Illiquidity Spillovers: Theory and Evidence From European Telecom Bond Issuance

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ABSTRACT

In a study of the European telecommunication-sector bond market, we find empirical evidence that a firm's new bond issue can temporarily inflate yield spreads of other bonds in its sector. We show that this effect seems unrelated to new fundamental information about the bond's issuer. Our results imply that an issuance of 15.5 billion Euros by Deutsche Telekom temporarily depressed the mark-to-market value of 100 billion Euros in outstanding European telecom debt by approximately 273 million Euros. This study is supported and motivated by a stylized model of a risk-averse liquidity-provider in which supply shocks, such as new issues, place price pressure on correlated securities.

JEL Classification: G12, G14.

Keywords: Bond issuance, Illiquidity spillovers, European telecom bonds.

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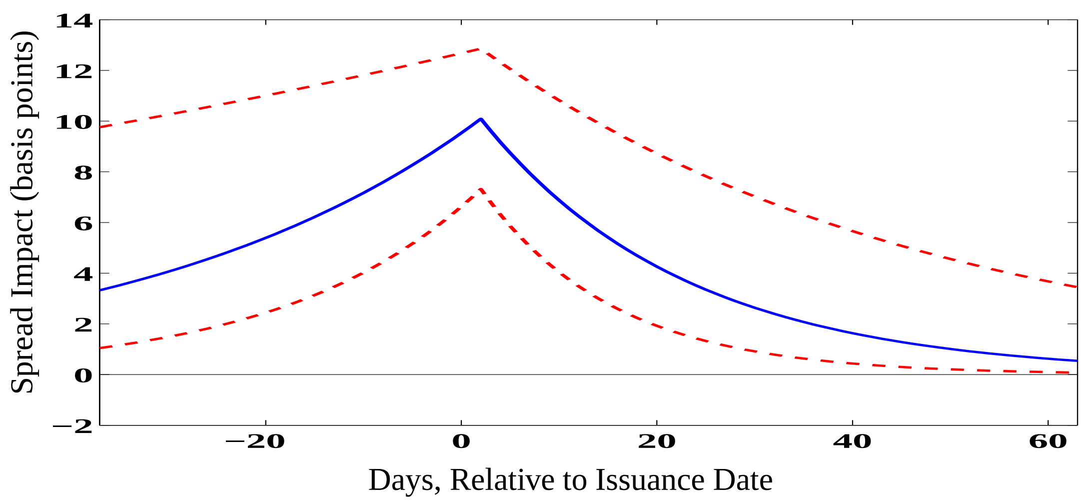
This paper provides an empirical analysis of the market-wide impact of security issuance on asset prices, based on a case study of European telecom debt issuance. Controlling for credit risk, we find that sector-wide yield spreads rise temporarily in response to debt issues by firms in this sector. This yield-spread impact, which appears to be unrelated to new fundamental information, has an estimated half life of fifteen days. This study is supported and motivated by a model of a risk-averse liquidity-provider with imperfect ability to locate long-term buyers of the issue, causing a temporary issuance price impact that spills over to correlated assets.

As an example, consider Deutsche Telekom's June 2000 bond issue of 15.5 billion Euros. We estimate that in the week of issuance, this issue raised yield spreads across the European telecom sector by 10.1 basis points, as illustrated in Figure 1, whose construction is explained in Section 2. In comparison, the mean bid-ask spread in this sector during our sample period was 2.6 basis points. The estimated mark-to-market impact of this issue across the 100-billion-Euro telecom bond market, while temporary, is 273 million Euros (0.28%).

Figure 1. The Impact of Debt Issuance on the Yield Spreads of Other Bonds



The solid line is the estimated impact of debt issuance on the yield spreads of other bonds (scaled to reflect the impact of Deutsche Telekom's issuance of 15.5 billion Euro of debt in June of 2000). The dashed lines are two-standard-error confidence bands.

As a result of the sensitivity of yield spreads to sector-wide debt issuance, investors in bonds of a given industry face the possibility that a bond issuance in that industry may temporarily depress the market value of their holdings. We estimate that 71% of the variance of changes in yield spreads over weeks of issuance can be attributed to the cross-market impact of issuance. This "new-issuance" risk factor is distinct from interest-rate risk, from the issuer's default risk, and from normal risk premia for bearing this default risk. The transience of the yield-spread impact, indicated in Figure 1, makes new-issuance risk of greater concern to investors with short holding periods, and to others sensitive to marking to market, such as leveraged hedge funds, than to buy-and-hold investors, such as pension funds or insurance companies. To our knowledge, we are the first to detect illiquidity-spillover effects associated with supply shocks in asset markets.[1]

Issuers are subject to the risk that simultaneous, or nearly simultaneous, issuance by competitors would raise their costs of funds. Rational managers may therefore choose to defer planned debt issuance – or preemptively issue debt – in anticipation of their rivals' issues. For example, *The Financial Times* reports[2] that Telecom Italia was "squeezed out by other telecom deals" and thus delayed a 10-billion-Euro issuance. Indeed, the distribution across time of issuance during our sample period is consistent, according to a test explained in Section 4, with issuers who try to separate themselves in time. (One could, however, argue that other institutional considerations motivate firms to separate issues in time.) The associated issuance game is left for future research.

In summary, we study the impact of debt issuance on sector-wide yield spreads. We propose a supporting theory of illiquidity spillovers based on imperfect intermediation, and check whether the behavior of the European telecom debt market near the times of debt issues is consistent with our theory.

Available debt-pricing models do not explicitly account for supply effects. In structural models,[3] default occurs when the firm's asset falls below some "default boundary," often the face value of debt. Reduced-form models[4] abstract from the direct causes of default, which is instead modeled as an arrival with an intensity that may depend only exogenously on firm-specific or macro variables. Neither family of models has been extended to treat new-issuance impacts. Similarly, empirical models of yield spreads do not provide an explicit role for issuance.[5]

[1]Amihud, Mendelson, and Lauterbach (1997) find positive liquidity spillovers across related stocks in reaction to improvements in the trading mechanism.

[2]See "Telecom Italia Seeks 5 Billion Euros," by R. Bream and A. Van Duyn, *The Financial Times* , March 28, 2001.

[3]This family of models was instigated by Black and Scholes (1973) and Merton (1974), and extended by Black and Cox (1976), Geske (1977), Fisher, Heinekel, and Zechner (1989), and Leland (1994), among many others.

[4]This class of models includes Litterman and Iben (1991), Jarrow and Turnbull (1995), Jarrow, Lando, and Turnbull (1997), Lando (1998), Madan and Unal (1998), and Duffie and Singleton (1999).

[5]Recent examples of this literature are Collin-Dufresne, Goldstein, and Martin (2001), Elton, Gruber, Agrawal, and Mann (2001), and Janosi, Jarrow, and Yildirim (2002). It should be noted, however, that two studies conjecture that supply shocks may affect yield spreads. Collin-Dufresne, Goldstein, and Martin (2001) investigate the determinants of credit-spread changes. They find that fitted credit-spread residuals are mostly driven by a single common factor, and conjecture that this missing factor is supply and demand shocks. Duffie and Singleton (1997), in a study of U.S. swap spreads, observe that a substantial fraction of swap-spreads variation is left unexplained by their model, and suggest that swap-market-specific supply and demand shocks may be driving the unexplained spread changes.

We propose the following mechanism for the price impact of new debt issues. In many markets, liquidity-providing intermediaries, such as underwriters, trading desks of investment banks, hedge funds, and other effective liquidity providers, initially absorb significant portions of a new issue of corporate securities. These intermediaries may hold, temporarily at least, significant amounts of other bonds in the same sector. A risk-averse intermediary's incentive to hold these correlated assets falls due to his newly acquired position in the issued bond. We hypothesize that liquidity-providing intermediaries face delays in their attempts to find and negotiate with suitable buy-and-hold investors. Our search-based asset-pricing model is in the spirit of Duffie, Gârleanu, and Pedersen (2003). Over time, intermediaries sell their excess holdings of the issued bond to buy-and-hold investors, such as pension funds, insurance companies, other institutional investors, and high-net-worth individuals. As the intermediary offloads the new issue, his incentive to hold sector-specific assets returns to pre-issuance levels, after controlling for changes in other explanatory factors, and the spread impact of the issuance on the correlated bonds decays. A decline in the market prices of these correlated bonds precedes a scheduled issuance through anticipation. More risky, and more highly correlated, assets are more severely affected by the issue. In short, the issuance places temporary price pressure on the issued bond and on correlated bonds. We refer to these effects as the *same-bond* impact of issuance and the *other-bond* impact of issuance, respectively. In Section 1, we formalize this intuition in a stylized model. This illiquidity-spillover theory compliments research on systematic illiquidity, such as Amihud, Mendelson, and Lauterbach (1997), Duffie, Gârleanu, and Pedersen (2003), and Weill (2003a).

The European telecom sector provides an ideal setting for estimating issuance-related illiquidity spillovers in a market with relatively low return volatility, compared to equity markets. Between October 1999 and July 2001, net debt issuance in this sector surpassed 175 billion Euros, a 300% increase over the 60 billion Euros of debt outstanding in September 1999, representing 46% of all European non-financial corporate debt issued during that period. *The Economist* reports[6] that these issues supported bids for government-auctioned cellular bandwidth licenses.

Our empirical model, motivated by our theory, uses the quantity of issuance as an explanatory variable for changes in the yield spreads of European telecom bonds. Our issuance measure is the product of an issue's market value with its duration, which is approximately the reduction in the issue's market value due to a 100-basis-point parallel shift in the term structure of the issuer's credit spreads. While this issuance measure does not perfectly capture all sources of unhedge-able risk to investors, we discuss its relative advantages in Section 2.2.

[6] See "A $250 Billion Gamble: The Telecom Sector has Overreached Itself," *The Economist*, January 25, 2001.

We estimate a time-series model explaining yield-spread changes with leads and lags of this issuance measure, controlling for changes in the issuing firm's leverage, the issuing firm's equity returns, changes in the slope of the risk-free term structure, and changes in the short-maturity interest rate. Due to possible endogeneity, explained in Section 2.3, we exclude the yield-spread changes of newly issued bonds from this estimation. We estimate two versions of this model. In one version, the impact of issuance on yield spreads is constrained to have a parametric dependence, inspired by our theory, on time from issue; the second version is unconstrained. For both versions, the estimated dependence of yield-spread changes on issuance of other bonds supports our theory, in that it is economically and statistically significant, transitory, and peaks on the week of issuance. This finding is robust to alternative model specification. Riskier bonds are more strongly affected by new issues. Issues of more credit-risky debt have a stronger effect, per unit of issuance, on other-bond yield-spreads, as do larger issues.

To the extent that equity returns reasonably control for fundamental information about the issuer that may be revealed during the issuance process, the issuance effect that we document does not appear to be related to such information. Two other characteristics of this issuance effect point away from an information-based explanation: (i) The effect is transitory, and (ii) The effect peaks on the day of issuance, not on the day of announcement. In Section 4, we discuss alternative information-based explanations for our empirical results, such as the impact of fundamental information, industry debt capacity, and investor risk aversion.

In a related analysis, described in Section 3.3, we characterize a component of yield-spread changes that is unique to newly issued bonds. For these bonds (which were not used in the estimation of our empirical model), we subtract the yield-spread changes predicted by our empirical model from the observed yield-spread changes. Using the sample average of prediction residuals, we reject the null hypothesis that this "same-bond issuance component" is zero. We find that the corresponding impact on yield-spreads levels is decreasing in the time since issuance, consistent with one's natural conjecture regarding the same-bond impact of issuance.

These issuance-related price impacts could potentially be exploited by arbitrageurs. Indeed, the behavior of our modeled liquidity provider is consistent with that of an arbitrageur. This "arbitrage," however, involves risk. The estimated Sharpe ratios associated with a strategy that exploits the other-bond effects of issuance, even ignoring transaction costs, is only 0.34.

Our work augments the literature examining same-asset supply effects in equity markets, which is concerned with separating the impact of fundamental information embedded in supply shocks from the impact of the supply shocks themselves. Some authors find patterns of price pressure and attribute them to temporary imbalances in supply and demand arising from market segmentation. Scholes (1972), Harris and Gurel (1986) and Mitchell, Pulvino, and Stafford (2004) find price pressure in the reaction of equity prices to large block trades, to changes in the composition of the S&P 500 Index, and to mergers, respectively. A second group of authors contends that supply shocks cause permanent movements along downward-sloping demand curves. This group includes Shleifer (1986) and Wurgler and Zhuravskaya (2002), in studies of changes in the composition of the S&P 500 index, Ofek and Richardson (2000), in a study of IPO lock-up expirations, and Kaul, Mehrotra, and Morck (2000), in a study of technical changes in the construction of the TSE 300 index. Other authors find a combination of transitory and permanent reactions to supply shocks. Among these are Kraus and Stoll (1972) and Holthausen, Leftwich, and Mayers (1990), in studies of large-block trades; Hess and Frost (1982), in a study of seasoned equity offerings; and Chen, Noronha, and Singal (2004), in a study of S&P 500 Index recompositions. Our study appears to be the first to document the spillover of price pressures to related assets. Prior studies have focused on equity markets. Bonds, which have less volatile returns than do equities, are more likely to reveal issuance spillover price impacts. Even relative to bond markets, our "laboratory" for this study, the European telecom sector between October 1999 and July 2001, is especially well suited to the task, given the large size and high number of issues.

The remainder of the paper is organized as follows: In Section 1, we describe the empirical setting and provide a theoretical model of the price impacts of issuance. Section 2 describes the data and presents a regression model relating yield-spread changes to issuance. Section 3 reports the empirical findings. Section 4 discusses the results.

1. A Model of the Price Impact of Issuance

We consider an economy with a risk-averse liquidity-providing intermediary, who immediately absorbs a significant portion of the issuance of a security, and a continuum of equally risk-averse long-term investors, who obtain positions in the issued security as they are contacted by the intermediary. The intermediary supplies immediacy to the issuer, and other market participants, in the spirit of Demsetz (1968). Both agent types have a time preference rate r that is equal to the continuously compounded risk-free rate on perfectly liquid money-market assets, and a coefficient of constant absolute risk aversion α.

