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EURO CORPORATE BOND RISK FACTORS

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SUMMARY

This paper investigates the determinants of credit spread changes in euro-denominated bonds. We adopt a factor model framework, inspired by the credit risk structural approach, as credit spread changes can be easily viewed as an excess return on corporate bonds over Treasury bonds. We try to assess the relative importance of market and idiosyncratic factors as an explanation of movements in credit spreads. We adopt a heterogeneous panel with a multifactor error model and propose a two-step estimation procedure, which yields consistent estimates of unobserved factors. The analysis is carried out with a panel of monthly redemption yields on a set of corporate bonds for a time span of 3 years. Our results suggest that the euro corporate market is driven by observable and unobservable factors. The unobservable factors are identified through a consistent estimation of individual and common observable effects. The empirical results suggest that an unobserved common factor has a significant role in explaining the systematic changes in credit spreads. However, in contrast to evidence regarding US credit spread changes, it cannot be identified as a market factor.

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Supporting information may be found in the online version of this article.

1. INTRODUCTION

The credit risk, or risk of default, of a bond arises for two reasons: both the magnitude and the timing of payoffs to investors may be uncertain. In other words, the risk of issuer default is accompanied by recovery rate uncertainty. The effects of default risks on prices depend on how the default event is defined and on the specification of the recovery in the event of a default. Considering this uncertainty, corporate bonds should offer higher yields than comparable default-free, i.e. government, bonds. Consequently, a corporate bond trades at a lower price than a corresponding (in terms of maturity and coupon) government bond. The difference between the yield on the risky bond and the yield on the corresponding default-free bond is called the credit spread. Theoretical credit risk models tackle the default risk in different ways. Structural models, in their most basic form, assume default the first time that some credit indicator falls below a specified threshold value. In Merton's model (Merton, 1974) default occurs at the maturity date of debt provided the issuer's assets are less than the face value of maturing debt at that time. Reduced-form models treat default as governed by a counting (jump) process coupled with an associated (possibly state-dependent) intensity process and thus whether or not an issuer actually defaults is an unpredictable event.

Several works deal with the empirical estimation of the structural models. Among others, Eom *et al.* (2003) use US market data to empirically test five structural models (Merton, 1974; Geske, 1977; Leland and Toft, 1996; Longstaff and Schwartz, 1995; Collin-Dufresne *et al.*, 2001) of corporate bond pricing. They clearly show that all five models considered have relevant spread prediction errors. In particular, all the models tend to underestimate the spread of higher-rated corporate bonds, while they overestimate the spread of bonds which are considered riskier.

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This paper studies the determinants of credit spread changes in the euro corporate bond market. In particular, we are interested in understanding the relevance of the implications of the structural credit risk models for the European market. Cross-sectional dependence, which is a common finding in the literature on delta credit spreads, is modeled by a heterogeneous panel data model with a multifactor error structure. This approach allows consistent estimation of the effect not only of the observed factors but also of the unobserved factors.

Collin-Dufresne *et al.* (2001) find that unobserved factors are a relevant component in American corporate bond data. They show that variables postulated in the structural approach have a rather limited explanatory power. They consider variables other than those prescribed by the structural models in order to capture other effects, such as the liquidity premium and the dynamic of interest rates. They adopt a heterogeneous parameter model for each issue and find out that the residuals from these regressions are highly cross-correlated. They conclude that a common systematic factor, identified as a local demand/supply shock, drives the credit spread changes.

Elton *et al.* (2001) move in a different direction. They stress the fact that credit spread changes are determined not only by credit risk but also by risk premium. Credit spread changes can easily be viewed as an excess return of corporate bonds over Treasury, i.e. risk-free bond proxy. Accordingly, they approach the problem in the framework of a traditional equity factor model to assess the influences of stock return common factors on credit spread. So far, most of the empirical works on credit spreads deal with US data, and relatively little is known about the extent to which these results apply to the euro market. Even though the empirical analysis of the US corporate bond market is an obvious reference, the European market is characterized by marked differences. While in Europe the bond market is dominated by government bonds and bonds issued by the financial intermediaries, the bond market in the USA is dominated by the non-financial corporate sector. In addition, municipal and agency bonds are major components of this market. Annaert and De Ceuster (2000), using data on aggregate index by rating categories and maturity buckets, stress that the European bond market shows broad similarities to the US market. However, their results are based on a rather limited time period (they consider daily data ranging from March 1998 to May 1999) when the Euro corporate bond market still lacked appropriate development. Houweling *et al.* (2005) analyze the excess yield on corporate bonds. They use several proxies to test whether liquidity is priced in the euro-denominated corporate bond market. Under both the assumptions on constant and time-varying liquidity premium, they find strong evidence of priced liquidity. De Jong and Driessen (2005) consider liquidity proxies of equity markets and show that returns on corporate bonds are correlated with market-wide fluctuations in the liquidity of the equity market.

In this paper, we first provide evidence that the methodology of Collin-Dufresne *et al.* (2001), as applied to the European delta credit spreads data, supports the idea that a set of unobserved factors influences the changes in the credit spreads. Moreover, in contrast to the conclusions of Collin-Dufresne *et al.* (2001), we find evidence that the significant unobserved component cannot be identified as a 'market factor'. Individual regressions show substantial parameter heterogeneity across bonds. In general, the presence of unobserved factors, as evidenced by the analysis of fitted residuals of both univariate regressions and the fixed-effects panel data model, suggests the adoption of a consistent econometric estimation procedure. When an unobserved common factor structure exists, the estimates of individual slope coefficients are inconsistent (see Coakley *et al.*, 2002, 2006; Pesaran, 2006; Kapetanios and Pesaran, 2007; Bai, 2009; Kapetanios *et al.*, 2011). The delta credit spreads are expressed as a function both of individual components and of observed and unobserved common factors, where the former are linearly dependent on the latter. This amounts to a heterogeneous panel data model, i.e. varying slope coefficients, with a multifactor error structure. In this framework, as studied in various papers (e.g. Coakley *et al.*, 2002, 2006; Pesaran, 2006; Bai, 2009), we can consistently estimate the effects of observed common and individual factors only if we take into account the relations between observed and unobserved factors, as shown in Pesaran (2006). We adopt a two-step procedure. First, the effects of observed factors are estimated by means of the Common Correlated Estimator, as proposed

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by Pesaran (2006). Second, the unobserved factors are then estimated by principal components analysis (see Bai, 2003). We show that the unobserved factors are consistently estimated in average norm.

We find that the variables suggested by the theory are in general both economically and statistically significant in explaining variations in individual issues' credit spreads. However, the factors predicted by the structural model are not as relevant as in the case of the US market. The analysis of the estimated residuals by means of the Bai and Ng (2002) information criteria suggests the presence of one unobserved common factor. Even though the estimated factor is not identified, we suspect that this is related to the market liquidity conditions. This confirms the conclusions of de Jong and Driessen (2005), that the European corporate bond excess holding returns have a significant exposure to liquidity risk, and that a liquidity premium helps to explain part of the 'credit spread puzzle', i.e. corporate bond yield spreads wider than what would be predicted by historical default losses.

Moreover, the existing literature presents strong empirical evidence of a liquidity premium on the European market analogously with what has been found for the USA.

The remainder of this paper is organized into a further eight sections. In Section 2 we discuss the meaning of the credit spread changes. Section 3 introduces the structural credit risk models. In Section 4 we illustrate the individual and common factors used in the analysis. Preliminary empirical evidence is presented in Section 5. The econometric model is introduced in Section 6. Section 7 describes the data. Results are discussed in Section 8. Section 9 concludes our findings.

