs $(1 + t_0^B)/(1 + t_0^A)$, but that this effect is dampened by the effect of consumption risk stemming from the covariance of $\widetilde{V}^B$ and c_2.[6] We call this effect of consumption risk on prices the *static* effect, as it is already present in the static problem at date 1. The following proposition summarizes these key results of our analysis.

Proposition 1: *The price ratio between the risky and the safe bond depends negatively on their relative current transaction costs. The static effect of consumption risk on the price ratio is negative, and it is the smaller the*

[6]It may seem that this comparative statics result depends on the simplifying assumption that liquidity sellers are perfectly inelastic. But in fact, the result continues to hold in a more realistic model with imperfectly elastic liquidity traders as in Admati and Pfleiderer (1988). The analysis in such a model becomes more complicated, because date-2 consumption becomes endogenous, and therefore the pricing kernels in the valuation equation more complicated. But now there is even an additional reason for the impact of the interaction term. Imagine a liquidity seller who can decide which asset to liquidate. Clearly, she will always want to liquidate less of the less liquid asset that she has in her portfolio. This says that, as she owns the asset, trading costs prevent her from selling it and thereby raise her demand for it, precisely the opposite from what trading costs would do ex ante (when they reduce demand, since one knows that one may have to liquidate the asset at high cost). This means that if trading costs increase at time 1, one tends to sell less of the asset. Consequently, if there is *simultaneously* an unexpected increase in the risk of an asset *and* an unexpected increase in its trading cost, then the effect of the increase in risk will be softened by the higher trading cost, since the latter will prevent investors from rushing out of that asset to the same extent than they would have done otherwise.

larger is the relative current transaction cost of the risky bond.

The third term in (19) introduces a new element compared to the analysis of date-1 prices. The term $(1-t_1^i)/(1+t_1^i)$ is the ratio of the anticipated bid and ask prices of bond i in date-1 trading. As such, it is decreasing in asset i's trading cost t_1^i, and can be considered a measure of its future liquidity. So the third term indicates that the more liquid asset B is expected to be relative to asset A in date-1 trading, the higher its relative price as of period 0 – and vice versa. Interestingly, the price relevance of this expected liquidity differential is increasing in the investors' probability of having to liquidate the asset prematurely. The greater π, the more exposed are investors to liquidity risk, and the lower is the relative price of the less liquid bond.

This shows that the effect of future transactions costs on current prices is qualitatively the same as that of current transactions costs: the higher (future or current) transactions costs, the lower the current price. This extends the "trading cost view" of liquidity (Amihud and Mendelson (1986)), by showing that.the negative effect of future transactions costs on future prices feeds back into current prices.

However, (19) also shows that if (and only if) the two bonds differ in their future liquidity, the variance of final consumption has an additional effect on the date-0 price ratio. We call this the *dynamic* effect of consumption risk on prices, because it only arises in the dynamic pricing problem of date 0.

If bond B is *less liquid* than bond A ($t_1^B > t_1^A$), then an increase in consumption risk (measured by $-Eu'(c_2)c_2$) reduces the third term, thereby amplifying the negative static price effect of an increase in risk on the relative price of bond B. Therefore in this case, the interaction between risk and the *future* transaction cost differential operates in the opposite direction relative to that between risk and the *current* transaction cost differential, which was shown to dampen the price effect of a risk increase.

However, the opposite occurs if the bond B is *more* liquid than bond A ($t_1^B < t_1^A$). Then, an increase in consumption risk increases the third term, and therefore dampens the negative static effect of increases in risk on the relative price of bond B. This shows that the interaction between risk and the future liquidity differential of the two bonds may lead to a positive dynamic effect of risk on the relative price of the riskier bond. Intuitively, this happens because when the riskier bond is comparatively very liquid, an increase in risk can induce a flight into the market that is expected to be more liquid instead of the safer market. From the standpoint of a date-0 investor, in fact, there are two sources of risk to be feared: interim consumption risk and final consumption risk. Interim consumption risk is increasing in the size of date-1 transaction costs. On an *ex-ante* basis, an increase in final consumption risk may induce investors to prefer an asset with high interim liquidity and high final consumption risk to an asset with the opposite characteristics. In short, if "liquidity" and "quality" are separated, the flight to liquidity counteracts the flight to quality as an investor reaction to increases in risk.