This economy has two equally risky assets. The intermediary normally maintains a constant exogenously determined inventory, s, of each asset, for example in order to satisfy market-making demands.[7] However, only asset 1 exists before time 0. At time $-T$, for some $T > 0$, it becomes known that asset 2 will be issued at time 0. On issuance day, the intermediary will absorb $s+S$ units of asset 2, targeting s units as market-making inventory, and seeking to unload the "excess" S units to long-term investors. Let $q_1(t)$ and $q_2(t)$ denote the intermediary's inventories in assets 1 and 2, respectively, at time t. In the equilibrium that we describe, we have $q_1(t) = s$ for all t, $q_2(t) = 0$ for $t < 0$, $q_2(0) = s+S$, and will shortly derive a model of $q_2(t)$ for $t > 0$.

In a contrasting Walrasian model, markets would clear immediately, and an intermediary with a large inventory of securities would instantly find buyers. In practice, however, underwriters, marketmakers, and other liquidity providers temporarily absorb supply imbalances, eventually unloading their positions as they locate longer-term investors. This intuition can be captured in a search-based approach to asset prices, introduced in Duffie, Gârleanu, and Pedersen (2003). See, also, Vayanos and Wang (2002) and Weill (2003a). We assume that the intermediary and the long-term investors face search frictions, hindering their ability to trade quickly.

In our equilibrium, any of the finite quantity (continuum) of potential investors, would, if contacted by the intermediary, purchase a bond. The intermediary's allocation S of the new issue is assumed to be sized for one bond per investor. After time 0, the quantity of yet-to-be-served investors is thus $S-(q_2(0)-q_2(t)) = q_2(t)-s$. Technological constraints preclude the intermediary from contacting the investors simultaneously. Instead, the intermediary and a given investor make contact at a Poisson arrival time with a constant search intensity λ. Assuming that these contacts are independent across pairs of investors, and that certain measurability assumptions are satisfied (conditions are given by Sun (2000), p. 18, Theorem C), the Law of Large Numbers implies that the total contact-and-sale rate is, almost surely,[8] the search intensity λ multiplied by the quantity $q_2(t)-s$ of yet-to-be-served investors. Thus, after time 0,

$$\dot{q}_2(t) = \frac{\mathrm{d}q_2(t)}{\mathrm{d}t} = -\lambda\big(q_2(t)-s\big). \tag{1}$$

Solving Eqn. (1),

$$q_2(t) = s \ + \ S\,\mathrm{e}^{-\lambda t}, \qquad \text{for } t \geq 0. \tag{2}$$

Let $R_i\big(q_1(t), q_2(t), t\big)$ denote the intermediary's time-t reservation value for asset i, for $i \in \{1,2\}$, and let Z_1 and Z_2 be a long-term investor's reservation values for assets 1 and 2, respectively.

[7] Amihud and Mendelson (1980) suggest that dealers target constant inventory levels. This hypothesis was supported by Madhavan and Smidt (1993) and Hasbrouck and Sofianos (1993) in studies of NYSE dealers and by Hansch, Naik, and Viswanathan (1998) in a study of dealers at the London Stock Exchange.

[8] We henceforth suppress the "almost surely" qualification.

We will calculate[9] $R_i\big(q_1(t), q_2(t), t\big)$ and Z_i in equilibrium. Long-term investors are assumed, for simplicity, to be buy-and-hold traders, who are passive except at their one (at most) encounter with the intermediary, and have a prior holding of one unit of asset 1. Thus, when meeting an intermediary they diversify their portfolio by purchasing one unit of asset 2 (since the two assets are equally risky), and their reservation values for the bonds are independent of the intermediary's inventory.[10] We let $s > 1$, so gains from trade are guaranteed at every meeting between a long-term investor and the intermediary.

We assume, for simplicity, that the long-term investor gets a fixed fraction, b, of the gains from trade. As a result, the time-t trade price of asset i, $V_i(t)$, is

$$V_i(t) = b\, R_i\big(q_1(t), q_2(t), t\big) + (1-b)\, Z_i, \tag{3}$$

which is an equilibrium outcome of the simultaneous-offer bargaining game described in Kreps (1990). While there are other equilibria in which the bargaining power b depends endogenously on the outside options of investors and intermediary, the thrust of our results depends mainly on non-zero bargaining power for the investor, and we avoid the complexities that a more detailed analysis of the bargaining setting would entail.

1.1. Equilibrium Price Impact of Issuance

Suppose that the two risky assets are consol bonds, with correlated default times, that pay coupons at unit rate until default. We assume zero recovery at default, for simplicity. For each bond, default occurs at the first arrival of a Poisson process with a constant intensity. Specifically, the default times are $T_1 = \min(\tau_A, \tau_C)$ and $T_2 = \min(\tau_B, \tau_C)$, respectively, where τ_A, τ_B, and τ_C are independent exponential random variables with parameters η, η, and η_C, respectively. The relative magnitude of η_C determines the degree of correlation in the default risk of the two bonds. Thus, by time t, the bonds have paid cumulative dividends of $D_1(t)$ and $D_2(t)$, respectively, where

$$\mathrm{d}\begin{bmatrix} D_1(t) \\ D_2(t) \end{bmatrix} = \begin{bmatrix} 1_{t<T_1} \\ 1_{t<T_2} \end{bmatrix} \mathrm{d}t. \tag{4}$$

[9] For the purposes of calculating the intermediary's reservation value of asset 1, one needs to conjecture the impact of an infinitesimally small trade on the future path of the intermediary's inventory of asset 1. For simplicity, we assume that the intermediary adjusts off-the-equilibrium-path deviations from his target inventory of asset 1 in a similar manner to his trades in asset 2.

[10] If the long-term investors were not pure buy-and hold investors, the most likely candidate to provide them with liquidity at a later date would be the intermediary. Thus, the long-term investor's reservation values for the assets would be positively correlated with those of the intermediary, and the nature of our findings remain unchanged.

Changes in the intermediary's wealth, $W(t)$, are driven by interest on current wealth, proceeds from bond sales, consumption at rate $c(t)$, and dividends on bond inventories, so

$$\mathrm{d}W(t) \quad = \quad (rW(t) - \dot{q}_2(t)V_2(t) - c(t))\,\mathrm{d}t + \sum_{i=1}^{2} q_i(t)\,\mathrm{d}D_i(t). \tag{5}$$

For $t > 0$, the intermediary chooses his consumption process, $\{c(t)\} : t > 0\}$, to solve the infinite-horizon, time-homogenous problem

$$J\big[W, q_1(t), q_2(t)\big] = \sup_{\{c\}} \; E_t \left[\int_t^{\infty} \mathrm{e}^{-r(u-t)} \; \frac{-\mathrm{e}^{-\alpha c(u)}}{\alpha} \; \mathrm{d}u \right], \tag{6}$$

where E_t denotes expectation given the information set $\mathcal{F}_t = \{1_{\{T_1 \geq s\}}, 1_{\{T_2 \geq s\}}, s \leq t\}$. A transversality condition, stated in Appendix A, prevents the intermediary from unlimited borrowing and consuming. For $t < 0$, the intermediary chooses his consumption process, $\{c(t)\} : t < 0\}$, to solve the finite-horizon problem

$$M\big[W, q_1(t), t\big] = \sup_{\{c\}} \; \left\{ E_t \left[\int_t^{0} \mathrm{e}^{-r(u-t)} \; \frac{-\mathrm{e}^{-\alpha c(u)}}{\alpha} \; \mathrm{d}u + \mathrm{e}^{-r(0-t)} J(W(0), q_1(0), S+s) \right] \right\}, \tag{7}$$

where $J(\cdot,\cdot,\cdot)$ is the solution to (6). Consumption is required to be $(\mathcal{F}_t)$-adapted, and integrable.

We focus on the case $t > 0$, leaving the solution of (7) to Appendix A. Suppressing from the notation the dependence of q_1, q_2, V_1, and W on t, the Bellman equation associated with problem (6) is

$$\begin{aligned}
\sup_{x \in (-\infty, +\infty)} \Bigg\{ & \frac{\partial J\big(W, q_1, q_2\big)}{\partial W} \left(rW + \sum_{i=1}^{2} q_i - \dot{q}_2 V_2 - x \right) \; + \; \frac{\partial J\big(W, q_1, q_2\big)}{\partial q_2} \dot{q}_2 \\
& + \; rJ\big(W, q_1, q_2\big) - \frac{\mathrm{e}^{-\alpha x}}{\alpha} \; + \; \eta \left[J\big(W, 0, q_2\big) - J\big(W, q_1, q_2\big) \right] \\
& + \; \eta \left[J\big(W, q_1, 0\big) - J\big(W, q_1, q_2\big) \right] + \eta_C \left[J\big(W, 0, 0\big) - J\big(W, q_1, q_2\big) \right] \Bigg\} \; = \; 0.
\end{aligned} \tag{8}$$

The last three terms in the objective function in (8) exploit the fact that the default of a bond, or a joint default at τ_C, has the same effect as reducing the corresponding bond inventory to zero, since we assume zero recovery at default.[11]

[11]The general case of fractional recovery of face value at default at rate γ is attained if $J\big(W, 0, q_2\big), J\big(W, q_1, 0\big)$, and $J\big(W, 0, 0\big)$ are replaced by $J(W + q_1\gamma, 0, q_2)$, $J(W + q_2\gamma, q_1, 0)$, and $J(W + (q_1 + q_2)\gamma, 0, 0)$, respectively.

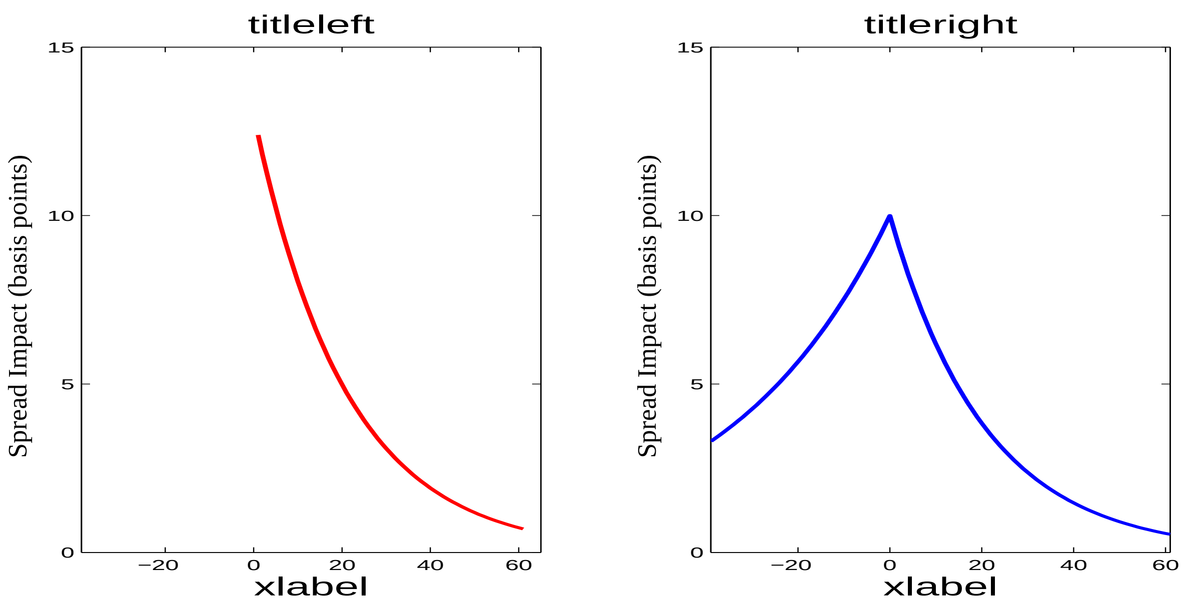
Assuming differentiability of J, which we later verify, the intermediary's reservation values are determined by equating marginal rates of substitution, in that $R_i(q_1,q_2,t)$ for $t>0$ is given by

$$\frac{\partial J\left[W,q_1,q_2\right]}{\partial q_i} = R_i(q_1,q_2,t)\ \frac{\partial J\left[W,q_1,q_2\right]}{\partial W}, \qquad\qquad i \in \{1,2\}. \tag{9}$$

The long-term investor's reservation values, Z_1 and Z_2, for the bonds are derived from a similar problem, detailed in Appendix A. The trade prices of the bonds, V_1 and V_2, are then given by (3). We solve the Bellman equation (8), using (3) and (9), and verify the solution's optimality. The analysis, inspired by DeMarzo and Yan (2003), is relegated to Appendix A.

Figure 2 presents the solution of the impact of bond issuance on spreads, assuming that no default occurs within the given time horizon. An issuance is associated with a temporary widening in the yield-spreads of both the issued bond and the correlated bond.

Figure 2. The Yield-Spread Impact of Debt Issuance



The yield-spread impact of issuance on the issued bond (left) and on a correlated bond (right). The underlying exponential variables τ_A, τ_B, and τ_C have parameters 0.035, 0.035, and 0.005, respectively, implying a default intensity of 400 basis points, a mean time to default of 25 years, and an annual default-event correlation, $\text{corr}\left(1_{T_1<1}, 1_{T_2<1}\right)$, of 0.09, which is typical of the U.S. telecom sector [see DeServigny and Renault (2002)]. Other parameters: $\lambda = 0.2, r = 0.1, \alpha = 0.05, s = 5, S = 400, b_1 = b_2 = 0.5$.

A closed-form solution of the yield-spread impact is not available in this setting, although in Appendix B we provide the intuition with an explicit solution of an analogous problem, differing only in that the cumulative dividend processes D_1 and D_2 are Brownian Motions. In that solution, the magnitude of the other-asset price pressure increases in the riskiness of the issued asset and of the other asset, in the correlation of the dividends of the two assets, and in the issuance size; and decreases in the rate of contact between the intermediary and long-term investors. These predictions motivate an empirical analysis of these comparative statics in Section 3.5.

The remainder of this paper is an empirical study of the market-wide price impact of issuance. We check which of the phenomena that we observe surrounding debt issuance are consistent, at least in character, with our illiquidity-spillover theory.

2. An Empirical Model of the Impact of New Debt Issuance

In this section we describe the data, discuss our measure of "risk-adjusted" issuance, and present two versions of a reduced-form empirical model of changes in bond yield spreads which is based on that measure and aims to estimate the other-bond impact of issuance on yield spreads.