2. DELTA CREDIT SPREAD AND EXCESS RETURNS

We define credit spread as the difference between the yield to maturity on a corporate bond and the yield to maturity on a government bond of the same maturity:

$$cs_t = c_t - g_t \tag{1}$$

where c_t is the redemption yield of a corporate bond at time t and g_t is the corresponding (i.e. with the same maturity) redemption yield on a government bond.[1] The return on a coupon bond j for a holding period equal to one is given by

$$r_{j,t} = \frac{\left(P_{j,t} + C_{j,t}\right) - P_{j,t-1}}{P_{j,t-1}}$$

where $P_{j,t}$ is the gross price at time t for bond j and $(C_{j,t})$ is the interest or coupon payments of bond j at time t. We employ the approximation developed by Shiller (1979) of the asset price as a function of the corresponding yield:

$$r_{j,t} \cong -d_{j,t}\left(b_{j,t} - b_{j,t-1}\right) \tag{2}$$

where $b_{j,t}$ is the redemption yield[2] of bond j at time t and $d_{j,t}$ is the modified duration of bond j at time t.[3]

[1] In particular, g_t is given by the redemption yield on the estimated euro government curve (see Section 5).

[2] The redemption yield of bond j at time t, $b_{j,t}$, equals the internal rate of return that discounts its cash flows, including the interest or coupon payments $(C_{j,t})$ and the repayment of principal (PC_j), back to the bond's current price, i.e. $P_{j,t} = \sum_{\tau=t}^{T-1} C_{j,\tau}\left(1 + b_{j,t}\right)^{-\tau} + \left(C_{j,T} + PC_j\right)\left(1 + b_{j,t}\right)^{-T}$.

[3] There are two measures of duration. Macauley duration (md) and modified duration (d). Macauley duration is the weighted average time to maturity of a bond, where the weights are given by the present values of the cash flows:

$$md_{j,t} = \sum_{\tau=t}^{T} \tau \times w_\tau = \sum_{\tau=t}^{T} \tau \times \frac{\frac{C_{j,\tau}}{\left(1+b_{j,t}\right)^\tau}}{\sum_{\tau=t}^{T} \frac{C_{j,\tau}}{\left(1+b_{j,t}\right)^\tau}}$$

The modified duration is simply the Macauley duration as defined above divided by $(1+b_{j,t})$.

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The modified duration is simply the Macauley duration as defined above divided by $(1+b_{j,t})$. Hence the modified duration indicates the percentage change in the price of a bond for a given change in yield. Using expression (2), the difference between the return on the corporate bond and the government position can be written as

$$r_{c,t} - r_{g,t} \cong -d_{c,t}(c_t - c_{t-1}) + d_{g,t}(g_t - g_{t-1}) \tag{3}$$

where $r_{g,t}$ and $r_{c,t}$ are the respective returns on the government and corporate bonds.

We know that if other factors are held constant, the lower the yield to maturity and the coupon rate, the higher the duration will be. In general, the corporate bonds have a higher coupon rate and a higher yield than a government bond with the same maturity. Hence the duration of government bonds can be thought of as the duration of corporate bonds plus, in general, a positive spread, $\gamma(t)$:

$$d_{g,t} = d_{c,t} + \gamma(t)$$

Then, expression (3) becomes

$$r_{c,t} - r_{g,t} \cong -d_{c,t}\delta_t + \gamma(t)(g_t - g_{t-1}) \tag{4}$$

where δ_t, i.e. the *delta credit spread*, is defined as

$$\delta_t = cs_t - cs_{t-1} \tag{5}$$

The second term on the right-hand side of (4) is negligible with respect to the excess return. Hence the excess return of corporate bonds over government bonds is proportional to the change in credit spread:

$$r_{c,t} - r_{g,t} \cong -d_{c,t}\delta_t \tag{6}$$

The delta credit spread represents a proxy for corporate bond excess loss, i.e. the return on a government bond minus the return on a corporate bond with the same maturity. Recently, empirical analysis and its practitioners have substantially changed their focus from bond yields to delta credit spreads. One example, as stressed by Collin-Dufresne *et al.* (2001), is represented by European mutual funds which invest both in corporate and in government bonds. As a consequence, their portfolios are extremely sensitive to changes in credit spreads and less so to changes in bond yields. Another example is represented by the hedge fund trading strategy of taking highly leveraged positions in corporate bonds while hedging away interest rate risk by shorting government bonds.

3. STRUCTURAL MODELS OF CREDIT RISK

The issuer of a fixed-income security might default prior to the maturity date. This means that both the magnitude and the timing of payoffs to investors may be uncertain. How these default risks affect corporate bond pricing depends on how the default event is defined and how recovery in the event of a default is specified. The pricing models can be classified into *reduced-form* models, those that are based on an assumed default intensity, and *structural* models, where there is an explicit characterization of the default event, such as the first time that a firm's assets fall below the value of its liabilities (Duffie and Singleton, 2003).

In this paper we refer to the structural-model approach and to risk premia theory in order to identify the main factors (individual and common) that drive credit spread changes. The seminal papers of Black and Scholes (1973) and Merton (1974) introduced the first model of the structural-form approach. In the Black–Scholes–Merton model we may think of equity and debt as derivatives with respect to the total market value of the firm, and as being priced accordingly. We are in the setting of standard Black–Scholes model, i.e. a market with continuous trading which is frictionless and

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Table I. Explanatory variables

Variable	Description	Predicted sign
Individual specific regressors		
cs	Beginning of month credit spread	–
Avgret	Average of daily excess return over preceding 180 days	–
Stdret	Standard deviation of daily excess return over preceding 180 days	+
rat	Rating	+
Dcsrat	Delta credit spread for rating	+
Dcsect	Delta credit spread for sector	+
Common factors		
5dss	5-year delta swap spread	+
Nofissue	Mv in number of issues in the IBOXX index	–
10Gov	Mv in 10-year German government benchmark rate	–
Slope	Mv of German government curve slope	–
Conv	Mv of German government curve convexity	
Upg	Mv in upgraded euro corporate bonds	–
Downg	Mv in downgraded euro corporate bonds	+
Mseuro	Morgan Stanley euro monthly return	–

Note: For the variable rating (*rat*), we assign a value to each rating class, from 10 (AAA) to 1. Mv stands for monthly variation.

competitive.[4] The original owners of the firm choose a capital structure consisting of pure equity and of debt in the form of a single zero-coupon bond maturing at time T, of face value D. In the event that the total value V_T of the firm at maturity is less than the contractual payment D due on the debt, the firm defaults, giving its future cash flows, worth V_T, to debt holders. The debt can then be viewed as a difference between a riskless bond and a Black–Scholes price of a European put option on the firm's asset. The option representation of the bond price implies that it is increasing in V and in D while it is decreasing in the riskless interest rate, in time-to-maturity, and in the firm's value volatility. The Black–Scholes–Merton model has important drawbacks.[5] In particular, it mainly focuses on the value and the capital structure of the firm, which is a difficult structure to represent. This objection apart, the structural approach provides an intuitive framework to determine the main factors that drive credit spread changes. In the next section, we present the set of variables used in the analysis of the euro corporate bond market, which are inspired by the structural-model approach (see Avramov *et al.*, 2007).

4. INDIVIDUAL AND COMMON FACTORS

The contingent-claim approach views debt as a combination of a risk-free loan and a short put option on the firm. Variables governing the firm-value process affect default probabilities and recovery rates, and ultimately drive credit spreads. Structural model variables typically include interest rates, term-structure slope, market return, market volatility, as well as firm leverage and volatility. Below we present the variables that we use in the empirical analysis and that are supposed to affect credit spread changes. In Table I we report the individual-specific and the observed common factors used in the analysis, along with the predicted sign effects on delta credit spreads.

[4] In detail, the agents are price-takers, there are no transaction costs, there is unlimited access to short selling and no indivisibilities of assets, and borrowing and lending through a money-market account can be done at the same riskless, compounded rate r.