The following proposition summarizes the foregoing discussion of the third term in the pricing equation (19), the dynamic effect of risk.

Proposition 2: *The price ratio between the risky and the safe bond depends negatively on their relative future transaction costs. Furthermore, if the risky bond is less liquid than the safe bond, the negative static effect of consumption risk on the price ratio (described in Proposition 1) is amplified by the dynamic effect. Hence, the total effect of consumption risk is negative. If the risky bond is more liquid than the safe bond, the dynamic effect of an increase in consumption risk counteracts the static effect, and the total effect is ambiguous.*

Conceptually, it is important to note that Proposition 2 makes a cross-sectional statement (what is the effect of consumption risk on prices for different bonds with different average transactions costs?), whereas Proposition 1 is concerned with the time-series behavior of a given bond (how do current transactions costs and risk factors impact current prices?). The main implication of the proposition is that if the risky bond is expected to be less liquid than the safe bond ($t_1^B \geq t_1^A$), the third term in (19) reinforces the negative static effect of risk on prices discussed in Proposition 1.

In our data, this is the case for all countries except for Italy. For the 10-year bond, Germany is the benchmark (bond A) and combines the risk-free status with low liquidity costs, mostly thanks to its large issuing volumes and the dominant Bund futures market. Hence, for all bonds except for Italy, we expect the yield spread over the Bunds to be increasing in the risk factors. For the 5-year bond, France has emerged as the benchmark, again a country with AAA rating, considered to be riskless, and the same predictions arise. In our data only Italy has higher average liquidity (i.e. lower average transactions costs) than the benchmark in both the 5-year and the 10-year range. This is due to a very high debt volume and a historically efficient electronic market for government debt. According to the theory, therefore, the effect of consumption risk should be weakest for Italy, and we cannot even exclude a priori that it takes a different sign from that of the other countries.

In the econometric analysis, we cannot measure risk factors directly. Instead, we use different proxies for the world investment opportunity W, and argue that its price depends negatively on its riskiness. In fact, from the arbitrage relationship (7), the price of the world-wide investment opportunity relative to the safe asset is

$$\frac{p_0^W}{p_0^S} = \frac{p_1^W}{p_1^S + r_0}$$

Hence, the impact of risk on date-0 prices is the same as that on date-1 prices. As discussed following (16), p_1^S increases if the risky investments become riskier (the flight to quality). Similarly, p_1^W decreases (a move into either safe assets or the B bond, depending on how the risk of the worldwide asset affects the B bond). Overall, therefore, p_0^W/p_0^S decreases with the risk of worldwide investment. Formally, using (15),

$$\frac{p_0^W}{p_0^S} = \frac{1}{1+r_1}\frac{p_1^S}{p_1^S + r_0}\frac{E\widetilde{V}^W u'(c_2)}{Eu'(c_2)} \tag{20}$$

which is decreasing in the riskiness of $\widetilde{V}^W$ and c_2.

Proposition 3: *The relative price of the worldwide investment opportunity decreases in consumption risk.*

The effects described in the previous propositions can be illustrated by specializing the model to CARA utility and using first-order approximations, that is, by reducing the analysis to a consumption-CAPM framework. Using

$$u'(c_2) \approx u'(Ec_2) + (c_2 - Ec_2)u''(Ec_2)$$

we get

$$\begin{aligned} E\widetilde{V}^i u'(c_2) &\approx u'(Ec_2)\left[E\widetilde{V}^i - a\text{cov}(\widetilde{V}^i, c_2)\right], i = B, V \\ Ec_2 u'(c_2) &\approx u'(Ec_2)\left[Ec_2 - a\text{var}(c_2)\right], \\ EVu'(c_2) &\approx VE[u'(Ec_2) + (c_2 - Ec_2)u''(Ec_2)] = Vu'(Ec_2) \end{aligned}$$