2.1. The Data

Debt issuance in the European telecom industry between October 1, 1999 and July 15, 2001 was dominated by a dozen investment-grade firms: British Telecom, Deutsche Telekom, France Telecom, Portugal Telecom, Sonera (Finland), TDC (Denmark), Telefonica (Spain), TeleNor (Norway), Vodafone (the U.K.), Telecom Italia, KPN (the Netherlands), and Telia (Sweden).

Our database covers 347 bonds issued by these firms and listed in any of Bloomberg, DataStream, or the Reuters Fixed Income Database. Of these, 215 were issued during our sample period in 94 separate issuance-events. For each bond, we record: the principal amount; the denominated currency; the issuing firm; the rate and frequency of coupon payments; the dates of issuance, first coupon, and maturity; and whether the bond has floating coupon rates, options to convert or be called, a "step-up" coupon provision,[12] or a "greenshoe" provision.[13] We include bonds issued by wholly owned subsidiaries of these telecom firms or by a parent company, provided that the credit risk of the bonds is linked to the telecom firm. Table I provides summary statistics.

We calculated the duration of each bond in our sample, daily. In the case of floating-rate notes, we take the duration of the nearly[14] equivalent fixed-rate bond obtained by a fixed-for-

[12]Coupon rates on "step-up" bonds may react to changes in the issuer's credit quality, usually in terms of rating.

[13]A "greenshoe" provision is an option, granted by a securities issuer to its underwriter, to increase the stated size of the issue by as much as 10-15% to meet heavy investor demand or as compensation and incentive to underwrite.

[14]The portfolio of floating-rate note and interest-rate swap is not equivalent at default to the fixed-rate note, since the interest-rate swap need not be at-market at that time. This effect is typically very small.

floating swap of the original note. Ideally, these calculations require the discount factors derived from the issuing firm's yield curve. In the absence of firm-specific zero-coupon yield curves, we use the term structure of swap yields in the bond's denominated currency. (See Appendix C.)

We collected daily trade prices from Reuters and DataStream, and resolved inconsistencies between data sources by consulting *The Financial Times*. These sources capture only publicly traded debt, and prices for privately placed, or otherwise untraded, issues are unavailable. We discarded suspect prices and prices on bonds with fewer than five reported prices. We are left with at least partial time-series of prices for 192 of the bonds in our sample.

We computed the semi-annually-compounded yield-to-maturity for each priced bond, daily. We compute yield spreads of zero-coupon and fixed-rate bonds relative to the yield to maturity of a hypothetical reference bond of the same maturity and coupon structure, based on LIBOR and swap rates of the appropriate currency. The yield spreads on floating-rate bonds are taken to be those of the fixed-rate bonds obtained by a fixed-for-floating swap of the original notes.

We obtained the market value of the firms' equity from DataStream. We computed the aggregate face value of each firm's debt from our bond dataset. (The relatively low frequency of accounting statements renders them unsuitable for our purposes.)

2.2. A Risk-Adjusted Measure of Bond Issuance

In order to estimate the impact of new issuance on market-wide bond yield spreads, we first develop a measure of the quantity of issued debt. The main objective is to capture the risk of changes in the market value of the bond, after hedging interest-rate risk (which investors can hedge away relatively easily), and enable aggregation across a plethora of bonds of different maturities, ratings, domiciles, and structures.

Our issuance measure for a given bond is the product of the bond's market value at issuance with its duration. Specifically, consider bond n, issued at time t_0 with price V_0 and duration S_0. The *risk-adjusted issuance* of this bond at time t is

$$I_{n,t} \equiv \begin{cases} V_0\, S_0, & t = t_0, \\ 0, & \text{otherwise.} \end{cases} \tag{10}$$

This measure approximates the reduction in the issue's market value due to a 100-basis-point parallel increase in the term structure of the issuer's credit spreads. This issuance measure effectively ignores default-event risk. For investment-grade firms, the risk to market value of default is small relative to the risk associated with yield-spread changes. For a bond with a typical annual yield-spread volatility of 100% and an annual default frequency of 0.4%,[15] only 3% of the total variance of change in market value over a 20-day horizon that is due to credit risk (that is, default and yield-spread changes) is due to default risk, assuming 50% recovery at default. Nevertheless, we construct two alternative issuance measures that do account for default-event risk by incorporating the first and second moments of loss given default. Our primary issuance meausre, $I_{n,t}$, also fails to address heterogeneity (across bonds) in the volatility of credit-spread changes, another relatively unhedge-able risk factor. We have constructed an alternative issuance measure that accounts for yield-spread volatility, but which may introduce additional measurement error. The results from estimating our model using these three alternative issuance measures are similar to those obtained using our primary issuance measure, $I_{n,t}$ (see Section 3.4).

Since a plurality of the bonds in our sample are denominated in Euros or Euro-zone currencies (56.1% of principal and 43.9% of the number of bonds), we standardize our measure of risk-adjusted issuance by using the Euro equivalents of the bonds' market values.

We construct a measure of sector-wide risk-adjusted issuance, $\overline{I_t}$, aggregating across time-t telecom-sector issues. For this, let N denote the set of bonds included in our study, and let $I_{n,t}$ be the time-t measure of risk-adjusted issuance due to bond n, given in (10). Then,

$$\overline{I_t} \equiv \sum_{n \in N} I_{n,t} \tag{11}$$

is our definition of aggregate risk-adjusted issuance, which we use, in a regression framework, as an explanatory variable for yield-spread changes.

This measure ignores asymmetries among the firms in the covariances among the bonds' yield-spread changes.[16] For example, a certain note issued by British Telecom and an identical bond issued by France Telecom are treated as though they have the same impact on sector-wide bond yield-spreads. In addition to ignoring heterogeneity in covariance, this issuance measure ig-

[15] Hamilton (2003) reports that the average annual default frequency for the riskiest European investment-grade issues is only 0.4%. The default probability of higher-rated issues is lower.

[16] We experimented with a risk- and correlation-adjusted issuance measure. We picked one bond, with roughly five years of duration at July 2001, from each of the issuers. For each bond pair, we regressed one on the other to get the implied cross-firm yield-spread "sensitivity". In this alternative formulation, changes in the yield-spread of a, say, British Telecom bond are regressed on a risk- and correlation-adjusted issuance measure which aggregates $I_{n,t}$, the time-t risk-adjusted issuance due to bond n, weighted by the sensitivity of British Telecom's yield spreads to the changes in yield spreads of bonds issued by bond n's issuer. Due to the little variation in cross-firm yield-spread sensitivity, we obtain almost identical results when estimating our empirical model with the risk- and correlation-adjusted issuance measure and with the simpler and more tractable risk-adjusted issuance measure.

nores clientele effects induced by market segmentation among investor groups with idiosyncratic preferences by rating class, maturity, domicile, and so on. Such investor heterogeneity could be due to different levels of investor risk aversion, or different contractual prohibitions on allowable investments, as is the case for certain mutual funds or insurance companies.

2.3. A Time-Series Model of Yield-Spread Changes

Let $\Delta s_{f,n,t}$ denote the change from week $t-1$ to week t in the yield spread of bond n, issued by firm f in currency $C(n)$. We estimate the model

$$\Delta s_{f,n,t} = \gamma_0 + \gamma_1 \Delta DIST_{f,t} + \gamma_2 R^E_{f,t} + \gamma_3 \Delta S^{C(n)}_t + \gamma_4 \Delta r^{C(n)}_t + \phi(L) \log\left(\bar{I_t}\right) + \varepsilon_{f,n,t}, \qquad (12)$$

where the distance to default, $DIST_{f,t}$, is the difference between the firm's assets and liabilities, divided by asset volatility (See Appendix E); $R^E_{f,t}$ is the weekly return on the firm's equity; $S^{C(n)}_t$ is the ten-to-two-year slope of the term structure of swap yields for currency $C(n)$ (See Appendix C); and $r^{C(n)}_t$ is the three-month LIBOR rate for currency $C(n)$. The lead-lag polynomial $\phi(L)$ is

$$\phi(L) = \phi_{K_2} L^{K_2} + \cdots + \phi_1 L + \phi_0 + \phi_{-1} L^{-1} + \cdots + \phi_{-K_1} L^{-K_1}, \qquad (13)$$

for some positive integers K_1 and K_2, where L is the standard lag operator. The "error" terms, $\varepsilon_{f,n,t}$, are assumed to be of mean zero and uncorrelated with current and lagged values of the regressors. The use of the logarithm operator in (12) involves an abuse of notation, as we replace $\log\left(\bar{I_t}\right)$ with zero when $\bar{I_t}$ is zero. We refer to the combination of (12) and (13) as the "unconstrained model" of yield-spread changes.

Issue size and issue price may be co-determined, since a firm interested in raising more capital may "price to sell" an issue.[17] As a result, if the yield spreads of newly issued bonds were used in estimating (12), the issuance measure $\log\left(\bar{I_t}\right)$ would be endogenous to the model. To avoid the statistical implications of endogeneity, we exclude from the left-hand side of (12) yield-spread changes of a bond during the first K_2 weeks after its issuance.

The coefficients in $\phi(L)$ correspond to past and future issuance. For example, ϕ_2 and ϕ_{-6} measure the expected change in other-bond yield spreads due to the anticipated value of $\log\left(\bar{I_t}\right)$

[16]For a given firm f, the distance to default, $DIST_{f,t}$, and the issuance measure $\log\left(\bar{I_t}\right)$ may be multi-collinear, since both change when the firm issues debt. We include both in our regression, as the distance to default, unlike $\log\left(\bar{I_t}\right)$, captures all leverage changes in the firm's capital structure, and not just those due to debt issuance.

[17]For example, a bond is "viewed by the market as generously priced - which it needs to be to clear the volume;" in "GM Bonds Tap into Wall of Money'," by A. Roberts and J. Wiggins, *The Financial Times* , June 27, 2003.

two weeks in the future, and the observed value of $\log\left(\overline{I_t}\right)$ six weeks ago, respectively. We implicitly assume that the market is aware of an issuance at least K_1 weeks in advance. Our choice of K_1, as explained in Section 3.1, makes this a relatively benign assumption.

Equation (12) is a model of yield-spread changes. We are ultimately interested in the *cumulative* yield-spread impacts of issuance. A bond's expected cumulative yield-spread reaction to issuance j weeks after an issue of another bond is $\sum_{k=-K_1}^{j} \phi_k$ per unit of log issuance, where $-K_1 \leq j \leq K_2$. Let e_j denote a (K_1+K_2+1)-dimensional vector with ones in the first j elements and zeros elsewhere, and let $\Omega(\phi)$ denote the estimated covariance matrix for the $\phi(L)$ coefficients. The estimated variance of cumulative yield-spread reactions at week j due to an issuance, due to coefficient uncertainty alone, is $e_j' \, \Omega(\phi) \, e_j$.

The specification of the lead-lag polynomial (13) does not fully exploit the insights of our theory. We expect issuance to raise the yield spreads of other bonds as issuance day approaches. This effect should peak on issuance day, and decay after issuance day. If the impact of issuance on yield spreads is along these lines, then constraining our regression model to conform with these time-from-issuance patterns may improve the accuracy of the estimated coefficients.

More specifically, for a scaling parameter β, and coefficients λ_1 and λ_2 of exponential decay for the lead and lag coefficients, respectively, a reasonable parametric model of the impact of a new issue on other-bond yield spreads is

$$\beta\left(\mathrm{e}^{-\lambda_2 K_2} L^{K_2} + \cdots + \mathrm{e}^{-\lambda_2} L^1 + L^0 + \mathrm{e}^{-\lambda_1} L^{-1} + \cdots + \mathrm{e}^{-\lambda_1 K_1} L^{-K_1}\right), \tag{14}$$

per unit of log issuance. The cumulative other-bond reaction to issuance and its variance (due to coefficient uncertainty alone) can be obtained directly from (14).

In order to address yield-spread changes, we difference (14) to obtain a constrained specification for the lead-lag polynomial $\phi(L)$:

$$\phi(L) \;=\; \beta\left(\sum_{i=K_2}^{0}\left(\mathrm{e}^{-\lambda_2 i} - \mathrm{e}^{-\lambda_2(i+1)}\right) L^i + \sum_{i=1}^{K_1}\left(\mathrm{e}^{-\lambda_2 i} - \mathrm{e}^{-\lambda_2(i-1)}\right) L^i\right). \tag{15}$$

The "constrained model" of yield spread changes, given by the combination of (12) and (15), requires the estimation of three parameters, β, λ_1, and λ_2, for the other-bond impact of issuance. In contrast, the number of parameters required in the unconstrained model for the same purpose is equal to the number of lead and lag terms, K_1+K_2+1. Thus, the constrained model, if reasonably specified, is less likely to over-fit the data than is the unconstrained model.

2.4. Control Variables and their Predicted Signs

We discuss below the choice of control variables in (12), and predict the signs of their coefficients.

- **Distance To Default:** Under structural models, such as Merton (1974), Black and Scholes (1973), and Black and Cox (1976), the distance to default is a sufficient statistic for default-event risk. This concept was popularized by KMV [see Vasicek (1984), Kealhofer (2003a) and Kealhofer (2003b)]. We expect a negative coefficient for $DIST_{f,t}$.
- **Equity Return:** Favorable information is expected to increase the firm's equity returns and to decrease the yield spreads on its bonds. We include equity returns in (12) to control for the effect of fundamental information on yield-spread changes, and expect a negative coefficient for $R^E_{f,t}$.[18]
- **Short-Maturity Interest Rates:** Longstaff and Schwartz (1995), Duffee (1998), and Collin-Dufresne, Goldstein, and Martin (2001) find that yield spreads fall when treasury yields rise. We expect a negative coefficient for r^C_t.
- **Slope of the Term Structure of Swap Rates:** An increase in the slope of the risk-free term structure increases the expected future short rate. [See Litterman and Scheinkman (1991).] Therefore, if the short-maturity interest rate has a negative effect on yield spreads, so should the slope of the term structure. Additionally, a decline in the slope of the term structure may imply a weakening economy, when loss-given-default is expected to rise [see Frye (2000)] and yield spreads should widen. We expect a negative coefficient for S^C_t.