[5] First, it requires inputs related to the value of firms that are often unavailable. Second, it allows default only at the maturity date of the bond. Third, it assumes independence between interest rates and credit risk. Last but not least, because it assumes that the value of the asset follows a geometric Brownian motion, the model implies that the default is predictable shortly before default. The first structural model has been widely improved by relaxation of some of its restrictive assumptions, for example, inter alia, Black and Cox (1976), Turnbull (1979), Leland (1994), Longstaff and Schwartz (1995), Briys and De-Varenne (1997) and Collin-Dufresne and Goldstein (2001).

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4.1. Common Factors

1. *Changes in the government bond rate level.* This variable represents both a proxy for the so-called *flight to quality* flows and a proxy for business cycle. From one point of view, a lower level of government rates implies a market preference for less risky asset, i.e. wider credit spreads. From the other, lower rates also imply a higher loan demand, which widens the credit spreads. Empirical evidence that a negative relationship exists between changes in credit spreads and interest rates has been shown by Longstaff and Schwartz (1995), Duffee (1998) and Collin-Dufresne *et al.* (2001). We use the DataStream's monthly series of the 10-year Benchmark German Treasury rates (denoted as *Gov*) to compute the monthly changes (denoted as *10Gov*), and the monthly variation of the squared 10-year Benchmark German Treasury rates ($10Gov^2$).
2. *Changes in the slope of the government yield curve.* This is a proxy for the movement in the supply and demand of government bonds. A flat term structure of interest rates reduces the incentives to invest in the government sector and therefore causes a corporate spread widening. Duffee (1998) tests this relation for the US corporate bond market. Moreover, a steepening of the term-structure slope implies an increase in expected future spot rates, thereby reducing credit spreads. The changes in the slope of the yield curve is given by the monthly changes in the difference between DataStream's 10-year and 2-year Benchmark German Treasury rates (denoted as *Slope*).
3. *Changes in the convexity of the government yield curve.* We also include the convexity of the government yield curve to capture potential nonlinear effects. This is calculated as the monthly changes in the difference between the 5-year German Treasury rate and the average of the 10-year and the 2-year Benchmark German Treasury rates (denoted as *Conv*).[6]
4. *Changes in liquidity.* Collin-Dufresne *et al.* (2001) stressed the fact that the corporate bond market tends to have relatively high transaction costs and low volumes. These findings suggest that it would be appropriate to check for the existence of a liquidity premium. Given that the corporate bond market is an over-the-counter market, the standard measures of liquidity are unavailable. Following Houweling *et al.* (2005), we consider several proxies to measure variations in liquidity: the monthly change in the 5-year Euro swap spread (denoted as *5dss*), the monthly variation in the number of issues of the corporate bonds included in the IBOXX Euro Bond Index, (*nofissue*). Furthermore, considering that liquidity can be correlated with return volatility, we include the squared index monthly return (*ret2*), as suggested by Hong and Warga (2000).[7] We expect that a change in swap market liquidity would reflect a change in the same direction in corporate market liquidity, because of the strong link between swap and corporate markets.[8] A decrease in swap market liquidity (i.e. in the corporate market) implies a market preference for less risky assets. Hence we expect the factor loading of this liquidity proxy to be positive. Second, the issued amount of a bond is often assumed to give an indication of its liquidity. When the liquidity of the corporate bond market increases, corporate bond prices increase and credit spreads decrease.[9] Since we do not observe the issued amount for each bond, we consider a market liquidity measure such as the monthly variation in the number of issues of the corporate bonds included in the IBOXX Euro Bond Index.
5. *Changes in the business climate.* Even if the probability of default remains constant for a firm, changes in credit spreads can occur due to changes in the expected recovery rate. The expected recovery rate, in turn, should be a function of the overall state of business climate. We use the monthly return on Morgan Stanley Euro Index (denoted as *MSeuro*) as a proxy for the overall state of the economy.

[6] The data source is DataStream.
[7] Though ARCH modeling can be adopted here to estimate the conditional variance as a measure of volatility, there is no evidence of ARCH effects in the monthly index return.
[8] The issuers of corporate bonds typically fund on the swap market. Thus, if swap spreads widen, the long-term funding costs of corporate bond issuers should increase, and investor demand for credit bonds should decrease. Assuming a constant supply of bonds, the decline in demand for credit products will cause prices to decline and the spread to Treasury to widen (see Collin-Dufresne *et al.*, 2001).
[9] Houweling *et al.* (2005) provide an extensive survey on both the theoretical structure and the empirical applications in which the issued amount is considered as a liquidity proxy.

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6. *Credit spread.* To investigate the presence of a mean-reverting behavior in credit spreads, we include the beginning-of-month level of credit spread (denoted as *Spread*).
7. *Changes in credit quality.* Changes in credit quality, which include downgrading or upgrading in rating, is a part of credit risk. A general process of improvement or worsening in credit quality should inversely move the credit spreads. We proxy the change in credit quality by monthly changes in rating downgrading (denoted as *Downg*) and upgrading (denoted as *Upg*) of the Merrill Lynch Global High Grade Corporate Index.[10]

4.2. Firm-Specific Factors

1. *Mean and standard deviation of daily excess return of firm's equity.* These variables summarize the firm-level risk and return. Equity data reflect up-to-date information on firm value and should anticipate bond prices.[11] An increase in the equity daily excess return means higher firm profitability. In line with the analysis of Kwan (1996) and of Campbell and Taksler (2003), we expect stock returns to have a negative effect on credit spreads. It is well known that the equity volatility of a firm increases its probability of default. Hence a firm's volatility should drive up the yields on corporate bonds and widen the credit spreads.[12] Following Campbell and Taksler (2003), we match bond data with equity data to explicitly evaluate the effects of equity volatility on corporate bond yield spreads. The daily excess return of each firm's equity is computed relative to the Morgan Stanley Index of the country where the stock is exchanged.[13] For each firm's equity, we compute the mean (denoted as *Avgret*) and standard deviation (*Stdret*) of daily excess returns over the 180 days prior to (not including) the bond trade.
2. *Changes in credit market factors.* We test whether credit spread changes depend on bond characteristics such as rating and the industrial sector. Each bond is assigned to one of the IBOXX sub-indices on the basis of the bond's beginning-of-month rating or sector. We consider four rating categories (AAA, AA, A, BBB) (denoted as *Dcsrat*) and three industrial sectors (Industrial, Financial, Utility), and for each sub-index we consider the index monthly delta spread (*Dcsect*).

We do not use accounting variables to explain the credit spread changes. This choice is driven by two considerations. First, accounting data generally have either quarterly or yearly frequency. We think that interpolating the data does not provide so much information on credit spread changes. Second, most of the studies which use accounting variables do not find any statistical evidence of their explanatory power, and conclude that they are unlikely to explain the observed movements in credit spreads.

5. PRELIMINARY EMPIRICAL EVIDENCE FOR THE EUROPEAN CORPORATE BOND MARKET

The important conclusions of Collin-Dufresne *et al.* (2001) for the US corporate bond market seem to be a logical starting point for any empirical study of the European corporate bond market. In this section, we try to replicate their analysis with data on euro-denominated corporate bonds. The goal is to understand to what extent the European market resembles the American market, at least as regards the factors that affect the delta credit spreads.

[10] We exclusively take into account euro-denominated bonds, and the monthly changes are computed with respect to the index par amount. The data come from the Merrill Lynch Index Rating Migration Databook. This databook summarizes relevant information on the composition of the main Merrill Lynch Corporate Bond Indices.

[11] Ederington *et al.* (1987) claim that all data going into ratings prices should be anticipated by equity prices. Moreover, they argue that investors fully anticipate rating changes which almost never affect bond returns.