where $a > 0$ is the coefficient of absolute risk aversion. So we can approximate the price ratios solely in terms of first and second moments. For the pricing formula (19) we get

$$\frac{p_0^B}{p_0^A} \approx \frac{1+t_0^A}{1+t_0^B} \cdot \frac{\pi \frac{1-t_1^B}{1+t_1^B} c_1 u'(c_1) + (1-\pi) u'(Ec_2)\left[Ec_2 - a\mathrm{var}(c_2)\right]}{\pi \frac{1-t_1^A}{1+t_1^A} c_1 u'(c_1) + (1-\pi) u'(Ec_2)\left[Ec_2 - a\mathrm{var}(c_2)\right]} \cdot \left[1 - \frac{a}{V}\mathrm{cov}(\widetilde{V}^B, c_2)\right]. \tag{21}$$

The relevant second moments are

$$\begin{aligned} \mathrm{var}(c_2) &= \left(\frac{Q^B}{1-\pi}\right)^2 \sigma_B^2 + \left(\frac{Q^W}{1-\pi}\right)^2 \sigma_W^2 + \frac{2Q^B Q^W}{(1-\pi)^2}\sigma_{BW}, \\ \mathrm{cov}(\widetilde{V}^B, c_2) &= \frac{Q^B}{1-\pi}\sigma_B^2 + \frac{Q^W}{1-\pi}\sigma_{BW}, \\ \mathrm{cov}(\widetilde{V}^W, c_2) &= \frac{Q^W}{1-\pi}\sigma_W^2 + \frac{Q^B}{1-\pi}\sigma_{BW}. \end{aligned}$$

These formulas allow to illustrate the discussions of Propositions 1 and 2 by means of just the three parameters σ_W^2, σ_B^2, and σ_{BW}. To simplify even further, it is reasonable to assume that over the relatively short period of two years that our data cover, the variance of European bond prices as well as their correlation with international investments have remained relatively stable.[7] We therefore take the variance of world-wide investment opportunities, σ_W^2, to be the driving force of changes in final consumption risk, with the covariance σ_{BW} moving accordingly.

In order to test the model empirically, we combine Propositions 1, 2, and 3, to get predictions about the relation between the price of the world-wide investment opportunity (p_0^W/p_0^S), the liquidity of a bond ($1/t^i$), and its relative price (p_0^B/p_0^A). The link between p_0^W/p_0^S and p_0^B/p_0^A are the common risk factors in the respective covariances of final consumption risk, or, in the CARA approximation discussed before, the variance and covariance σ_W^2 and σ_{BW}.

Propositions 1 through 3 taken together imply that, if the bond is less liquid than the reference benchmark, the relative price of a bond p_0^B/p_0^A is positively correlated with the relative price of worldwide investment, p_0^W/p_0^S (the "international risk factor"), and that this correlation is smaller the more illiquid is bond B relative to bond A, since the cross effect of risk and liquidity

[7] See, e.g., Codogno, Favero and Missale (2003).

is negative. However, Proposition 2 shows that, if the bond is more liquid than the benchmark, the flight to liquidity induced by greater risk works in the opposite direction. Therefore, for highly liquid bonds it is an empirical matter whether changes in liquidity dampen or amplify the price impact of risk changes.

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. In taking the model to the data, we make three approximations.

First, in our estimation the dependent variables are yield differentials, while the model characterizes the effect of changes in risk and liquidity on the price ratios. Empirically, the difference is negligible, as a Taylor approximation of $\log(1+r)$ for small r shows. The time-series of yield differentials and ratios have a correlation of over 0.95 in our sample.

Second, the model distinguishes between the effects of the current transaction costs of bond i (t_0^i) and those of its future transaction costs (t_1^i) on its date-0 price (p_0^i), and assumes that future transaction costs are perfectly anticipated by investors. Since the estimation is conducted at daily frequency, and bid-ask spread differentials feature considerable persistence at that frequency, changes in current bid-ask spread differentials may lead investors to revise expectations of future bid-ask spread differentials in the same direction. However, for lack of a plausible model of expectation formation by investors we treat future transactions costs as given and focus on the impact of current transactions costs on current prices.