We considered several European bond indices for inclusion as controls for general trends in market-wide yield spreads, but opted not to include them in our model. First, they are heavily laden with European telecom-sector debt. Second, while index returns are available, the associated spreads are not, and seem difficult to estimate with available data.

We do not control for tax effects, as do, for example, Elton, Gruber, Agrawal, and Mann (2001). (They explain levels of spreads, and we explain changes in spreads; tax effects on spreads are relatively stable over time.) Our yield spreads are not relative to government bond yields, but rather to swap and LIBOR rates, which, in general, do not enjoy beneficial tax treatment.

We also consider variants of (12) with alternative control variables. See Section 3.4.

[18] Schaefer and Strebulaev (2003) find negative correlation between equity returns and yield-spread changes. Elton, Gruber, Agrawal, and Mann (2001) and Collin-Dufresne, Goldstein, and Martin (2001) find that higher market returns are associated with a tightening in yield spreads. They conjecture that equity and debt prices are affected by similar risk factors. We include the firm's equity returns instead to better capture firm-specific information.

3. The Impact of Debt Issuance on Yield Spreads

This section presents our empirical findings on the impact of debt issuance on yield spreads.

3.1. Estimation Methodology

In addition to excluding the yield spreads of newly issued bonds from the estimation of yield-spread changes model (see Section 2.3), we eliminate the yield-spreads of convertible and callable bonds due to the price distortions caused by their embedded optionality. We also eliminate the yield spreads of bonds with a face value of less than 500 million Euros, with more than five years since issuance, and with less than one year till maturity.[19] We estimate (12) with the remaining 1889 yield-spread observations.

In most cases, the *Financial Times* provides a brief description of a new bond issue on the first business day following issuance. Information about an upcoming issue may, however, be available earlier to market participants. *The Financial Times* reported the intent to issue the bonds in our sample a early as seven weeks before the issuance date.[20] With that in mind, we set the number of weekly lead terms of risk-adjusted issuance (K_1) to seven.

In contrast, no such guidance exists for K_2, the number of lag terms. We estimated the unconstrained version of the model with K_2 ranging from one to twenty-five, and calculated the Akaike information criterion and the Schwartz criterion for each specification. Based on these criteria, we estimate our model with twelve weekly lags of risk-adjusted issuance ($K_2 = 12$).

[19] Similar selection criteria are used by Collin-Dufresne, Goldstein, and Martin (2001) and in the construction of the Lehman Brothers' Liquid Corporate Bonds Index.

[20] Typically, the financial press mentions the firm's intent to issue debt several weeks before the actual issuance date. Tentative details are gradually revealed via conference calls with investors and "roadshows" in major cities. The deal's size and structure is announced on the issuance day, although issues can still be withdrawn, and the issue size may be increased after the issuance if a "greenshoe" option is exercized. For example, consider the largest one-day issuance in our database: 15.5 billion Euros in six tranches issued by Deutsche Telekom on June 28, 2000. On May 2, 2000 the Financial Times reports that DT intends to tap the bond markets to finance part of the winning bid for the UK's third generation mobile phone licenses. Another reference to DT's plans is made on May 28. On June 14, we learn that DT is planning to raise $8 billion in four currencies "towards the end of the month after roadshows in the US and Europe"; the underwriters are also announced. Another mention of this deal is made on June 15, 2000. On June 16: "Roadshows finish in the UK today, and will move to the US next week and Europe the week after." The deal will be in the $8B - $15B range; and tentative maturities are provided, as well as the possibility of a step-up provision. *The Financial Times* makes another short reference to this issue on June 21. On June 22: "...bond investors are now fully focused on the eagerly awaited jumbo financing from DT." A tentative issuance date is provided on June 24; "DT is expected to launch an $8B offering early next week." On June 27: "DT is poised this week to launch [its] bond issue... strong investor demand may lead it to increase its planned offering to $15B;" some details (currencies, number of tranches) are mentioned. On June 28, a 14.5 billion Euros deal is priced and the exact details are given. On June 29 we learn that the deal is oversubscribed and a size increase is likely. The deal is completed on June 30 with a final size of 15.5 billion Euros. All quotes are from the *Financial Times*.

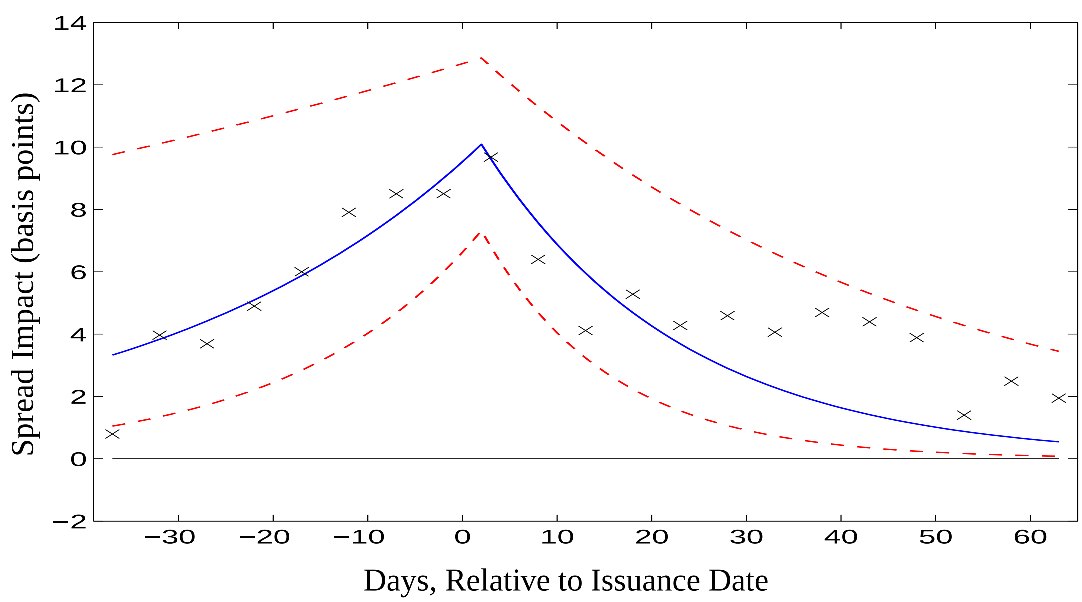
We use ordinary least squares to estimate our unconstrained model of yield-spread changes, and non-linear least squares to estimate its constrained version. Heteroskedasticity could arise from differences in investor clienteles over bonds' country of issue, denominated currency, or rating category, or from risk factors not captured by our control variables. Heteroskedasticity could lead our estimated coefficients to be inefficient and have downward-biased standard errors. We use the Newey and West (1987) autocorrelation- and heteroskedasticity-consistent covariance estimator, with ten lags, to estimate the covariance matrix of regression coefficients.

3.2. Estimation Results

Tables II and III present the estimated coefficients and the associated t statistics for the unconstrained model and the constrained model, respectively. In both models, the coefficients for the control variables are, as expected (see Section 2.4), negative and significant at the 5% level.

Figure 3 presents the estimated impact of issuance on other-bond yield spreads. The solid line plots the estimates of the constrained model. The dashed lines are two-standard-error confidence bands. The estimates obtained from the unconstrained model are marked '$\times$'.

Figure 3. The Cumulative Impact of Debt Issuance on Other-Bond Yield Spreads



The solid line is the estimated cumulative other-bond yield-spread impact of debt issuances (scaled to reflect the impact of Deutsche Telekom's issuance of 15.5 billion Euro of debt in June of 2000), as implied by the constrained model. The dashed lines are two-standard-error confidence bands. The estimates obtained from the unconstrained model are marked '$\times$'.

Figure 3 is scaled to reflect the impact of the largest issuance in our dataset, a Deutsche Telekom six-tranche bond, issued on June 28, 2000, with an aggregate face value of 15.5 billion Euros. The constrained model predicts that the cumulative other-bond yield-spread impact peaks, on the week of issuance, at 10.08 basis points (with a t statistic of 5.98), and has a half life of 16 days. The unconstrained model generates similar results. The cumulative other-bond yield-spread impact peaks, on the week of issuance,[21] at 9.67 basis points (with a t statistic of 6.14) and has a half life of 15 days. At the time of Deutsche Telekom's issue, the market value of outstanding European telecom-sector debt was 100 billion Euros. The mark-to-market impact of Deutsche Telekom's issue on this portfolio is 260 million Euros (0.26%), as estimated by the constrained model, and 273 million Euros (0.28%), as estimated by the unconstrained model. This effect corresponds to a price elasticity[22] of -3100. In comparison, Scholes (1972) finds a price elasticity of -3000 to large-block stock trades.

Our estimation may be affected by selection bias. The yield spreads that we use in our estimation are derived from reported transaction prices, but we are unable to control for the identity of the parties that engaged in these trades. Investors who are aware of the effects of illiquidity have an incentive to trade with the intermediary immediately surrounding the issuance, while subsequent trades are more likely to be those of less-sophisticated investors.

To further emphasize the economic significance of the impact of issuance on yield spreads, we attribute the yield-spread variance among the regressors in our model. We focus on weeks of issuance, when we expect the effect of issuance to be more pronounced. We find that our measure of the other-bond impact of issuance accounts for 70.6% of the yield-spread variation on issuance-weeks observations. Additionally, during our sample period, the mean bid-ask spread in the telecom sector was 2.61 basis points, substantially less than the sector-wide impact of Deutsche Telekom's issue and of other large issues. The standard deviation of the bid-ask spread was 1.82 basis points. We investigate the impact of issuance on bid-ask spreads in Appendix F.

We have seen that the estimated half life of the other-bond effect is 15 business days. In comparison, Harris and Gurel (1986) find that the half life of the stock price impact due to inclusion in the S&P 500 Index is eleven business days. Chen, Noronha, and Singal (2004) find that the price impact associated with deletion from the S&P Index is nearly fully reversed only after 60 trading days. The empirical microstructure literature finds less persistent reactions to inventory shocks

[21]We perform a Monte-Carlo simulation of the cumulative other-bond yield-spread impact of issuance, using the $\phi(L)$ coefficients from the estimation of the unconstrained model and their associated variances. We are unable to reject, at the 5% confidence level, the null hypothesis that the effect peaks on the week of issuance.

[22]We define the price elasticity as the percent change in supply associated with a 1% higher price.

of equity intermediaries. For example, Madhavan and Smidt (1993) estimate that it takes, on average, 7.3 trading days for imbalances in the inventories of NYSE specialists to be reduced by 50 percent. Hansch, Naik, and Viswanathan (1998) find that the half life of inventory imbalances of dealers at the London Stock Exchange is 2.5 days.

We estimate an economically and statistically significant rise in yield spreads in reaction to an issuance. To the extent that equity returns, which are included in (12) as an explanatory variable, provide a reasonable control for fundamental information that may be revealed during the issuance process about the issuing firm, this price pressure does not appear to be related to such information. Two other characteristics of this issuance effect point away from an information-based explanation: (i) The effect is transitory, and (ii) The effect peaks on the day of issuance, not on the day of announcement. This is consistent with our perception that the issuance effects that we document are mainly due to illiquidity, rather than to fundamental information revealed by the new issue. We further examine the informational content of debt issuance in Section 4.1.

3.3. Characterizing the Same-Bond Impact of Issuance

In estimating (12), we excluded the yield-spread changes of newly issued bonds, due to the endogeneity concerns that we raised in Section 2.3. Thus, we have not – so far – measured the same-bond impact of issuance.

Let $\mathcal{Y}_N$ and $\mathcal{Y}_S$ denote the samples of yield-spread changes of newly issued bonds and of seasoned bonds, respectively. One approach to the endogeneity problem is to explain the yield-spread changes in $\mathcal{Y}_N$ in a system of simultaneous equations, relating issuance size to issuance price and vice versa, along the lines of Green (2000), Chapter 16. The small number of observations in $\mathcal{Y}_N$, however, leaves relatively little power for this method. Instead, we characterize the same-bond impact of issuance by examining how well the coefficients from (12), estimated using the yield-spread changes in $\mathcal{Y}_S$ (which are not subject to the same-bond effect), explain the yield-spread changes in $\mathcal{Y}_N$ (which are subject to the same-bond effect). The null hypothesis is that the same-bond effect is zero, and that the two samples belong to the same population.

Let $\Delta\hat{s}_{f,n}(t)$ be the out-of-sample, null-model-predicted change from week $t-1$ to week t in the yield spread of bond n issued by firm f, defined as

$$\Delta\hat{s}_{f,n}(t) = \hat{\gamma}^f + \hat{\gamma}_1 \Delta DIST_{f,t} + \hat{\gamma}_2 R^E_{f,t} + \hat{\gamma}_3 \Delta S^{C(n)}_t + \hat{\gamma}_4 \Delta r^{C(n)}_t + \hat{\phi}(L)\log\left(\overline{I_t}\right), \qquad (16)$$

where $\hat{\gamma}_0$, $\hat{\gamma}_1$, $\hat{\gamma}_2$, $\hat{\gamma}_3$, $\hat{\gamma}_4$, and $\hat{\phi}(L)$ are the estimated coefficients from (12).