[12] Houweling *et al.* (2005) assume that bonds issued by companies whose equities are listed on a stock market are more liquid. Therefore, our sample should contain corporate bonds with higher liquidity and lower yields with respect to the full sample (see Section 7).

[13] We eventually consider the following Morgan Stanley indices: Msci Emu, Msci Denmark, Msci Finland, Msci Norway, Msci Sweden, Msci Switzerland, Msci Uk, Msci Usa, Msci Canada, Msci Japan And Msci Hong Kong.

They start from a simple model where the delta credit spread of each bond i at time t, y_{it}, depends on common observed factors, $\boldsymbol{d}_t$, and individual specific components, $\boldsymbol{x}_{it}$:

$$y_{it} = \boldsymbol{\alpha}_i^{'}\mathbf{d}_t + \boldsymbol{\beta}_i^{'}\mathbf{x}_{it} + e_{it} \quad t = 1, \ldots, T \quad i = 1, \ldots, I$$

In line with their analysis, we consider three different specifications. Nevertheless, we have to mention that for the European market we lack some of the data available for the US market. However, we attempt to follow their specification approach as closely as possible. They found that nearly half of the variation in spreads was unaccounted for by their regressors.

The first specification includes the following common factors ($\boldsymbol{d}_t$):

- the monthly change in the German government slope (*Slope*);
- the monthly change in 10-year German government bond yield-to-maturity (*10Gov*);
- the monthly change in the German convexity (*Conv*);
- the monthly return on Morgan Stanley Euro Index (*MSeuro*);

and the individual factors ($\boldsymbol{x}_{it}$):

- the average of daily excess equity return over the preceding 180 days (*Avgret*);
- the standard deviation of daily excess equity return over the preceding 180 days (*Stdret*).

The first model specification is

$$\begin{aligned} y_{it} \; &= \alpha_{1i} + \alpha_{2i}\mathrm{Slope}_t + \alpha_{3i}10\mathrm{Gov}_t + \alpha_{4i}\mathrm{Conv}_t + \alpha_{5i}\mathrm{MSeuro}_t \\ &+\beta_{1i}\mathrm{Avgret}_{it} + \beta_{2i}\mathrm{Stdret}_{it} + e_{it} \quad t = 1, \ldots, T \quad i = 1, \ldots, I \end{aligned} \tag{7}$$

The second specification includes additional explanatory variables to control for possibly omitted systematic common factors:

- the spread of the IBOXX euro corporate bond index at time t−1 (*Spread*);
- the 5-year delta swap spread (*5dss*);
- the monthly variation in the total issued amount of the IBOXX index (*Totaos*);
- the monthly variation in the square level of 10-year German government benchmark yield-to-maturity. ($10Gov^2$);
- the level of 10-year German government benchmark yield-to-maturity at time t-1 (*Gov*);
- monthly variation in upgraded euro corporate bonds (*Upg*);
- monthly variation in downgraded euro corporate bonds (*Downg*);
- the level of VDAX index at time *t-1* (*Vdax*);
- monthly variation in the VDAX index (*Dvdax*).

The VDAX index is a volatility index of the DAX options traded at the Eurex. The 5-year delta swap spread ($5dss_t$), and the monthly variation in the total issued amount of the IBOXX index ($Totaos_t$), are two liquidity proxies. The second model specification is

$$\begin{aligned} y_{it} \; &= \alpha_{1i} + \alpha_{2i}\mathrm{Slope}_t + \alpha_{3i}10\mathrm{Gov}_t + \alpha_{4i}\mathrm{Conv}_t + \alpha_{5i}\mathrm{MSeuro}_t \\ &+\alpha_{6i}\mathrm{Spread}_{t-1} + \alpha_{7i}5\mathrm{dss}_t + \alpha_{8i}\mathrm{Totaos}_t + \alpha_{9i}10\mathrm{Gov}_t^2 + \alpha_{10i}Gov_{t-1} \\ &+\alpha_{11i}\mathrm{Upg}_t + \alpha_{12i}\mathrm{Downg}_t + \alpha_{13i}\mathrm{VDAX}_{t-1} + \alpha_{14i}\mathrm{DVDAX}_t \\ &+\beta_{1i}\mathrm{Avgret}_{it} + \beta_{2i}\mathrm{Stdret}_{it} + e_{it} \quad t = 1, \ldots, T \quad i = 1, \ldots, I \end{aligned} \tag{8}$$

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Table II. Average adjusted R^2

	(7)	(8)	(9)
Rating			
AAA	0.242	0.190	0.266
AA	0.209	0.366	0.478
A	0.161	0.378	0.463
BBB	0.171	0.307	0.414
Industrial sector			
Financials	0.172	0.394	0.505
Industrials	0.177	0.338	0.435
Utilities	0.182	0.295	0.35
Maturity bucket			
Short (1–4 years)	0.193	0.402	0.487
Medium (4–10 years)	0.183	0.323	0.438
Long (>10 years)	0.142	0.352	0.44
Overall	0.176	0.354	0.451

Note: The adjusted R^2 of equations (7), (8), and (9) are computed and averaged by rating, sector and maturity bucket, and overall.

Finally, the third specification adds to the observed common factors what Collin-Dufresne *et al.* (2001) call a 'market factor' for the corporate bond market, i.e. the monthly change in the IBOXX BBB Index credit spread ($Iboxxbbb_t$):

$$\begin{aligned} y_{it} = {} & \alpha_{1i} + \alpha_{2i}\text{Slope}_t + \alpha_{3i}10\text{Gov}_t + \alpha_{4i}\text{Conv}_t + \alpha_{5i}\text{MSeuro}_t + \alpha_{6i}\text{Spread}_{t-1} \\ & + \alpha_{7i}5\text{dss}_t + \alpha_{8i}\text{Totaos}_t + \alpha_{9i}10\text{Gov}_t^2 + \alpha_{10i}\text{Gov}_{t-1} + \alpha_{11i}\text{Upg}_t \\ & + \alpha_{12i}\text{Downg}_t + \alpha_{13i}\text{VDAX}_{t-1} + \alpha_{14i}\text{DVDAX}_t + \alpha_{15i}\text{Iboxxbbb}_t \\ & + \beta_{1i}\text{Avgret}_{it} + \beta_{2i}\text{Stdret}_{it} + e_{it} \quad t = 1, \ldots, T \quad i = 1, \ldots, I \end{aligned} \tag{9}$$

Each specification includes the intercept and is estimated by OLS. We aggregate the bonds by maturity buckets, rating categories and industrial sectors. The individual regressions show considerable parameter heterogeneity.[14]

Table II reports the average adjusted R^2 of the specifications above. There are some differences from the results obtained by Collin-Dufresne *et al.* (2001) for the US market. Most of the explanatory variables have some ability to explain the delta credit spreads, and most of the estimated parameters are in line with the predictions of economic theory, but the explanatory power of all the specifications put together is slightly lower than that found for the US market. Collin-Dufresne *et al.* (2001) find an average adjusted R^2 of 21, 35 and 55% for their specifications, while we find overall an adjusted R^2 of 18%, 35% and 45%, respectively (Table II). Collin-Dufresne *et al.* (2001) found that the unexplained component of the movement in credit spread changes can be ascribed to the presence of a single common factor. Examining the cross-section correlation in the residuals, they showed that controlling for the influence of a market factor dramatically reduces the correlation among the fitted residuals. In particular, the percentage of the total residual variance explained by the first principal component drops from about 76% to about 40%. They conclude that the 'market factor' can be identified as a supply/demand shock.

We replicate their analysis in the case of the European market. We divide the residuals into nine bins, determined by three maturity groups (Short term: <4 years; Medium term: ≥4 years and <10 years; Long term: ≥10 years) and three industrial sectors (Financial, Industrial and Utility).[15] The

[14] For ease of exposition we do not report the individual regressions results, but they are available upon request from the authors.
[15] The results are very similar when we divide the residuals into bins based on maturity and rating.