Third, as already mentioned in the previous sections, we shall not measure consumption risk directly, but rather proxy changes in risk by the yield of U.S. corporate debt relative to government debt, so that U.S. corporate debt will play the role of the W-bond in the model. In Section 5.2 we report results obtained with other proxies for the "world price of risk", showing that our results are robust.

5.1 The baseline model

As a baseline, we estimate 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:

$$\begin{aligned} R_{\tau}^{i,j}-R_{\tau}^{B,j} &= \beta_1^{i,j}\left(R_{\tau-1}^{i,j}-R_{\tau-1}^{B,j}\right)+\left(1-\beta_1^{i,j}\right)\left(\beta_0^{i,j}+\beta_2^{i,j}\left(M_{\tau}^{i,j}-M_{\tau}^{B,j}\right)\right) \\ &\quad +\left(1-\beta_1^{i,j}\right)\left(\beta_3^{i,j}\left(L_{\tau}^{i,j}-L_{\tau}^{B,j}\right)+\beta_4^{i,j}\left(R_{\tau}^{SWUS,,j}-R_{\tau}^{US,j}\right)\right) \\ &\quad +\left(1-\beta_1^{i,j}\right)\beta_4^{i,j}\left(R_{\tau}^{SWUS,,j}-R_{\tau}^{US,j}\right)\left(L_{\tau}^{i,j}-L_{\tau}^{B,j}\right)+u_{\tau}^{i,j}. \end{aligned}$$

The index i varies across countries and the index j varies across maturities (five and ten years). 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. Yield spreads in the Euro area, $R_{\tau}^{i,j}-R_{\tau}^{B,j}$, are explained by their own lag (to capture persistence in the data), by the international factor $R_{\tau}^{SWUS,,j}-R_{\tau}^{US,j}$, by the liquidity differential $L_{\tau}^{i,j}-L_{\tau}^{G,j}$, and by the difference between the residual maturity in benchmark bonds used to form the yield differentials, $M_{\tau}^{i,j}-M_{\tau}^{G,j}$.

In the baseline specification, we measure the international risk factor as the spread between j-year fixed interest rates on U.S. swaps and the yield on j-year U.S. government bonds. 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. This variable is interacted with the international risk factor to allow for the non-linearity in their effect that our model suggests. This interaction is consistent with the preliminary investigation of the data reported in Section 3.

Finally, 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.[8]

[8] We also tried different methods of dealing with these problems such as omitting form

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). So, as predicted by the model, higher 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.

As discussed following Proposition 2, the chosen benchmarks (Germany for the 10-year bucket, France for 5 years) combine the (virtual) risk-free status with high liquidity. Therefore, Proposition 2 predicts that there is no flight to liquidity that reverses the sign of the coefficient, which is borne out by the data. As shown by Table 1, in our data only Italy has higher average liquidity (i.e. lower average transactions costs) than the benchmark in both the 5-year and the 10-year range. And consistent with Proposition 2, the effect of final consumption risk in both estimations is weakest for Italy. Similarly, the only other two countries with 10-year bond markets roughly as liquid as the German one are France and the Netherlands, and also for these countries the risk coefficient is comparatively low in Table 4.1.

The coefficient of the liquidity differential for the 10-year maturity is positive for all countries except Finland, but it is significantly different from zero only for Austria, Belgium, the Netherlands and Portugal. So, in keeping with the model's predictions, for these four countries a higher bid-ask spread is

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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.

associated with a higher yield spread relative to Germany. Importantly, in all four cases, the positive effect of the bid-ask spread on yield differentials is paired with a significantly negative coefficient of the interaction term between the liquidity measure and the international risk factor. This evidence illustrates the importance of non-linearities in the effect of liquidity indicators on yield differentials, showing that a higher transaction cost differential tends to dampen the effect that changes in risk have on yields. 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. The result rather confirms our finding that liquidity operates via two opposing channels, which cancel each other if not specified separately. In fact, in the estimates of Tables 4.1 the interaction coefficient, when applied to the empirical estimate of the international risk factor is quite precisely of the same size as the liquidity coefficient, only with the opposite sign. The same occurs in Table 4.2.