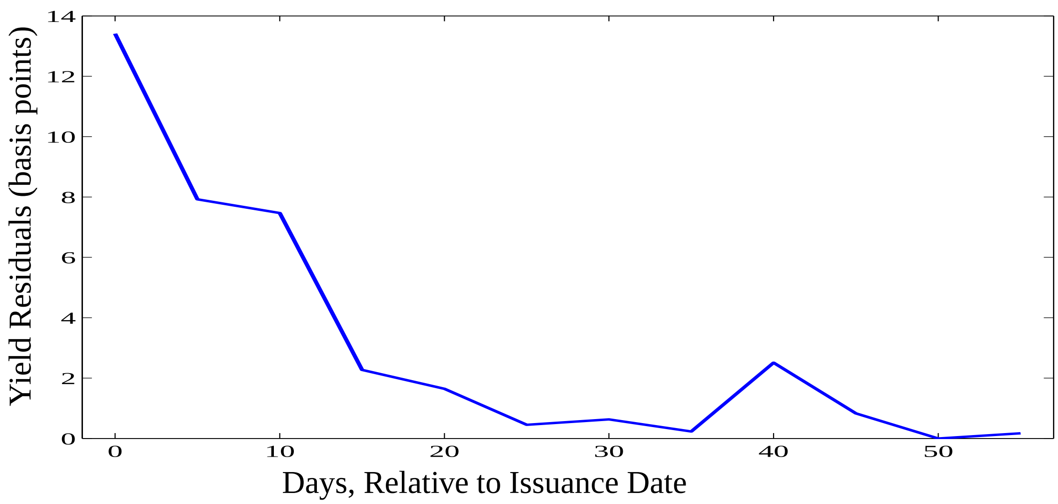
For the 286 observations in $\mathcal{Y}_N$, we subtract $\Delta\hat{s}_{f,t}$ from the observed yield-spread change, $\Delta s_{f,t}$, to obtain the null-model-predicted residuals,

$$\hat{\varepsilon}_n(t) = \Delta s_{f,n}(t) - \Delta\hat{s}_{f,n}(t). \tag{17}$$

Under the null, these residuals are standard least-squares "errors," and should be of mean zero. On the other hand, if the impact of same-bond issuance is as prescribed by our theory, then these residuals would have negative mean. Formally, we test the null hypothesis of no same-bond effect, $H_0 : E[\hat{\varepsilon}_{n,t}] \geq 0$, against the alternative hypothesis, $H_1 : E[\hat{\varepsilon}_{n,t}] < 0$. We reject the null hypothesis (t-statistic of 3.11), and interpret this rejection as consistent with a same-bond issuance effect, which widens bond yield spreads following their issuance.

The null-model-predicted yield-spread-changes residuals allow us to reconstruct the portion of yield-spread *levels* that is left unexplained by (12). For each bond, we calculate the difference between the unexplained portion of yield spreads on the week of issuance and the unexplained portion of yield spreads t weeks after issuance. We average these differences by the time since issuance to obtain the term structure of average null-model-predicted residuals, which we use as a gauge of the same-bond impact of issuance.[23] Figure 4 plots this term structure, which is decreasing almost monotonically, as does the same-bond impact of issuance that our theory predicts. This average impact peaks on the first week after issuance, at 13.54 basis points.

Figure 4. The Term Structure of Average Null-Model-Predicted Yield Residuals



The term structure of average null-model-predicted yield residuals, as implied by the comparison of the predictions of our model to the actual yield spreads of newly issued bonds.

[23]The differences in explained yield-spread levels between the bonds in $\mathcal{Y}_N$ and $\mathcal{Y}_S$ could be also due to different sensitivity to the regressors in (12).

3.4. Robustness of Model Specification

For robustness, we estimated our model on some data subsamples. We split our sample into halves, and estimated the yield-spread reaction to issuance within each sample period separately. The other-bond impact of issuance is more pronounced in the second half of the sample than in the first half (12.9 basis points vs. 8.8 basis points, scaled to reflect the impact of Deutsche Telekom's issuance), as is the impact attributed to the same-bond effect (18.7 basis points vs. 13.0 basis points). We explain this difference with the lower credit quality that prevailed in the second half of our sample.[24] We split the sample of yield spreads into those for bonds with less than five years and more than five years to maturity, and into those for bonds with less than one year and more than one year since issuance.

We also tried alternative model specifications. We replaced our default-event-risk proxy, the firm's distance-to-default, by the ratio of face value of the firm's debt to the market value of its equity. Despite reservations about the appropriateness of credit rating as credit-risk proxies in high-frequency studies,[25] we also included in our regression indicators for the firm's credit ratings. Due to concerns about the possible co-linearity of distance to default and risk-adjusted issuance, we also estimated a variant of (12) without an explicit proxy for the firm's riskiness.

We experimented with alternative issuance measures: (i) the product of risk-adjusted issuance and the sample yield-spreads volatility in the preceding 50 days; (ii) the expected loss due to default on the newly issued bond over the bond's life, assuming zero recovery; and (iii) the standard deviation of expected loss due to default over the same horizon. Appendix D describes the ratings-based calculation of physical-measure default probabilities, required for (ii) and (iii).

These alternative measures address incremental risk borne by investors due to new issues. These risks are different from the risk that we capture with our primary measure of risk-adjusted issuance although they are multi-collinear by construction. Consequently, these measures may also complement our primary measure of risk, $\log\left(\bar{I_t}\right)$, rather than replace it. We therefore add the successively orthogonalized parts of measures (i), (ii), and (iii) to our regression model.

In all cases, the other-bond yield-spread impact of issuance is positive and statistically significant, and we reject the null hypothesis of no same-bond impact. Table IV summarizes the results for the subsamples and model specifications mentioned in this section.

[24] Our theory suggests that the other-bond impact of issuance increases in the credit quality of outstanding bonds. The median credit rating of outstanding bonds is Aa3 in the first half and A2 in the second half of the sample.

[25] Hand, Holthausen, and Leftwich (1992) and others, including Hull, Predescu-Vasvari, and White (2003), have shown that credit-rating changes are lagged responses to perceived shifts in credit quality.

3.5. Comparative Statics for the Other-Bond Impact of Issuance

Our theory predicts how some factors affect the magnitude of the same-bond and other-bond impacts of issuance. We expect a more pronounced other-bond impact of issuance when the issuance is riskier; when the outstanding bond is riskier; when the yield-spread correlation between issuance and outstanding bond is higher; when the contact rate between the intermediary and the longer-term investors is lower; and when the issuance is larger. We examine these predictions below and outline the results in Table V. We do not examine the predictions regarding the same-bond effect, which we have not directly estimated. (Nevertheless, we reject the null hypothesis of no same-bond effect at the 5% confidence level in all of the specifications described below.)

We use the issuing firm's credit ratings as a proxy for a bond's riskiness. We arbitrarily classify firms that were rated A3[26] or above as "high-rated" and all other firms as "low-rated." We assume that yield-spread correlation is higher for bonds issued by the same firm than for bonds issued by different firms. We conjecture that the unobservable contact rate between intermediaries and long-term investors may be related to a bond's currency . We use the median size of the bonds in our sample, 250 million Euros, to differentiate between "large" and "small" bonds.

Issuance Credit Quality

We estimated a variant of (12) in which the risk-adjusted issuance, $\log\left(\overline{I_t}\right)$, is replaced by two separate measures: the risk-adjusted issuance due to high-rated firms and the risk-adjusted issuance due to low-rated firms. As expected, the estimated maximal other-bond impact of low-rated issues is higher than that of high-rated issues (15.97 basis points vs. 6.80 basis points, scaled to reflect the impact of Deutsche Telekom's June 2000 issue). Both impacts are significant at the 1% confidence level. Using a standard Wald test, we reject, at the 5% confidence level, the null hypothesis of identical impacts. Panel A of Figure 5 compares the other-bond impact of high-rated issues with that of low-rated issues.

Outstanding Bond Credit Quality

Our sample consists of 1393 yield-spreads observations of high-rated bonds and 496 yield-spreads observations of low-rated bonds. As expected, the estimated impact on high-rated bonds is less pronounced than the estimated impact on low-rated bonds (9.23 basis points vs. 13.22 basis points, scaled to reflect the impact of Deutsche Telekom's June 2000 issue). Both impacts are significant at the 1% confidence level. Using a standard Wald test, we reject, at the 5% confidence

[26]Changing this threshold does not affect the nature of our results.

level, the null hypothesis of identical impacts. Panel B of Figure 5 plots the other-bond impact of issuance on high-rated bonds and on low-rated bonds.

Yield-Spread Correlation

We split our data into 207 yield-spread observations of bonds issued by firms that issued another bond on the same week and 1418 yield-spread observations of all other bonds. As expected, the estimated other-bond impact of issuance is more pronounced for bonds issued by the same firm than for bonds issued by other firms (9.47 basis points vs. 8.70 basis points, scaled to reflect the impact of Deutsche Telekom's June 2000 issue). Both impacts are significant at the 1% confidence level. However, using a standard Wald test, we can not reject the null hypothesis of identical impacts at any reasonable confidence level. Figure 6 compares the other-bond impact of issuance on bonds issued by the same firm with the impact on bonds issued by other firms.

Issuance Currency

We estimated a variant of (12) in which $\log\left(\overline{I_t}\right)$ is replaced by two measures, the risk-adjusted issuance due to Dollar-denominated issues and the risk-adjusted issuance due to Euro- and Pound-denominated issues. The magnitudes of the maximal impacts (scaled to reflect the impact of Deutsche Telekom's June 2000 issue) are similar: 9.66 basis points for Euro- and Pound-denominated issues vs. 9.20 for Dollar-denominated issues. Both impacts are significant at the 1% confidence level. Using a standard Wald test, we can not reject the null hypothesis of identical impacts at any reasonable confidence level. Figure 7 compares the other-bond impact of Euro- and Pound-denominated issues with that of Dollar-denominated issues.

Issuance Size

We estimated a variant of (12) in which $\log\left(\overline{I_t}\right)$ is replaced by two measures, the risk-adjusted issuance due to large issues and the risk-adjusted issuance due to small issues. The other-bond impact of large issues is positive and significant at 7.99 basis points (scaled to reflect the impact of Deutsche Telekom's June 2000 issue). In contrast, the other-bond impact of small issues is insignificant, suggesting that the linear dependence of yield-spread changes on the logarithm of risk-adjusted issuance, implied by (12), is inaccurate, perhaps because intermediaries absorb a smaller fraction of small issues. This result is distinct from that of Scholes (1972), who finds that the magnitude of the price pressure following a large block sale is unrelated to the block's size.

4. Discussion

We hypothesize that it takes time for liquidity providers, who absorb large amounts of a new issue, to reduce their positions. In the meantime, yield spreads are elevated, both for the new bond and for bonds with correlated credit risk, which are also relatively difficult to hedge during this period. A gradual yield-spread increase precedes a scheduled issuance through anticipation.

The empirical evidence presented in Sections 3.2, 3.3, 3.4, and 3.5 seems to support this theory. The debt issuance places a temporary price pressure on the yield spreads of other bonds, as predicted by our theory. The other-bond yield-spread impact of issuance is economically and statistically significant, and reveals itself in many subsamples and model specifications. We reject the null hypothesis of no same-bond effect. We confirm other predictions of our theory. Riskier bonds are more strongly affected by new issues. Issues of more credit-risky debt have a stronger effect, per unit of issuance, on other-bond yield-spreads, as do larger issues.

This sensitivity of yield spreads to sector-wide debt issuance means that investors in bonds of a given industry face the possibility that a bond issuance in that industry may temporarily depress the market value of their holdings. For instance, *The Financial Times* mentions[27] such concerns:

> "With companies such as British Telecom, France Telecom and Telecom Italia seen as potential issuers, the impact of future supply on prices in the secondary market remains the biggest threat to existing bond holders."

It is not clear, however, whether *The Financial Times* refers to a transitory price impact of issuance, as implied by our theory, or to a permanent price impact, for example due to investors' risk aversion (see Section 4.3).

The transitory price impacts of issuance that we document could potentially be exploited by arbitrageurs. For example, our results imply that an arbitrageur who purchased bonds issued by, say, British Telecom on June 28, 2000 (that is, on the day of Deutsche Telekom's large issuance), and sold those bonds as the impact of Deutsche Telekom's issuance decayed would stand to make, according to the results of Section 3.2, an expected profit of roughly ten basis points, after controlling for other explanatory variables. This "arbitrage," however, involves risk. The estimated Sharpe ratio associated with such a strategy is only 0.34, assuming that the arbitrageur chooses the holding period that maximizes the Sharpe ratio of his trading strategy, and ignoring transaction costs.

[27] See "Deutsche Telekom charms U.S. Investors," by J. Chaffin, *The Financial Times* , June 30, 2000.

Our analysis also suggests that firms have incentives to strategically time their fund-raising in anticipation of debt issuance by other firms in the sector. Simultaneous, or nearly simultaneous, debt offerings by a firm's competitors, would, if our theory is correct, lower the price at which it can issue debt, and rational managers may decide to preemptively issue debt – or to postpone planned debt issuances – in anticipation of their rivals' debt issues. This is in line with location-based theories going back to Hotelling (1929), in which firms separate themselves geographically in order to mitigate the competition of their rivals. For example, we estimate that a hypothetical issuer, seeking to issue a 10 billion Euro bond with a duration of 15 years on the same week as Deutsche Telekom's 15.5 billion Euro issue, would have to "pay" investors 150 million Euros[28] (in the form of a higher yield spread), entirely due to the proximity in time to Deutsche Telekom's issue. Indeed, *The Financial Times* reported on March 28, 2001 that Telecom Italia delayed a 10-billion-Euro issuance because it was "squeezed out by other telecom deals."

We leave this "issuance game" as a topic for further research, but not before providing some incidental support for the idea that issuers respond to each other's issuance-timing choices. Our dataset comprises 94 issuance events. Under the null hypothesis that the dates $\{T_i\}_{i=1}^{94}$ of these events are independent, corresponding to each issuers ignoring the timing of other issue events, we construct samples of these dates via Monte Carlo simulation. The probability distribution of any T_i is assumed, under the null, to be that obtained by taking the probability that T_i is in a given calendar month to be the fraction of issues that actually occurred within that month, and taking the day within the month to be drawn uniformly within the month, except for a correction for day-of-the-week seasonality, based on the observed frequency distribution of issues by day of the week. By repeated independent simulations of samples of $\{T_i\}_{i=1}^{94}$, we obtain the probability distribution, under the null hypothesis, of the approximate total cost to all issuers associated with the other-bond impact of issuance,

$$L\left(\{T_i\}_{i=1}^{94}\right) = \sum_{n=1}^{94} \sum_{m=n+1}^{94} \beta \left(\mathrm{e}^{-\lambda_1 \Delta |T_{n,m}|} I_{n,T_n} + \mathrm{e}^{-\lambda_2 \Delta |T_{n,m}|} I_{m,T_m} \right), \tag{18}$$

where $\Delta T_{n,m}$ is the number of weeks between T_n and T_m, and β, λ_1, and λ_2 are the estimated coefficients of the constrained model. We reject, at the 5% confidence level, the null hypothesis of "independent timing," as the observed issuance impact is smaller than the 5% critical value of

[28]This amount represents the estimated 10 basis points of the other-bond impact of issuance, which translate, for this particular hypothetical issue, to 150 million Euros. We do not consider here the same-bond impact of issuance, since this hypothetical issuer would be subject to that cost regardless of the proximity to Deutsche telekom's issuance.

the null distribution of $L\left(\{T_i\}_{i=1}^{94}\right)$. (One could, however, argue that other institutional considerations, such as marketing costs, create an incentive for issues to separate themselves in time.)