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Table III. Principal component analysis

	PC_1	$PC_1 + PC_2$
First specification (7)	64.9	83.2
Second specification (8)	56.4	82.0
Third specification (9)	53.8	73.7

Note: The OLS residuals of regression equations (7), (8), and (9) are divided into nine bins. For each bin cross-section averages are computed so that the covariance matrix of the nine bins can be obtained. Principal components of cross-section average residuals are then calculated. Percentage of total variation in the average residuals as explained by the first two principal components (PC_1 and PC_2) is reported (see Section 5).

Table IV. Cross-section dependence

A	
% PC_1 for y	44.53%
% PC_2 for y	17.92%
B	
	Average pair-wise correlation
Delta credit spread	0.40
OLS residuals	0.36
Fixed-effect residuals	0.12
CD stat	55.40
	(0.00)

Note: Panel A reports the proportion of delta credit spread variability explained by the first and second principal component, i.e. PC_1 and PC_2. The first row of panel B shows the average of all pair-wise cross-section correlation coefficients of delta credit spread. 'OLS residuals' and 'Fixed effect residuals' are the average of all pair-wise correlation coefficients of the least-squares residuals from the individual linear regressions in the panel and of the FE residuals of the third specification (9), respectively. CD Stat is the Pesaran (2004) cross-section dependence test statistic. *P*-value is in parentheses.

estimates of individual bond regressions suggest that the estimated parameters are characterized by heterogeneity, both at the individual bond level and at bin level.[16] Table III shows that the first component explains a relevant part of the variability in the residuals (64.9%, 56.4%, and 53.8%, for the first, second and third specifications respectively) for all the specifications considered, while the analysis of the US market shows that when a 'market factor' is added the first component accounts for a small fraction of the remaining variation. In our case, the residuals' variability is still high (53.8%) even when we introduce a proxy for the market into the regression (9). Moreover, Collin-Dufresne *et al.* (2001) show that in the first two specifications the first principal component can be seen as an equally weighted portfolio across the categories used to build the bins. In the European case, there can be potentially more than one unobserved factor which influences the variation in credit spreads. Moreover, this analysis does not provide any clear clue to the identification of the unobserved factors. While Collin-Dufresne *et al.* (2001) use the evidence on the effects of the inclusion of the 'market factor' to support the presence of an unobserved common factor, we cope with this problem in a different way.

From the average correlation of the corporate bond delta credit spread, shown in Panel B of Table IV, it is evident that the delta credit spreads are cross-correlated. The first two principal components account for 62% of the total variance, and the correlation of the residuals of the second specification (8) is quite high. This seems to confirm the presence of substantial cross-section dependence. Table IV

[16] This contrasts with Elton *et al.* (2001), who for the US market show that there is a substantial homogeneity in the estimated parameters.

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also reports the test statistic for cross-dependence by Pesaran (2004).[17] The hypothesis that the residual credit spread changes are cross-sectionally independent is strongly rejected. All this suggests the adoption of a panel-data approach to the analysis of delta-credit spread cross-section dependence that evolves over time. Table IV reports the average correlations of fitted residuals from a fixed-effect model of third specification (equation (9)). Again, the average cross-section correlation of fixed-effects fitted residuals is far from negligible. This suggests that there could be some omitted explanatory variables which must be found among non-firm-specific factors.

6. THE MODEL

Following Pesaran (2006), we consider a linear heterogeneous panel data model where y_{it} is the observation on the delta credit spread at time t for the ith issue for $i = 1, 2, \ldots, I$ and $t = 1, 2, \ldots, T$:

$$\begin{aligned} y_{it} &= \boldsymbol{\alpha}_i^{'}\mathbf{d}_t + \boldsymbol{\beta}_i^{'}\mathbf{x}_{it} + e_{it} \\ e_{it} &= \boldsymbol{\gamma}_i^{'}\mathbf{f}_t + \varepsilon_{it} \end{aligned} \tag{10}$$

where $\mathbf{d}_t$ is a $n \times 1$ vector of observed common effects, $\mathbf{x}_{it}$ is a $k \times 1$ vector of observed individual specific regressors, $\mathbf{f}_t$ is the $m \times 1$ vector of unobserved common factors and ε_{it} is the idiosyncratic error assumed to be independently distributed of $\mathbf{d}_t, \mathbf{x}_{it}$. As in Pesaran (2006), we suppose that the individual specific factors are correlated with common (observed and unobserved) factors through

$$\mathbf{x}_{it} = \mathbf{A}_i^{'}\mathbf{d}_t + \boldsymbol{\Lambda}_i^{'}\mathbf{f}_t + \mathbf{v}_{it} \tag{11}$$

where $\mathbf{v}_{it}$ are the specific components of $\mathbf{x}_{it}$ distributed independently of the common effects and across i. The factor loading matrices $\mathbf{A}_i$ and $\boldsymbol{\Lambda}_i$ have fixed and bounded components. We hypothesize that Assumptions 1, 2, 5a and the rank condition (equation (21) in Pesaran (2006)) hold. Moreover, we assume:

Assumption 1. The common factors $(\mathbf{f}_t, \mathbf{d}_t)$ are orthogonal, $E\left(\mathbf{f}_t\mathbf{d}_t^{'}\right) = \mathbf{0}, \forall t$.

Assumption 2. The slope coefficients $\boldsymbol{\alpha}_i$ and $\boldsymbol{\beta}_i$ follow the random coefficient model

$$\boldsymbol{\alpha}_i = \boldsymbol{\alpha} + \boldsymbol{\xi}_i, \;\; \boldsymbol{\xi} \sim \text{i.i.d.}(0, \Omega_\xi), \;\; i = 1, 2, \ldots I \tag{12}$$

$$\boldsymbol{\beta}_i = \boldsymbol{\beta} + \boldsymbol{\nu}_i, \;\; \boldsymbol{\nu}_i \sim \text{i.i.d.}(0, \Omega_\nu), \;\; i = 1, 2, \ldots I \tag{13}$$

where $\boldsymbol{\xi}_i$ and $\boldsymbol{\nu}_i$ are independent with $\| \boldsymbol{\alpha} \| < C$, $\| \boldsymbol{\beta} \| < C$, $||\Omega_\xi|| < C$ and $||\Omega_\nu|| < C$, where C is a finite positive constant. Ω_ξ and Ω_ν are $(n \times n)$ and $(k \times k)$ symmetric positive definite matrices, respectively. The random deviations, $\boldsymbol{\xi}_i$ and $\boldsymbol{\nu}_i$, are distributed independently of $\boldsymbol{\gamma}_j$, $\boldsymbol{\Gamma}_j$, ε_{jt} and $\mathbf{v}_{jt}$ for all i,j and t.

Assumption 3. The factor loadings $\boldsymbol{\gamma}_i$ are treated as parameters. In particular: $||\boldsymbol{\gamma}_i|| \leq \bar{\gamma} < \infty$, and $\boldsymbol{\Gamma}' = (\boldsymbol{\gamma}_1, \ldots, \boldsymbol{\gamma}_I)$ is such that $\left(\frac{\boldsymbol{\Gamma}'\boldsymbol{\Gamma}}{I}\right)^{-1}$ exists.

[17] Pesaran (2004) proposes a test for cross-section dependence based on a simple average of the all pairwise correlation coefficients of the ordinary least square (OLS) residuals from the individual regressions in the panel. This test is applicable to a variety of panel data models and, despite the Breusch and Pagan LM test, it can be used when the cross-section dimension is large relative to the time series dimension. The cross-section dependence statistic (CD stat) is computed as $\text{CD} = \sqrt{\frac{2T}{I(I-1)}}\left(\sum_{i=1}^{I-1}\sum_{j=i+1}^{I}\hat{\rho}_{ij}\right)$ where $\hat{\rho}_{ij}$ is the sample estimate of the pair-wise correlation of the residuals ($\hat{e}_i$ and $\hat{e}_j$). Under the null hypothesis of no cross-section dependence, the CD statistic is distributed (as I and $T \to \infty$ with no particular order) as a standard normal distribution.