It is tempting to attribute the significance of the coefficients to a "small country" effect – meaning that investors prefer to move out of smaller countries and into Bunds when their perceived risk increases.[9] But this is rejected by the results for the 5-year bonds, where the benchmark are French OATs. Also in this case, the sign of the liquidity differential is positive and significant and the interaction term is negative and significant for five countries: Austria, Spain, Italy, Netherlands and Portugal. In the case of Italy and Spain the liquidity variables are significant for 5-year differentials but not for 10-year differentials, the opposite is true for Belgium. What is remarkable is that in all cases, for both maturities (i.e., in 9 out of 16 cases), liquidity is significant if and only if the interaction between liquidity and international

[9] As, for example, the quote from the Wall Street Journal Europe in footnote 2 would suggest.

risk is significant.

It could be observed that our SUR estimation is inefficient when valid cross-equation restrictions can be imposed on our model. In fact, one could even think of imposing cross-equation restrictions on all the coefficients except the intercept, so as to have a dynamic panel estimation with fixed effects. 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. The Wald statistics reported in Table 5 clearly illustrate that the heterogeneity of coefficients in Tables 4.1 and 4.2 does not allow one to impose validly any set of cross-equation restrictions. The gain in efficiency generated by a panel estimation comes at the cost of the inconsistency of estimates, caused by invalid restrictions. Hence the evidence in Table 5 that panel estimation delivers strongly significant coefficients for all the variables we have considered, with a very similar pattern in the 5-year and the 10-year models, cannot be considered as a meaningful addition to the information provided by the SUR estimates of Tables 4.1 and 4.2. This underscores the important differences in the effects of risk and liquidity variables for the different sovereign issuers in our sample.

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.[10].

[10]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

However, it must be recognized that swap spreads are a special measure of risk, in that they include the counterparty risk of swap dealers 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 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.

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structure.

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 present a simple general equilibrium model. The model predicts that yield differentials should increase in both liquidity and risk, with an interaction term whose magnitude depends on the size of the liquidity differential with respect to the reference country.

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 crucial to detect their effect.

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 is quite complex, since it depends both on the current and on the future liquidity of the asset – with opposite signs. High illiquidity today decreases the impact of changes in volatility on the asset's price, but high illiquidity tomorrow exacerbates it because it makes it more risky in the light of a prospective "flight-to-quality". For less liquid markets, our theoretical analysis shows that the former effect always dominates: illiquidity dampens the price effect of consumption risk changes. However, the analysis also shows that for very liquid markets, the dampening effect of liquidity does not necessarily dominate. 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.

References

[1] Acharya, Viral V., and Lasse Heje Pedersen (2004), "Asset Pricing with Liquidity Risk," *Journal of Financial Economics*, forthcoming.

[2] Admati, R. Anat, and Paul Pfleiderer (1988), "A Theory of Intraday Patterns: Volume and Price Variability," *Review of Financial Studies* 1(1), 3-40.

[3] Amihud, Yakov and Haim Mendelson (1986), "Asset pricing and the bid-ask spread," *Journal of Financial Economics* 17, 23-249.

[4] Amihud, Yakov and Haim Mendelson (1991), "Liquidity, maturity, and the yields on U.S. Treasury securities," *Journal of Finance* 46, 1411-1425.

[5] Arora, Vivek and Martin Cerisola (2001), "How Does U.S. Monetary Policy Influence Sovereign Spreads in Emerging Markets?" *IMF-Staff-Papers* 48(3): 474-98.

[6] Barnes, Kevin and William Cline (1997), "Spreads and Risks in Emerging Markets Lending," Institute of International Finance Working Paper No.97-1.