Issuers could also break each large issue into several smaller ones, in an attempt to reduce the impact of issuance on their existing bonds. For example, British Telecom issued 11 billion Euros in debt on December 5, 2000 and only six weeks later, on January 18, 2001, issued an additional 10 billion Euros in debt. Debt issuance, however, involves fixed costs, such as the cost of road shows, that could limit the viability of this option.

An interesting corollary to the predictions of our model is the price impact due to the "mechanical" trades of bond index funds surrounding debt issuance. A new debt issue is assigned a positive "market weight" in bond indices, and thus the weights of other bonds are lowered. These shifts in index weights would compel bond index funds, and investors trying to match the performance of these funds (such as a fund manager whose compensation depends on his returns relative to an index) to buy the new issue, thus becoming the long-term investors in our model. If these investors are budget-constrained, they would have to sell other bonds in their portfolio at the same time. This portfolio re-balancing results in temporary supply and demand imbalances, similar in spirit to those that we model as due to illiquidity. This is consistent with past studies that attribute temporary price pressures surrounding index recompositions to supply and demand imbalances. [See, for example, Harris and Gurel (1986) and Chen, Noronha, and Singal (2004).]

In what follows we consider possible alternative explanations for our empirical findings.

4.1. The Informational Content of Issuance Events

Our underlying theoretical model presumes that the event of issuance conveys no new fundamental information about the bond's issuer, or about other issuers in the sector. (Our model does not rule out that the event of announcement of the intent to issue debt may release information to market participants.) We now contrast this assumption with an alternative, that bond issuance may provide information to market participants about the distribution of future cash flows within the sector, thereby rationally affecting yield spreads on bonds throughout the sector.

This argument is analogous to the information hypothesis discussed in the literature concerning demand-curves for equities. Scholes (1972) proposed this hypothesis to explain equity-price reactions to large-block sales. He suggested that an offer to trade a large block of shares may signal news about the stock, entailing a price reaction. Mikkelson and Partch (1985), Shleifer (1986), Jain (1987), Kalay and Shimrat (1987), Bagwell (1992), and Kaul, Mehrotra, and Morck (2000) consider the information hypothesis as explaining price reactions to various supply shocks.

In general, debt issuance may cause investors to lower their conditional expectations of future cash flows, or to raise their conditional expectations of systematic default risk or future asset volatility. This information, which is relevant for other firms in the sector as well as for the issuing firm, should be incorporated in asset prices when the intention to issue debt is announced, not on the date of issuance. Other information could be released closer to the issuance date. Prior to issuance, the issuing firm is subject to a "due diligence" process. On the date of issuance, the successful placement of debt resolves any uncertainty that the firm might be forced to withdraw its issue and explore more costly avenues for raising capital.

Mirroring an argument of Kalay and Shimrat (1987), if debt issuance provides market participants with new fundamental information about the bond's issuer, this information should be rationally incorporated into equity prices. Intuitively, positive information causes equity returns to rise and yield spreads to contract; negative information has the opposite effect. We include equity returns in (12) to control for the possible effects of fundamental information on bond yield spreads, and estimate statistically significant and negative coefficient. Thus, the other-bond impact of issuance on yield spreads appears to be unrelated to new fundamental information. This is consistent with our perception that the issuance effects that we document are mainly due to illiquidity, rather to fundamental information revealed by the new issue. Two other characteristics of this issuance effect point away from an information-based explanation: (i) The effect is transitory, and (ii) The effect peaks on the day of issuance, not on the day of announcement.

A less noisy conduit, relative to equity returns, for detecting any new fundamental information that the issuance process may release is credit default swap[29] (CDS) rates. These rates could provide an excellent control for bond-specific information. We could also estimate a regression model that relates prices of default swaps to issuance, controling for the determinants of CDS rates suggested by Houweling and Vorst (2003). We obtained data on CDS contracts for the six largest telecom firms. Unfortunately, we are unable to use this data due to its limitations: most of the observations are from the last six months of our sample period, and only a few are transactions, not quotes.

[29] A credit default swap transfers third-party credit risk from a lender to an insurer, in exchange for regular periodic payments (essentially an insurance premium). If the third party defaults, the insurer will have to purchase the defaulted asset from the lender, and pay him the remaining interest on the debt, as well as the principal.

4.2. Industry Debt Capacity

The issuance of debt may change expected recovery rates on all bonds in the sector. Shleifer and Vishny (1992) propose a model of industry debt capacity, in which sector-specific risk factors may cause firms in a given industry to experience simultaneous reductions in cash flows. If a firm defaults, the best users of its assets, other firms in the sector, suffer from debt overhang due to their own reduced cash flows. They, and other less efficient users of the assets, offer fire-sale prices for the distressed assets. A reduction in sector-wide conditional expected recoveries lowers debt prices permanently across all firms in the sector.

However, the transitory nature of the yield-spread issuance impact that we found, as well as the fact that is peaks on issuance dates, not on announcement dates, are inconsistent with the notion that the issuance effects that we estimate are due to the channel described by the Shleifer and Vishny (1992) theory. In their model, debt prices are affected once the intent to issue is declared, rather than when the debt is placed. Additionally, yield spreads would revert to pre-issuance levels quickly only if the total level of industry debt also quickly declined after each debt issue, perhaps through debt retirement or equity-for-debt exchanges. In our study, however, the industry-wide level of debt actually increased tremendously throughout our sample period.

4.3. Investor Risk Aversion and General Equilibrium

Consider a representative investor holding the "market portfolio" of all assets in the economy. This investor buys each new issue. He would avoid reducing his holdings of correlated securities only by a corresponding reduction in price. The price effects of debt issuance in this representative-agent model would not be transitory, unless other security issues were simultaneously retired from the market. As discussed previously, however, the level of telecom-sector assets increased steadily throughout the sample period. The transient nature of the other-bond yield-spread impact of issuance fails to lend support to this hypothesis.

In summary, the above alternative explanations for the observed cross-issuer yield-spread impact in bond markets are not supported by the data. The transience of the yield-spread impact of issuance and the peaking of the effect on the issuance date, rather than on the announcement date, all point away from explanations based on information, complete-markets general equilibrium, and the Shleifer and Vishny (1992) theory of industry debt capacity. The theory of price pressures from market segmentation and investor specialization is consistent with the data in our study.

Table I: **Description of our Telecom-Sector Bond Database**

Issuer	Fixed		Floating		Other		Total		Extant Bonds		New Issuance	
Name	#	Prin.	#	Prin.	#	Prin.	#	Prin.	#	Prin.	#	Prin.
British Telecom	28	16.6B	13	6.6B	10	17.3B	51	40.6B	13	6.2B	38	34.4B
Deutsche Telekom	46	33.0	12	4.4	-	-	58	37.5	1	2.0	57	35.5
France Telecom	39	21.9	8	8.8	11	17.6	58	48.4	24	10.5	34	37.9
KPN	17	19.0	4	6.7	7	1.2	28	26.9	8	5.0	20	21.9
Vodafone	21	17.4	4	8.1	2	2.6	27	28.1	12	8.9	15	19.2
Telecom Italia	7	17.8	2	1.5	9	9.9	18	29.3	11	16.6	7	12.7
Telefonica	25	8.8	6	1.8	13	2.3	44	12.8	31	4.3	13	8.5
TeleNor	13	1.6	3	2.1	2	0.1	18	3.8	12	1.4	6	2.4
TDC	12	2.2	2	0.1	1	0.1	15	2.4	2	0.4	13	2.0
Sonera	4	1.5	1	0.5	-	-	5	2.0	1	0.3	4	1.7
Telia	15	1.4	4	0.9	-	-	19	2.3	12	0.8	7	1.5
Portugal Telecom	2	2.0	-	-	4	0.8	6	2.8	5	1.8	1	1.0
Total	229	143.2B	59	41.5B	59	51.8B	347	236.9B	132	58.2B	215	178.7B

Breakdown of bonds in data set by type. Principals are in billions of Euros. The first column lists the twelve firms included in our data set. Columns two through nine show the number and aggregate principal of the bonds in the data set, separated by type and issuer. Bond types are: fixed rate (second and third columns), floating rate (fourth and fifth columns), or callable, convertible, or miscellaneous (sixth and seventh columns). Columns eight and nine list the total number of bonds and aggregate principal for each issuer in our database. Columns ten and eleven describe the 132 bonds extant as of September 30, 1999; columns twelve and thirteen describe the 215 bonds that were issued during our sample period (October 1, 1999 - July 15, 2001).

Table II. Estimated Coefficients for the Unconstrained Model

Panel A: Control Variables

Variable	Estimate	t-Stat
Intercept	-0.0091	$2.64^{\ddagger}$
$\Delta DIST_{f,t}$	-0.0064	$2.25^{\dagger}$
$R^E_{f,t}$	-0.1735	$5.69^{\ddagger}$
$\Delta r_t^{C(n)}$	-0.0519	$3.23^{\ddagger}$
$\Delta S_t^{C(n)}$	-0.0851	$2.56^{\dagger}$

Panel B: Lead-Lag Polynomial Coefficients

Variable	Estimate	t-Stat	Variable	Estimate	t-Stat
ϕ_{-12}	4.9892e-04	$1.84^{\S}$	ϕ_{-2}	−8.9128e-04	$3.60^{\ddagger}$
ϕ_{-11}	4.2564e-04	$1.68^{\S}$	ϕ_{-1}	−1.2825e-03	$5.11^{\ddagger}$
ϕ_{-10}	−9.7085e-04	$3.96^{\ddagger}$	ϕ_{0}	4.5798e-04	$1.98^{\dagger}$
ϕ_{-9}	−1.9909e-04	0.76	ϕ_{1}	1.3787e-06	0.01
ϕ_{-8}	−1.1674e-04	0.43	ϕ_{2}	2.3221e-04	1.06
ϕ_{-7}	2.4656e-04	0.85	ϕ_{3}	7.4617e-04	$3.14^{\ddagger}$
ϕ_{-6}	−2.0659e-04	0.83	ϕ_{4}	4.3058e-04	$1.90^{\S}$
ϕ_{-5}	1.2263e-04	0.52	ϕ_{5}	4.7200e-04	$2.01^{\dagger}$
ϕ_{-4}	−3.9177e-04	1.60	ϕ_{6}	−1.0830e-04	0.43
ϕ_{-3}	4.5501e-04	$1.78^{\S}$	ϕ_{7}	1.5495e-03	$5.15^{\ddagger}$

Estimated coefficients for control variables (Panel A) and lead-lag polynomials (Panel B) from ordinary least squares (OLS) estimation of

$$\Delta s_{f,n}(t) \quad = \quad \gamma_0 + \gamma_1 \Delta DIST_{f,t} + \gamma_2 R^E_{f,t} + \gamma_3 \Delta S_t^{C(n)} + \gamma_4 \Delta r_t^{C(n)} + \phi(L) \log\left(\bar{I}_t\right) + \varepsilon_{f,n,t}.$$

The dependent variable, $s_{f,n}(t)$, is the yield spread for bond n and issued by firm f. The bond is denominated in currency C(n). Control variables include the distance to default of the issuing firm, $DIST_{f,t}$; the equity return of the issuing firm, $R^E_{f,t}$; the ten-to-two-year slope of the $C(n)$-currency risk-free term structure, $S_t^{C(n)}$; and the three-month risk-free rate for currency $C(n)$, $r_t^{C(n)}$. The lead-lag polynomial $\phi(L)$, given in (13), captures twelve lags and seven leads of our measure of other-bond risk-adjusted issuance, $\bar{I}_t$.

Standard errors are computed using the Newey-West (1987) covariance estimator with ten lags. All 1889 observations are measured along non-overlapping weekly intervals. The adjusted R^2 is 33.8%. The signs §, †, and ‡ represent significance at the 10%, 5%, and 1% levels, respectively.

Table III. Estimated Coefficients for the Constrained Model

Panel A: Control Variables

Variable	Estimate	t-Stat
Intercept	-0.0089	$2.73^{\ddagger}$
$\Delta DIST_{f,t}$	-0.0061	$2.34^{\dagger}$
$R^E_{f,t}$	-0.0923	$2.54^{\dagger}$
$\Delta S^{C(n)}_t$	-0.1633	$7.56^{\ddagger}$
$\Delta r^{C(n)}_t$	-0.0488	$2.23^{\dagger}$

Panel B: Other-Bond Impact Coefficients

Variable	Estimate	t-Stat
β	0.0039	$5.98^{\ddagger}$
λ_1	0.0285	$2.66^{\ddagger}$
λ_2	0.9584	$2.86^{\ddagger}$

Estimated coefficients from non-linear least squares estimation of

$$\Delta s_{f,n}(t) = \gamma_0 + \gamma_1 \Delta DIST_{f,t} + \gamma_2 R^E_{f,t} + \gamma_3 \Delta S^{C(n)}_t + \gamma_4 \Delta r^{C(n)}_t + \phi(L) \log\left(\bar{I}_t\right) + \varepsilon_{f,n,t}.$$

The dependent variable, $s_{f,n}(t)$, is the yield spread for bond n, issued by parent company f. Control variables include the leverage ratio of the issuing firm, $LEV_{f,t}$; the European equity market return, R^M_t; the three-month risk-free rate for currency $C(n)$, $r^{C(n)}_t$; and the ten-to-two-year slope of the $C(n)$-currency risk-free term structure. The non-linear lead-lag polynomial $\phi(L)$ is given in (15).

Standard errors are computed using the Newey West (1987) covariance estimator with ten lags. All 1889 observations are measured along non-overlapping weekly intervals. The Adjusted R^2 is 30.8%. The signs §, †, and ‡ represent significance at the 10%, 5%, and 1% levels, respectively.