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Under Assumptions 1, 2, 5a and the rank condition in Pesaran (2006), and Assumptions 2 and 3 above, both the Common Correlated Effect (CCE) estimator and the CCE mean group (CCEMG) estimator are consistent (see Theorems 1 and 2 in Pesaran, 2006). Asymptotic normality of CCE is obtained under the further assumption that $\sqrt{T}/I \to 0$. The rank condition guarantees the consistency of the CCE estimator of $\boldsymbol{\beta}_i$ (Pesaran, 2006). However, as shown in Pesaran and Tosetti (2011, Remark 2), CCE continues to be applicable even though the rank condition is not satisfied. When this happens, the mean of the slope parameters $\boldsymbol{\beta}_i$ can be consistently estimated and their asymptotic distribution can be obtained if it is further assumed that the unobserved factor loadings are i.i.d. across i and of ε_{jt}, v_{jt} and $\mathbf{f}_t$, $\mathbf{d}_t$ for all i, j, t (see also Kapetanios *et al.*, 2011). Assumption 1 is needed to avoid an identification problem that arises when estimating $\hat{\mathbf{e}}_i$ in (16) (Pesaran and Tosetti, 2011, make a similar assumption in the case of spatial correlation and common factors).

6.1. Estimation

We are interested in estimating the effects of observed common and individual components on the credit spread changes. However, it is also relevant for understanding the determinants of delta credit spread to determine the number of the unobserved common factors and to estimate these too. To this end, we propose an estimation procedure which is articulated in two steps. First, the individual slope coefficients, $\boldsymbol{\alpha}_i$ and $\boldsymbol{\beta}_i$, are estimated by the CCE estimator (Pesaran, 2006); second, a consistent (in average norm) principal component estimate of the unobserved factors is obtained (see Proposition 1). The number of factors is assumed to be unknown but fixed.

The CCE estimator is obtained by augmenting the OLS regression of y_{it} on $\mathbf{x}_{it}$ and $\mathbf{d}_t$ with the cross-section averages $\bar{\mathbf{z}}_t = \frac{1}{I}\sum_{i=1}^{I}\mathbf{z}_{it}$:

$$\hat{\boldsymbol{\beta}}_i = \left(\mathbf{X}_i'\mathbf{M}\mathbf{X}_i\right)^{-1}\mathbf{X}_i'\mathbf{M}\mathbf{y}_i \tag{14}$$

where $\mathbf{y}_i = (y_{i1}, \ldots, y_{iT})'$ and $\mathbf{X}_i = (\mathbf{x}_{i1}, \ldots, \mathbf{x}_{iT})'$ and

$$\mathbf{M} = \boldsymbol{I}_T - \mathbf{H}\left(\mathbf{H}'\mathbf{H}\right)^{-1}\mathbf{H}' \tag{15}$$

$\mathbf{H} = (\mathbf{D}, \bar{\mathbf{Z}})$, $\mathbf{D}$ and $\bar{\mathbf{Z}}$ being, respectively, the $T \times n$ and $T \times (k+1)$ matrices of observations on $\mathbf{d}_t$ and $\bar{\mathbf{z}}_t$. Although $\bar{y}_t$ and ε_{it} are not independent (i.e. endogeneity bias), their correlation goes to zero as $I \to \infty$.

We estimate the model (10)–(11) with the hypothesis that the observed common factors are orthogonal to the unobserved ones, i.e. $E\left(\mathbf{f}_t\mathbf{d}_t'\right) = \mathbf{0}, \forall t$. In order to deal with error cross-section dependence due to unobserved common factors we adopt the following procedure:

1. We consistently estimate the slope parameter $\hat{\boldsymbol{\beta}}_i$ by means of the CCE estimator of equation (14), based on an estimate of $\mathbf{f}_t$ by means of cross-section averages, $\bar{\mathbf{z}}_t$, and $\mathbf{d}_t$.
2. We estimate the residuals as

$$\hat{e}_i = \mathbf{M}_d\left(\mathbf{y}_i - \mathbf{X}_i\hat{\boldsymbol{\beta}}_i\right) \tag{16}$$

where $\mathbf{M}_d$ is given by

$$\mathbf{M}_d = \boldsymbol{I}_T - \mathbf{D}\left(\mathbf{D}'\mathbf{D}\right)^{-1}\mathbf{D}'$$

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The presence of unobserved common factors correlated with the individual specific regressors does not cause the inconsistency of the parameter estimates of the observed common effects part ($\boldsymbol{\alpha}_i$), given by

$$\hat{\boldsymbol{\alpha}}_i = \left(\boldsymbol{D}'\boldsymbol{D}\right)^{-1}\boldsymbol{D}'\left(\mathbf{y}_i - \mathbf{X}_i\hat{\boldsymbol{\beta}}_i\right) \tag{17}$$

The proof of the consistency of $\hat{\boldsymbol{\alpha}}_i$ is reported in the supplementary Appendix B, available online as supporting information. A similar result is independently obtained by Pesaran and Tosetti (2011) in a heterogeneous panel data model with spatially dependent idiosyncratic errors. This implies that the fitted residuals $\hat{\mathbf{e}}_i$ are consistent for $\mathbf{e}_i$.

3. The unobserved common factors are estimated, up to a non-singular transformation (i.e. *rotation indeterminacy*), by the method of least squares. The estimator of $\boldsymbol{F}$ is equal to the first J eigenvectors associated with the first J largest eigenvalues of the matrix $\hat{\boldsymbol{E}}\hat{\boldsymbol{E}}'$, where $\hat{\boldsymbol{E}} = (\hat{\mathbf{e}}_1, \hat{\mathbf{e}}_2, \ldots, \hat{\mathbf{e}}_I)$ is a $(T \times I)$ matrix. In order to consistently estimate the number of factors, we make use of the information criteria proposed by Bai and Ng (2002). Thus by the definition of eigenvalues and eigenvectors $\hat{\boldsymbol{F}}$ satisfies

$$\left[\frac{1}{IT}\hat{\boldsymbol{E}}\hat{\boldsymbol{E}}'\right]\hat{\boldsymbol{F}} = \hat{\boldsymbol{F}}\boldsymbol{V}_{IT} \tag{18}$$

where $\mathbf{V}_{IT}$ is a diagonal matrix which consists of the J largest eigenvalues of $\hat{\boldsymbol{E}}\hat{\boldsymbol{E}}'$ arranged in decreasing order.[18] When we look at the asymptotic characteristics of the estimated factors we assume that m is known.[19]

In order to prove the the average (norm) consistency of $\hat{\boldsymbol{F}}$ for $\boldsymbol{F}$, we make the following assumptions:

Assumption 4.

1. $E[\|F_t\|^8] < \infty$.
2. $sup_i E[\|\varepsilon_i\|^8] < \infty$.
3. $sup_i E\frac{\|\mathrm{X}_i\|^4}{T^2} = O(1)$.
4. $sup_i E\left\|\left(\frac{\mathrm{X}_i'\mathrm{M}\mathrm{X}_i}{T}\right)^{-1}\right\|^4 = O(1)$.

The moment conditions (Assumptions 4(i), 4(ii)) ensure that the estimation error of the individual slope coefficients possesses bounded fourth-order moments.