[7] Biais, Bruno, Antoine Renucci and Gilles Saint-Paul, "Liquidity and the Cost of Funds in the European Treasury Bill Market," IDEI Woking Paper.

[8] Blanco Roberto (2001), "The Euro-Area Government Securities Market. Recent Developments and Implications for Market Functioning," Banco de Espana - Servicio de Estudios Working Paper 0120.

[9] Brennan, Michael J., and Avanidhar Subrahmanyam (1996), "Market microstructure and asset pricing: On the compensation for illiquidity in stock returns," *Journal of Financial Economics* 41, 441-464.

[10] Chordia, Tarun, Richard Roll and Avanidhar Subrahmanyam (2000), "Market liquidity and trading activity," *Journal of Finance* 56, 501-530.

[11] Codogno, Lorenzo, Carlo Favero, and Alessandro Missale (2003). "Government bond spreads", *Economic Policy* October 2003, 503-532.

[12] Datar, Vinay, Naik Narayan, and Robert Radcliff (1998), "Liquidity and stock returns: An alternative test," *Journal of Financial Intermediation* 1, 203-219.

[13] Daves, Philip R., and Michael C. Erhardt (1993), "Liquidity, reconstitution, and the value of U.S. Treasury strips," *Journal of Finance* 48(1), 315-329.

[14] Diamond, Douglas W., and Dybvig, Philip H. (1983), "Bank runs, deposit insurance, and liquidity," *Journal of Political Economy* 91(3), 401-419.

[15] Duffee G.R.(1998) "The Relation between Treasury Yields and Corporate Bond Yield Spreads", The Journal of Finance, 53, 6, 2225-2241

[16] Dungey, Mardi, Vance L. Martin, and Adrian R. Pagan (2000), "A Multivariate Latent Factor Decomposition of International Bond Yield Spreads," *Journal of Applied Econometrics* 15, 697-715.

[17] Dunne, Peter G., Michael J. Moore and Richard Portes (2002), "Defining Benchmark Status: An Application Euro-Area Bonds," CEPR Discussion Paper no. 3490, August.

[18] Eichengreen, Barry, and Ashoka Mody, (2000), "What Explains Changing Spreads on Emerging Market Debt?" in *Capital flows and the emerging economies: Theory, evidence, and controversies*, S. Edwards ed., NBER Conference Report series, Chicago and London: University of Chicago Press, 2000, 107-34.

[19] Eleswarapu, Venkat (1997), "Cost of transacting and expected returns in the Nasdaq market," *Journal of Finance* 52, 2113-2128.

[20] Ellul, Andrew, and Marco Pagano (2003), "IPO underpricing and after-market liquidity," CSEF Working Paper No. 99, http://www.dise.unisa.it/WP/wp99.pdf, *Review of Financial Studies*, forthcoming.

[21] Gallmeyer, Michael, Burton Hollifield and Duane Seppi (2004), "Liquidity Discovery and Asset Pricing," Carnegie Mellon University, unpublished manuscript.

[22] Goldreich, David, Bernd Hanke, and Purnendu Nath (2005), "The price of future liquidity: Time-varying liquidity in the U.S. Treasury market," *Review of Finance*, 9(1).

[23] Kamara, Avraham (1994), "Liquidity, taxes, and short-term Treasury yields," *Journal of Financial and Quantitative Analysis* 29(3), 403-417.

[24] Krishnamurthy, Arvind (2000), "The bond/old-bond spread," Northwestern University Working Paper.

[25] Lucas, Robert (1978), "Asset prices in an exchange economy," *Econometrica* 46(6), 1429 – 45.

[26] Pagano, Marco, and Ernst-Ludwig von Thadden (2004), "The European Bond Markets under EMU," *Oxford Review of Economic Policy* 20(4), 531-554.

[27] Strebulaev, Ilya (2001), "Market imperfections and pricing in the U.S. Treasury securities market," London Business School Working Paper.