Table IV. Robustness Tests for the Impact of Issuance on Other-Bond Yield Spreads

	Other-Bond Impact				Same-Bond Impact		
	No. Obs.	Maximum Impact Estimate	Maximum Impact *t*-Stat	Adj. R^2	No. Obs.	*t*-Stat	Impact
Base Case	1889	9.68	6.14‡	33.8 %	286	3.11‡	12.96
Halves of Sample							
1st Half of Sample	751	8.83	4.96‡	53.0%	104	2.28†	13.02
2nd Half of Sample	1138	12.87	4.75‡	55.7%	182	2.56†	18.73
Time to Maturity							
< 5 Years Till Maturity	792	8.95	3.53‡	30.9%	158	2.53†	10.93
≥ 5 Years Till Maturity	1097	10.32	5.20‡	41.3%	128	3.62‡	12.87
Time Since Issuance							
< 1 Year Since Issuance	1358	8.29	4.57‡	32.1%	286	2.91‡	11.77
≥ 1 Year Since Issuance	531	10.84	4.25‡	53.9%	0		
Manner of Controling for Risk							
Debt-To-Equity	1889	9.20	5.73‡	36.3%	286	3.04‡	12.47
Rating Indicators	1889	10.52	6.48‡	34.6%	286	3.32‡	14.12
No Explicit Control	1889	11.23	6.01‡	33.1%	286	2.84‡	11.54
Manner of Controling for Issuance							
Expected Default Loss	1889	11.91	6.50‡	34.6%	286	2.90‡	12.49
Default-Loss Volatility	1889	10.76	6.17‡	34.2%	286	3.06‡	12.45
Yield-Spread Volatility	1889	8.94	5.92‡	33.7%	286	3.14‡	13.15
Credit Risk and Orthogonalized:							
- Expected Default Loss	1889	8.64	5.30‡	49.1%	286	2.05†	11.15
- Default-Loss Volatility	1889	9.72	6.29‡	49.0%	286	2.57†	11.52
- Yield-Spread Volatility	1889	8.92	4.07‡	53.6%	286	3.03‡	14.62

Robustness checks on estimated maximum cumulative other-bond and same-bond yield-spread reactions (in basis points) to debt issuance, scaled to represent the impact of the June 2000 Deutsche Telekom 15.5 billion Euro issue. The first row reports the estimated base-case reactions. The second and third rows display the estimated impacts for the first half (before August 23, 2000) and second half of the sample period. The fourth and fifth rows present the impacts for bonds maturing within five years and for bonds that mature more than five years into the future. The sixth and seventh rows present the impacts for bonds with less than one year since issuance and for bonds with more than one year since issuance. The eighth, ninth, and tenth rows contain results for when we replace, in Eq. (12), distance to default by leverage, by credit rating (as determined by Moody's), and when leverage is omitted altogether. The eleventh, twelfth and thirteenth rows display estimated yield-spread impacts for alternate issuance measures: (i) the expected loss on the issue due to default within the next year; (ii) the volatility of default losses on the issue within the next year; (iii) the product of our primary issuance measure, $\overline{I_t}$, and the sample volatility of yield spreads in the preceding 50 days. The final three rows copntain the results when issuance is measured by a combination of the primary measure, $\overline{I_t}$, and an orthogonalization of one of measures (i), (ii), and (iii) onto $\overline{I_t}$. We report, in each row, the number of observations used in estimating the other-bond impact of issuance, the maximum other-bond impact and the associated t-stat, and the adjusted R^2. We also provide the number of observations used in testing the null hypothesis of no same-bond impact, the t-Stat of that test, and the difference between the unexplained yield spreads on weeks 1 and 12 after issuance. The signs §, †, and ‡ represent significance at the 10%, 5%, and 1% levels, respectively.

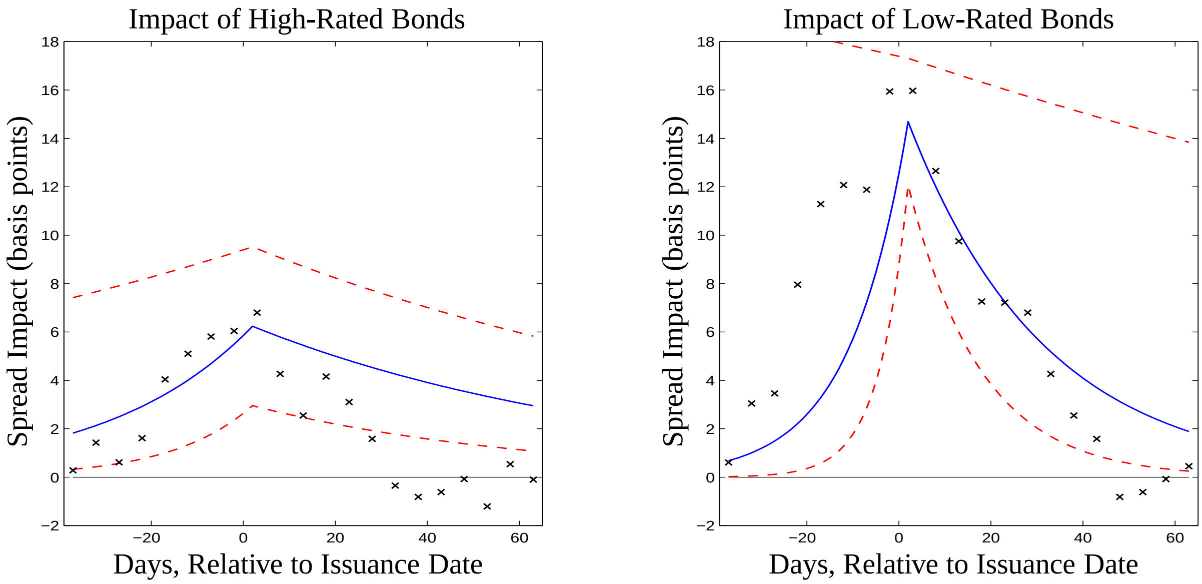
Table V. Comparative Statics for for the Impact of Issuance on Other-Bond Yield Spreads

	Other-Bond Impact				Same-Bond Impact		
	No. Obs.	Maximum Impact Estimate	Maximum Impact t-Stat	Adj. R^2	No. Obs.	t-Stat	Impact
Base Case	1889	9.68	6.14‡	33.8 %	286	3.11‡	12.96
Issuance Credit Quality							
Low-Rated Bonds	1889	6.80	4.44‡	56.3%	286	2.38†	15.20
High-Rated Bonds		15.97	7.91‡				
Outstanding Bond Credit Quality							
Low-Rated Bonds	496	13.22	3.45‡	44.2%	286	2.96‡	12.61
High-Rated Bonds	1393	9.23	5.52‡				
Yield-Spread Correlation							
Same Firm	207	9.47	5.36‡	40.6%	264	2.55	11.32
Other Firms	1418	8.70	1.69§				
Issuance Currency							
Euro-Denominated Bonds	1889	9.66	6.68‡	55.4%	286	2.40†	11.98
Dollar-Denominated Bonds		9.20	5.26‡				
Issuance Size							
New Issuance > 250M Euro	1889	7.99	5.22‡	53.5%	286	2.74‡	13.40
New Issuance ≤ 250M Euro		1.82	0.50				

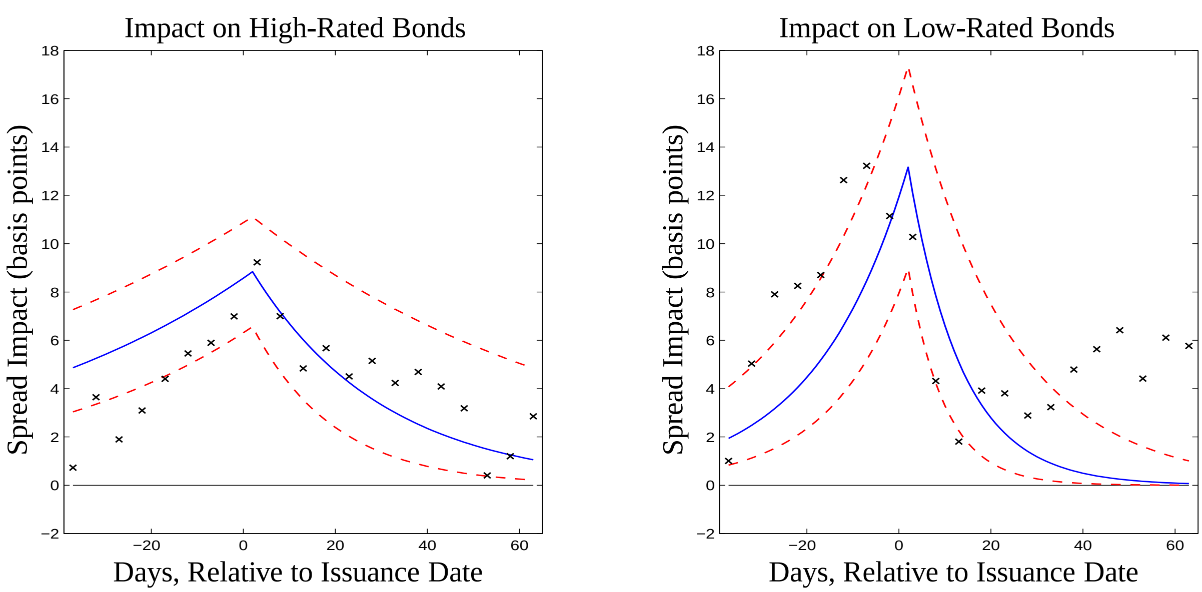
Comparative statics on the estimated maximum cumulative other-bond and same-bond yield-spread reactions (in basis points) to debt issuance. All yield-spread reactions are scaled to represent the impact of the the June 2000 Deutsche Telekom 15.5 billion Euro issue. The first row reports the estimated base-case reactions. The second and third rows present the estimated impacts when issuance of low-rated bonds (rated Ba1 and below) is separated from that of high-rated bonds (rated A3 and above). The fourth and fifth rows display the estimated impacts on low-rated bonds and on high-rated bonds. The sixth and seventh rows present the estimated impact of issuance on bonds issued by the same firm and on bonds issuance by all other firms. The eighth and ninth rows present the estimated impacts when Euro-denominated issuance is separated from Dollar-denominated issuance. The final two rows contain the estimated impacts when new issues are separated into big (face value of more than 250 million Euros) and small (face value below 250 million Euros). We report, in each row, the number of observations used in the estimation of 12, the maximum other-bond impact of issuance and the associated t-Statistic, and the adjusted R^2. We also provide the number of observation used in testing the null hypothesis of no same-bond impact and the t-Stat of that test, and the difference between the unexplained yield spreads on weeks 1 and 12 after issuance. The signs §, †, and ‡ represent significance at the 10%, 5%, and 1% levels, respectively.