Proposition 1. Suppose that Assumptions 1, 2, 5a and the rank condition in Pesaran (2006) and Assumption 4 hold. Let $\mathbf{W} = \left(\boldsymbol{\Gamma}'\boldsymbol{\Gamma}/I\right)\left(\mathbf{F}'\hat{\mathbf{F}}/T\right)\mathbf{V}_{IT}^{-1}$. Then $\mathbf{W}$ is an $(m \times m)$ invertible matrix, and

$$\frac{1}{T}\left\|\hat{\mathbf{F}} - \mathbf{F}\boldsymbol{W}\right\|^2 = \frac{1}{T}\sum_{t=1}^{T}\left\|\hat{f}_t - \boldsymbol{W}'\mathbf{f}_t\right\|^2 = O_p(1/min(I,T))$$

In supplementary Appendix we provide the proof (B) and a Monte Carlo exercise (C), which shows the small-sample properties of the estimated factors.

[18] The scaling by IT does not affect $\hat{\mathrm{F}}$.
[19] Their asymptotic distributions are not affected when the number of factors is unknown and is estimated (see Bai, 2003).

7. DATA DESCRIPTION

Our corporate bond data are extracted from the IBOXX Euro Bond Index. This index is issued by seven major investment banks.[20] Each bank is due to buy and sell every single asset belonging to the index. The index bond prices are determined by the following criteria. First, the highest and the lowest prices are excluded, and the price is subsequently given by the average of the other five prices. Moreover, each asset included should have at least 500 million euros of amount outstanding and its time to maturity should be greater than 1 year. Such criteria should guarantee that only tradable and liquid bonds are included in the analysis. In this way we try to reduce the liquidity premium of the Euro corporate market. The IBOXX database[21] contains issue- and issuer-specific variables such as callability, maturity, coupon, industrial sector, rating, subordination level, issuer country, duration and several measures of credit spread.

We eliminate all the bonds downgraded to high-yield debt, as our goal is to explain the behavior of investment grade euro corporate bonds.[22] The bonds under consideration have standard cash flows – fixed-rate coupon and principal at maturity. We exclude all unrated bonds, step-up notes, floating rate debt and convertible bonds. We also exclude bonds with call options, put options or sinking fund provisions. Moreover, we require issuers with publicly traded stock in order to estimate equity volatility and equity excess return.[23] Last, in order to undertake principal component analysis of the residuals, we restrict our sample to a balanced panel. We only take into account those issues which are continuously included in the index from the last observation backward. We end up with 207 bonds for 33 monthly observations. The sample extends from March 2002 to November 2004. We use the fitted government curve spread provided by the IBOXX database. This spread is equal to the difference between the yield-to-maturity of the corporate bond and the corresponding (i.e. with the same maturity) yield-to-maturity on the estimated euro government curve.[24] The use of 'estimated redemption yield spread' only makes sense if the approximated corporate bond prices are truly close to the observed one. This is not generally the case in the Euro bond market. In fact, no matter what technique is used to interpolate (Nelson–Siegel, Cubic Spline with five knots) the results are quite poor. In the supplementary Appendix A, we present some evidence concerning the magnitude of the estimation errors of redemption yield spread based on estimated corporate spot rates.

Table V presents summary statistics on the bonds and issuers in the sample. Given the reduction of the sample to match the equity data and to deal with a balanced panel dataset, one may wonder if these bonds are representative of the overall Euro corporate market. A comparison of our sample to all non-callable and non-putable bonds included in the IBOXX index for the period considered (on average about 374 issues) suggests that they are very close. Table VI compares bonds in the sample with bonds included in the IBOXX index. The two samples have very similar distribution across credit ratings (panel A) and industrial sectors (panel B). The distribution across maturity bucket of our sample has a slight tendency toward medium- and short-term bonds (panel C). However, the average bond maturity in our sample is very close to the average bond maturity of the full sample (5.66 in Table V).

Although the criteria of the IBOXX index should guarantee the liquidity of their components, Table V shows that the full sample contains outliers. The standard deviation of the full sample is twice

[20] ABN AMRO, Barclays Capital, BNP Paribas, Deutsche Bank, Dresdner Kleinwort Wasserstein, Morgan Stanley and UBS Investment Bank.

[21] The database was built by the optimization group at Fideuram Investimenti SGR, Milan.

[22] During the sample period considered, 12 issuers and 24 issues downgraded to high-yield bonds. Two issuers and four issues defaulted.

[23] We exclusively consider the corporate bonds issued by firms listed in the Morgan Stanley World All Country Index. The data source is DataStream.

[24] The euro government curve is estimated by a cubic spline. Moreover, only German and French government bonds enter the term structure estimation process.

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Table V. Summary statistics

	Mean		SD	
	Sample A	Sample B	Sample A	Sample B
Credit spread change	−1.58	5.55	22.61	44.02
Coupon (%)	5.55	5.37	0.74	0.92
Years to maturity	5.21	5.66	2.28	3.45
	Min.		Max.	
	Sample A	Sample B	Sample A	Sample B
Credit spread change	−492.20	−740.20	465.70	2529.80
Coupon (%)	3.50	2.13	7.25	9.75
Years to maturity	0.94	0.92	14.07	29.94
Equity volatility	1.98%	0.62%	0.00%	6.59%
Equity excess return	−0.10%	0.36%	−1.95%	1.27%

Note: Summary statistics of the corporate bonds both in the selected sample (sample A) and in the IBOXX corporate bond index (sample B), i.e. all the non-putable and non-callable corporate bonds included in the IBOXX index (see Section 7). The credit spread changes are measured in basis points.

Table VI. Sample composition for rating and sector

A		
Rating	% sample A	% sample B
AAA	1.80%	5.42%
AA+	0.37%	1.39%
AA	3.35%	5.29%
AA-	13.37%	12.46%
A+	16.35%	15.25%
A	12.88%	14.99%
A-	19.81%	15.00%
BBB+	14.58%	12.67%
BBB	12.87%	12.02%
BBB-	4.63%	5.50%
B		
Industrial sector	% sample A	% sample B
Financials	38.16%	37.92%
Industrials	49.76%	50.80%
Utilities	12.08%	11.28%
C		
Maturity bucket	% of sample	% of full sample
Short (1–4 years)	35.53%	34.84%
Medium (4–10 years)	62.08%	60.06%
Long (> 10 years)	2.39%	5.10%

Note: See note to Table V.

our sample standard deviation. The maximum monthly credit spread change is about 466 basis points for our sample and 2530 basis points for the full sample. Therefore the extra return of a corporate bond with respect to a government bond can be 25% in a month, if we consider the full sample.

8. RESULTS

Following the estimation procedure outlined in Section 6, first we estimate the slope parameters α_i and $\boldsymbol{\beta}_i$ in the model (10), with the variables in specification A (see Table VII), by means of the CCE estimator in (14) and (17). Second, from the estimated variance–covariance matrix of the consistently estimated residuals in (16), we calculate the principal components as in (18). According to the information criteria

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Table VII. Regression results

	A	B
cons	−16.381	−9.757
	(−1.355)	(−0.880)
cs	−0.185	−0.155
	(−12.382)	(−12.090)
avgret	−925.144	−968.940
	(−1.837)	(−2.130)
stdret	1228.086	819.269
	(2.035)	(1.580)
dcsrat	0.742	0.708
	(7.151)	(8.120)
5dss	−3.522	
	(−0.833)	
nofissue	−46.669	
	(−1.842)	
iret2	11.982	
	(0.326)	
10gov	4.347	2.442
	(1.685)	(1.090)
slope	−6.389	−5.000
	(−1.762)	(−1.590)
conv	−18.061	−14.758
	(−2.094)	(−1.370)
upg	0.226	0.127
	(0.775)	(0.460)
downg	0.438	0.737
	(2.400)	(2.590)
mseuro	−15.746	−18.993
	(−1.678)	(−2.060)
Wald test	417.32	400.26
	(0)	(0)
Test parameter constancy	5858.91	5694.35
	(0)	(0)
Pesaran–Yamagata Test	8.47	9.82
	(0)	(0)
Observations	6831	6831

Note: Estimates of two random coefficient models. The abbreviations for the variables used in the regressions are presented in Table I. The Pesaran and Yamagata (2008) test statistic for the null hypothesis of slope homogeneity is reported. t-statistics for parameter estimates and p-values appear in parentheses.