[28] Vayanos, Dimitri (2004), "Flight to Quality, Flight to Liquidity, and the Pricing of Risk," NBER Working Paper No. W10327.

[29] Warga, Arthur, 1992, "Bond return, liquidity, and missing data," *Journal of Financial and Quantitative Analysis* 27, 605-617.

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

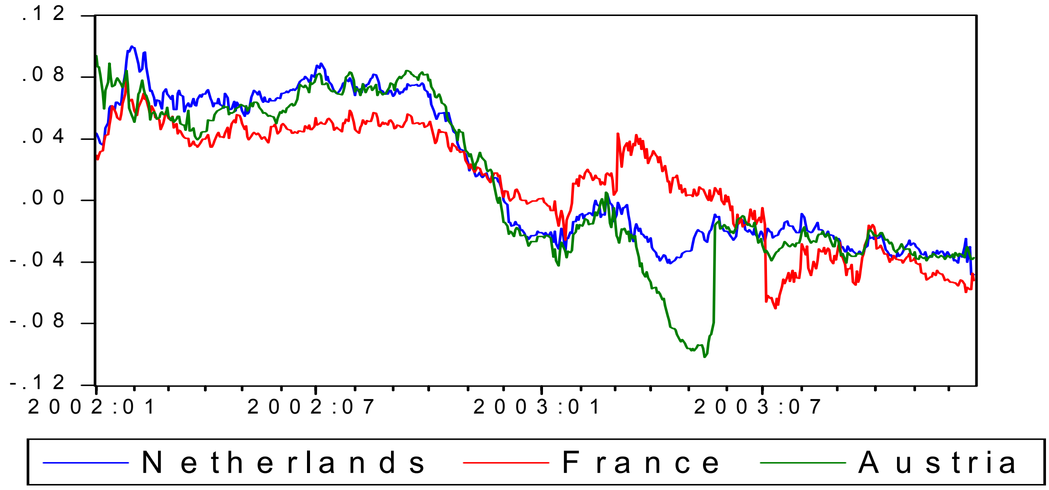
The table reports robustness analysis for the SURE system on 10-year yield differentials. We consider three alternative Risk Factors. R F 1 is the swap spread, R F 2 is the differential between yields on seasoned US BAA bonds and seasoned US AAA bonds (the source for these data is the FRED database), 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*	$-0.042^{\dagger}$
		(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*	$-0.10^{\dagger}$
		(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



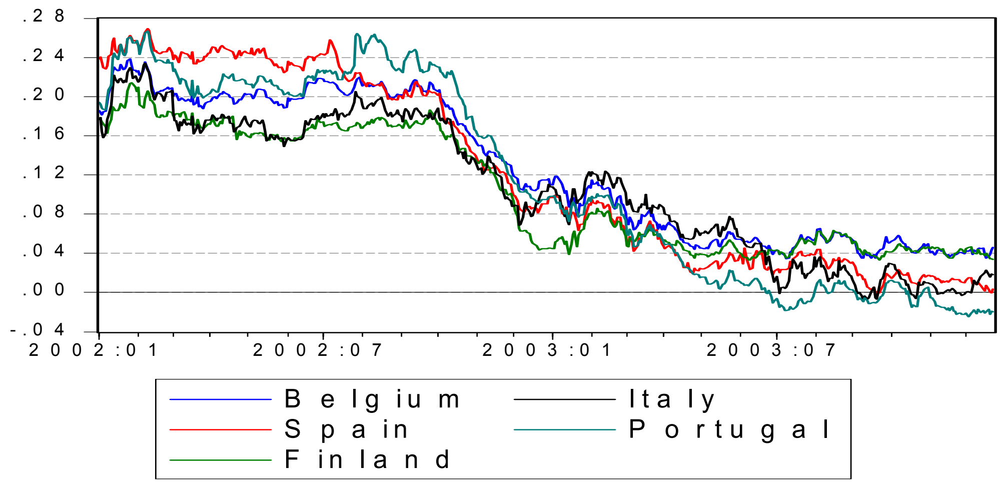


Figure 2. Bid-ask spread differentials in the Euro area, relative to German bonds (10-year benchmark bonds)