IC_p of Bai and Ng (2002), the euro delta credit spread presents just one systematic unobserved risk factor (see Table VIII).

We estimate two different specifications, denoted in Table VII by columns A and B. Specification A includes all the regressors (individual and observed common factors); specification B

Table VIII. Information criteria for common factors

No. of factors	IC_{p1}	IC_{p2}	IC_{p3}
1-4 1	8.06	8.06	8.04
2	8.12	8.13	8.10
3	8.23	8.24	8.19
4	8.40	8.42	8.35
5	8.27	8.30	8.21
6	8.34	8.37	8.27
7	8.51	8.54	8.42
8	8.59	8.63	8.50

Note: The reported figures refer to the information criteria of Bai and Ng (2002). The factors are computed with the principal component estimator using the residuals of specification A in Table VII.

excludes the non-significant liquidity proxies (the 5-year delta credit spread (*5dss*), the monthly variation in the number of issues of the corporate bonds included in the IBOXX Euro Bond Index (*nofissue*), and the squared index monthly return (*ret2*)). The test of slope homogeneity in the version proposed by Pesaran and Yamagata (2008), where the cross-section dimension could be large relative to the time series dimension, shows that the null hypothesis of parameter constancy is strongly rejected.

In general, the variables suggested by the theory are both economically and statistically significant in explaining variations in individual issues' credit spreads:

- *Changes in the government bond rate level*. The monthly variation in 10-year German government benchmark yield-to-maturity (*10gov*) is not significant at the 5% significance level.
- *Changes in the slope of the government yield curve*. The slope in the German government yield curve is significant (10% significance level) and has a negative impact on the delta credit spread. That is, when the curve is flattening this increases the credit spreads. This is in accordance with the findings of Duffee (1998) and of Collin-Dufresne *et al.* (2001).
- *Changes in the convexity of the government yield curve*. When liquidity proxies are included, the convexity of the government yield curve has a negative and statistically significant effect on the change in delta credit spreads. This captures possible nonlinearities in the relation between delta credit spreads and yield curve movements.
- *Changes in liquidity*. The liquidity proxies included in specifications A and B are not significant at the 5% significance level. This suggests that these proxies are inadequate to catch the influence of liquidity conditions on the movements of delta credit spreads. However, this does not exclude that liquidity is a relevant factor in explaining the movements in the excess returns over corporate bonds.
- *Mean and standard deviation of daily excess return on equity*. The standard deviation of daily excess return (*stdret*) over the preceding 180 days, a volatility proxy, is significant for specification A, i.e. when the non-significant liquidity proxies are not included in the regression.
- *Changes in credit quality*. It is particularly interesting that while the change in rating downgrade is always strongly significant, the change in rating upgrade is not. It should be noted that the effect of the monthly variation in the total amount outstanding of the upgraded bonds is smaller than the variation in the downgraded ones. This suggests that shocks in the credit market have an asymmetric effect on delta credit spreads.
- *Changes in the business climate*. The Morgan Stanley Euro Index price return is significant in both the cases considered. The sign of the estimated coefficient is in accordance with the theory, which says that the market sentiment has a positive impact on the excess returns of corporate bonds.
- *Changes in credit market factors*. The monthly delta spread of the IBOXX sub-indices, based on the bond's beginning-of-month rating classification, to which each bond issue in the sample is assigned, is strongly significant.[25] This could be interpreted as a market factor which explains a large part of the variation in the delta credit spreads. An increase of 100 basis points in this credit market factor augments the delta credit spreads by about 70 basis points.
- *Credit spread*. The initial credit spread level is negative and strongly significant in both the regressions; this is in accordance with a mean-reverting behavior of the delta credit spreads (the same is found by Collin-Dufresne *et al.*, 2001).

In order to interpret the estimated risk factor, we repeat the analysis carried out in Section 5. We run regressions for each bond issue with the same explicative variables used in specification A (Table VII) with and without the estimated factor. Then we compute the principal components from the residuals,

[25] The results obtained with the IBOXX sub-indices for industrial sectors are less significant, so we choose to retain the indexes for rating categories alone.

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Table IX. Correlation of delta credit spreads with the estimated factor

Rating	Average partial correlation
AAA	0.30
AA	0.36
A	0.40
BBB	0.49

Note: The average partial correlation of delta credit spreads with the estimated factor, controlling for the explanatory variables contained in Table VII specification A.

and we find that the percentage of total variation attributed to the first principal component drops from about 64%, computed from the model residuals without including the estimated factor, to 45% when we include the estimated factor among the regressors. Thus the inclusion of the estimated factor seems to account for a common unobserved component. Moreover, the average correlation coefficient between actual and fitted values, as obtained by the inclusion of the estimated factor in each bond regression, along with the explanatory variables of specification A, is about 0.53. From the empirical distribution, 55% of estimated correlations are larger than 0.5.

Our guess is that the estimated factor accounts for *latent* liquidity effects. This is reinforced by the observation that the liquidity proxies are unable to control for liquidity distortions. We consider different liquidity proxies and observe that none has a significant effect. Moreover, we argue that the liquidity distortions are possibly induced by the presence of imperfections in the euro corporate bond market. This idea comes mainly from the evidence, stressed in Section VII, that corporate bonds in the euro market could be mispriced (see Table A1 in the supplementary Appendix A).

In order to understand the extent to which the factor is correlated with credit quality, we calculate the average partial correlation between the delta credit spreads and the estimated factor, after having controlled for all the explanatory variables contained in Table VII, specification A. Table IX reports, as expected, that the average partial correlation increases as the credit rating decreases, i.e. as the liquidity conditions worsen. This is close to what was found by de Jong and Driessen (2005). They show that both US and European corporate bonds are exposed to systematic liquidity shocks, and that a liquidity risk premium helps to explain part of the credit spread. Importantly, the liquidity exposure is larger for lower-rated bonds. In conclusion, we think that an aggregate factor potentially driving liquidity in the bond market could be correlated with the estimated common factor, in line with the findings of Collin-Dufresne *et al.* (2001).

9. CONCLUSION

In this paper we investigate the determinants of credit spread changes denominated in euros. We point out that the change in credit spreads can be viewed as a proxy of the excess return of the corporate bonds over government bonds. For this reason we conduct our empirical analysis in a factor model framework. We also follow a data-driven approach recently developed for the US market which addresses the question of which variables are most correlated with credit spread movements. Notable differences from the American market emerge. First, the estimated parameters seem to be quite heterogeneous across the bonds and bins used in the analysis. Second, the unexplained part of the movements in the delta credit spreads, which is due to unobserved common factors, is not correlated with the market. Nonetheless, we find highly cross-correlated residuals from the single-bond regressions. This suggests a heterogeneous panel data model with a multifactor error structure. In this setup, we distinguish observed and unobserved common factors, and in order to consistently estimate the influences of individual factors we adopt a recently developed estimator (Pesaran, 2006). Starting

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from these estimates, we show that the unobserved factors can be consistently (in average norm) estimated. Overall, our analysis shows that a systematic risk factor exists in the euro corporate bond market, and that this factor is independent of the main common factors predicted by the theory. The estimated factor can be thought of as capturing the liquidity bias, which in turn can be caused by the lack of a fully developed market. This interpretation also seems to be supported by the price misalignments found in the euro corporate bond market.

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