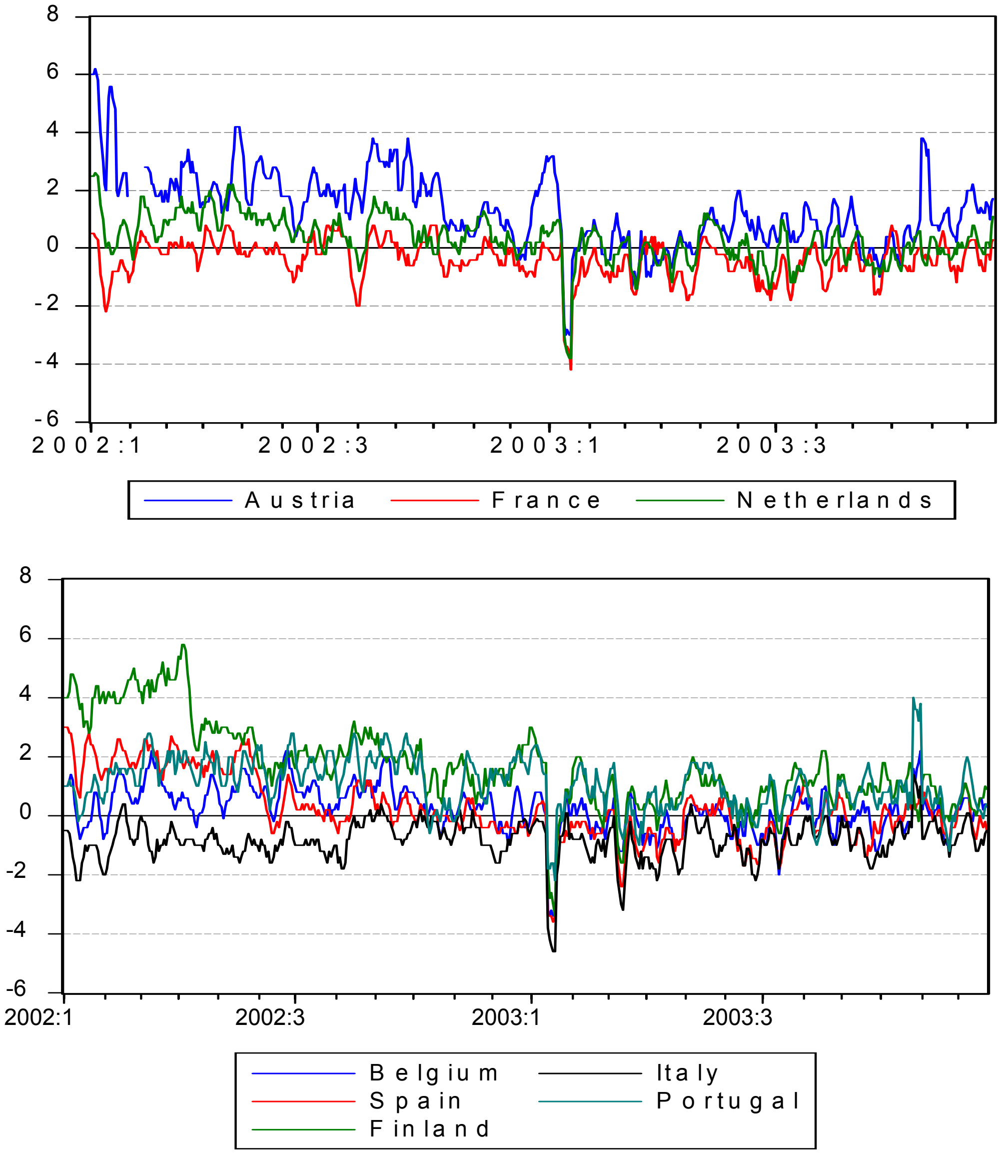


Figure 3. First principal components of Euro-area yield differentials and the spread between the 10-year fixed interest rate on swaps and US government bond yield