Figure 5. The Other-Bond Yield-Spread Impact of Issuance by Credit Rating

Panel A: The Credit Quality of the Issued Bond

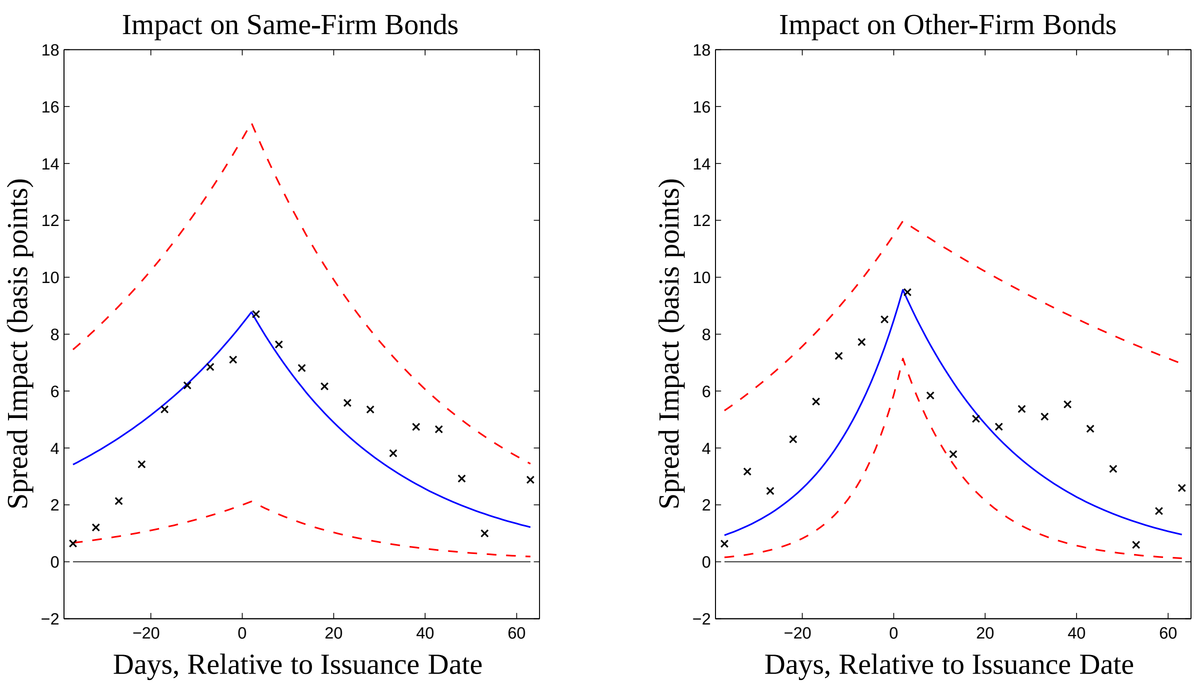


Panel B: The Credit Quality of the Outstanding Bonds



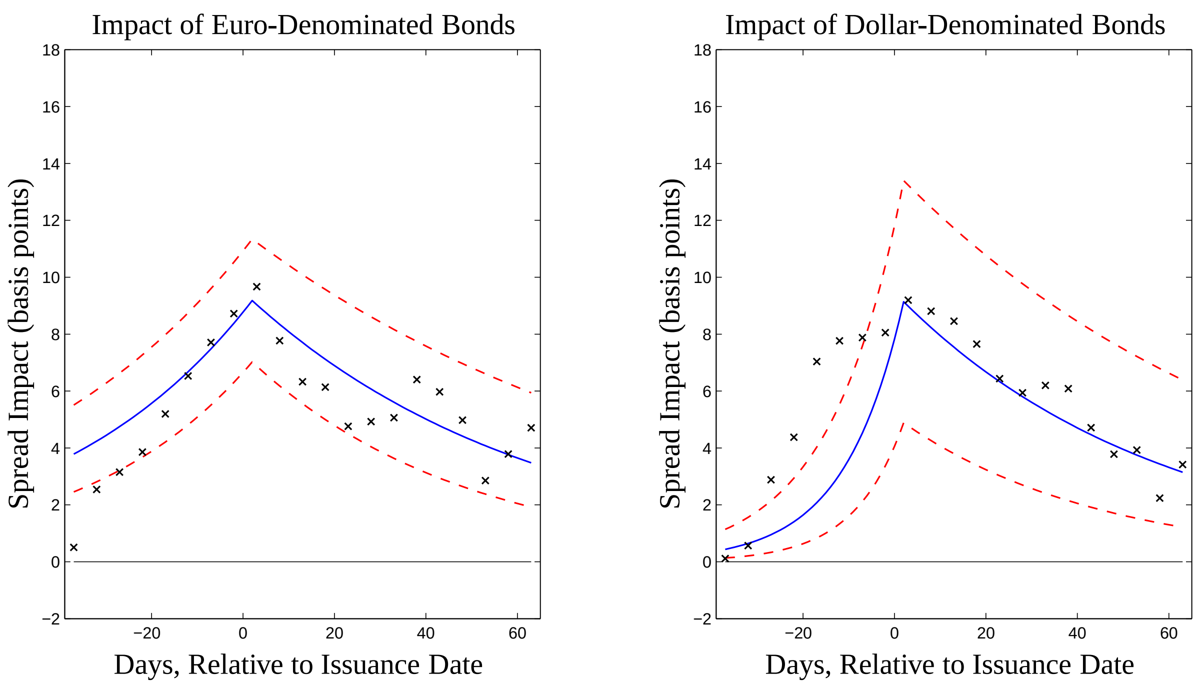
The estimated impact of high-rated issuance (Panel A, left) and low-rated issuance (Panel A, right) on the yield-spreads of other bonds, and the estimated impact of issuance on the yield spreads of high-rated bonds (Panel B, left) and the yield spreads of low rated bonds (Panel B, right). Here, "High rated" and "Low Rated" mean a Moody's credit rating of A3 or above, and Ba1 or below, respectively. For all plots, the solid line is the estimated cumulative other-bond yield-spread impact of debt issuance (scaled to reflect the impact of Deutsche Telekom's issuance of 15.5 billion Euro of debt in June of 2000), as implied by the constrained model. The dashed lines are two-standard-error confidence bands. The estimates obtained from the unconstrained model are marked 'x'.

Figure 6. Other-Bond Yield-Spread Impact of Issuance by Yield-Spread Correlation



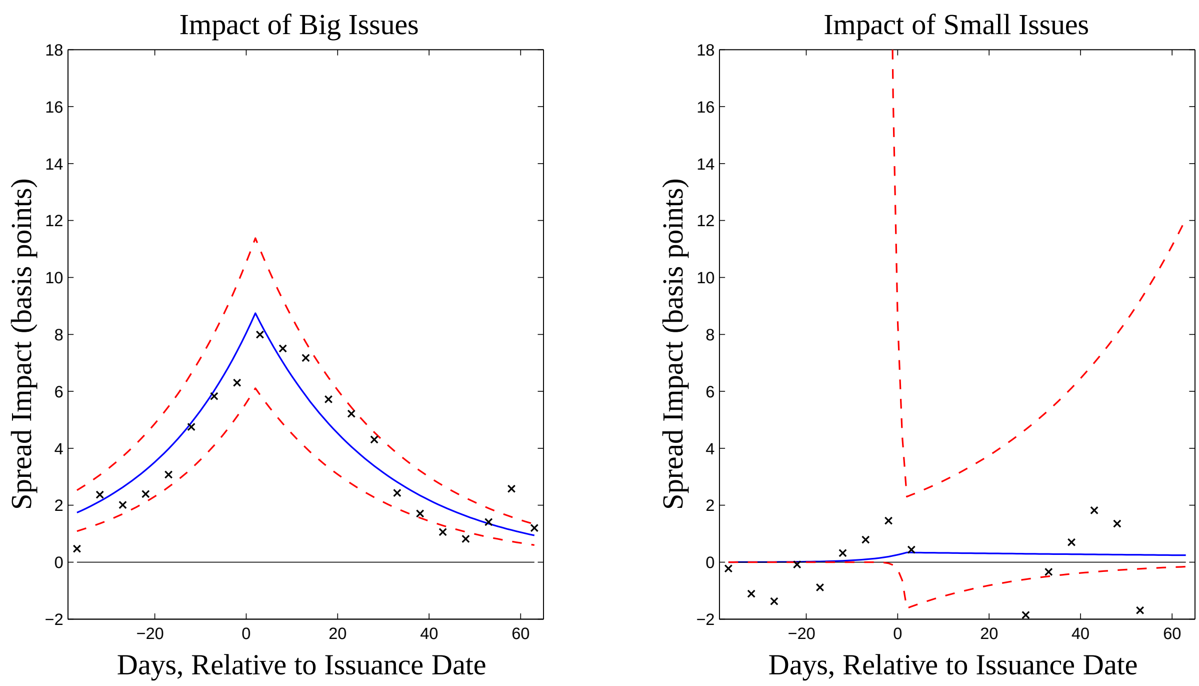
The estimated impact of issuance on the yield-spreads of other bonds issued by other firms (left) and on the yield spreads of other bonds issued by the same firm (right). The solid lines are the estimated cumulative other-bond yield-spread impact of debt issuance (scaled to reflect the impact of Deutsche Telekom's issuance of 15.5 billion Euro of debt in June of 2000), as implied by the constrained model. The dashed lines are two-standard-error confidence bands. The estimates obtained from the unconstrained model are marked 'x'.

Figure 7. Other-Bond Yield-Spread Impact of Issuance by Currency of Issuance



The estimated impact of Euro-denominated issuance (left) and Dollar-denominated issuance (right) on the yield-spreads of other bonds. The solid lines are the estimated cumulative other-bond yield-spread impact of debt issuances (scaled to reflect the impact of Deutsche Telekom's issuance of 15.5 billion Euro of debt in June of 2000), as implied by the constrained model. The dashed lines are two-standard-error confidence bands. The estimates obtained from the unconstrained model are marked 'x'.

Figure 8. Other-Bond Yield-Spread Impact of Issuance by Size of Issuance



The estimated impact of large debt issuance (a face value of more 250 million Euros) and of small debt issuance (a face value of less than 250 million Euros) on the yield-spreads of other bonds. The solid lines are the estimated cumulative other-bond yield-spread impact of debt issuance (scaled to reflect the impact of Deutsche Telekom's issuance of 15.5 billion Euro of debt in June of 2000), as implied by the constrained model. The dashed lines are two-standard-error confidence bands. The estimates obtained from the unconstrained model are marked 'x'.

Appendices

A. Derivation of the Results in Section 1.1

This appendix is organized as follows. In Section A.1, we state the long-term investor's portfolio-consumption problem for $t > 0$ when one of the bonds has defaulted and conjecture a solution to the associated Bellman equation. In Section A.2, we state the long-term investor's problem for $t > 0$ when neither bond has defaulted and conjecture a solution to the associated Bellman equation, using the solution from Section A.1.

In Section A.3, we state the long-term investor's portfolio-consumption problem for $t > 0$ when one of the bonds has defaulted and conjecture a solution to the associated Bellman equation. In Section A.4, we use the solution of Section A.3 to conjecture a solution to the Bellman equation associated with the intermediary's problem for $t > 0$ when neither bond has defaulted, which is stated in Section 1. In Section A.5, we consider the pricing of bond 1 for $t < 0$.

In Section A.6, we verify the optimality of the solutions that we provided in Sections A.1, A.3, A.2, and A.4.

A.1. The Long-Term Investor's Problem when One Bond has Defaulted

The single bond in this economy pays cumulative dividend $D(t)$ by time t, where $\mathrm{d}D(t) = 1_{\{t<T\}}\mathrm{d}t$ and T is an Poisson arrival time with intensity η.

Consider a long-term investor with g bond units, and let $B^*(t)$ denote his time-t wealth. Changes in wealth are driven by interest on current wealth, consumption at rate $a(t)$, and dividends. Thus,

$$\mathrm{d}B^*(t) \quad = \quad (rB^*(t) - a(t))\,\mathrm{d}t + g\,\mathrm{d}D(t). \tag{A.1}$$

This investor chooses his consumption process, $\{a(t) : t \geq 0\}$, to solve

$$H^*\big(B^*(t), g\big) = \sup_{\{a\}} \; E_t\left[\int_t^\infty \mathrm{e}^{-r(u-t)} \; \frac{-\mathrm{e}^{-\alpha a(u)}}{\alpha} \; \mathrm{d}u\right], \tag{A.2}$$

where E_t denotes expectation given the information set $\mathcal{F}_t = \{1_{\{T\geq s\}}, s \leq t\}$. Consumption is required to be $(\mathcal{F}_t)$-adapted, integrable (see Section A.6), and to satisfy a transversality condition stated in Section A.6.

Suppressing from the notation the dependence of B^* on t, the Bellman equation is

$$\sup_{x\in(-\infty,+\infty)}\left\{\frac{\partial H^*(B^*,g)}{\partial B^*}(rB^*+g-x) \;+\; rH^*(B^*,g)\right.$$
$$\left. -\;\frac{e^{-\alpha x}}{\alpha}+\;\eta\left[H^*(B^*,0)-H^*(B^*,g)\right]\;\right\} \;=\; 0, \tag{A.3}$$

nd the first-order condition is

$$\frac{\partial H^*(B^*,g)}{\partial B^*}=e^{-\alpha x}. \tag{A.4}$$

After the bond defaults, the long-term investor faces a simple savings problem.[30] Thus,

$$H^*(B^*,0)=\frac{-e^{-\alpha B^*}}{\alpha r}. \tag{A.5}$$

Consider the candidate solution to the Bellman equation given by

$$H^*(B^*,g) \;=\; -\frac{e^{-\alpha r(B^*+k)}}{\alpha r}, \tag{A.6}$$

where k is a constant to be determined. For this conjectured value function,

$$\frac{\partial H^*(B^*,g)}{\partial B^*}=-\alpha rH^*(B^*,g). \tag{A.7}$$

Combining (A.7) with the first-order condition (A.4), the associated consumption policy is

$$a(t)=r(B^*(t)+k). \tag{A.8}$$

We can identify the constant k from the Bellman equation (A.3). We insert (A.5), (A.6), (A.7), and (A.8) into (A.3), rearrange terms, and obtain

$$\alpha r^2k+\eta e^{\alpha rk}=\alpha rg+\eta. \tag{A.9}$$

The left-hand-side of (A.9) is a bijection and the right-hand-side is a constant, so this equation has a unique solution in k (which we find numerically). The consumption policy given in (A.8) and (A.9) is $(\mathcal{F}_t)$-adapted. Thus, (A.6) is a candidate solution to the long-term investor's problem.

[30]Here, the wealth process is given by $dW(t)=(rW(t)-c(t))dt$. Consider the value function $J(W,t)=\frac{-e^{-\alpha rW}}{\alpha r}$, with $J_W=-\alpha rJ$. The Bellman equation is $\sup_{\{x\}}\left[J_W(rW_t-x)-\frac{e^{-\alpha x}}{\alpha}-rJ\right]=0$, and its first order condition is $-J_W+e^{-\alpha x}=0$. Thus, $c(t)=rW(t)$. The verification argument, under the transversality condition $\lim_{T\to\infty}E_t[e^{-rT}J(W,T)]$, is straightforward.

Subject to the verification argument in Section A.6, the long-term investor's reservation value Z for one unit of bond is given by the unique solution (which does not depend on the wealth B^*) to

$$H^*(B^*, g) = H^*(B^* - Z, 1 + g). \tag{A.10}$$

A.2. The Long-Term Investor's Problem when Neither Bond has Defaulted

After trading with the intermediary, the long-term investor's portfolio consists of one unit each of bond 1 and bond 2. Let $W^*(t)$ be the long-term investor's wealth at time t, given by

$$\mathrm{d}W^*(t) \quad = \quad (rW^*(t) - s(t))\,\mathrm{d}t + \mathrm{d}D_1(t) + \mathrm{d}D_2(t). \tag{A.11}$$

The long-term investor chooses his consumption process, $\{a(t) : t \geq 0\}$, to solve

$$J^*\big(W^*(t), 1, 1\big) = \sup_{\{a\}} \; E_t \left[\int_t^\infty \mathrm{e}^{-r(u-t)} \; \frac{-\mathrm{e}^{-\alpha a(u)}}{\alpha} \; \mathrm{d}u \right], \tag{A.12}$$

where E_t denotes expectation given the information set $\mathcal{F}_t = \{1_{\{T_1 \geq s\}}, 1_{\{T_2 \geq s\}}, s \leq t\}$. Consumption is required to be ($\mathcal{F}_t$)-adapted and to satisfy a transversality condition stated in Section A.6.

Suppressing from the notation the dependence of W^* on t, the Bellman equation is

$$\begin{aligned} \sup_{x \in (-\infty, +\infty)} \Bigg\{ & \frac{\partial J^*(W^*, 1, 1)}{\partial W^*} (rW^* + 2 - x) \; + \; rJ^*(W^*, 1, 1) - \frac{\mathrm{e}^{-\alpha x}}{\alpha} \\ & + \; \eta \left[J^*(W^*, 0, 1) - J^*(W^*, 1, 1) \right] \\ & + \; \eta \left[J^*(W^*, 1, 0) - J^*(W^*, 1, 1) \right] \; + \; \eta_C \left[J^*(W^*, 0, 0) - J^*(W^*, 1, 1) \right] \Bigg\} \; = \; 0, \end{aligned} \tag{A.13}$$

and its first-order condition is

$$-\frac{\partial J^*(W^*, 1, 1)}{\partial W^*} + \mathrm{e}^{-\alpha x} = 0. \tag{A.14}$$

A default of a single bond reduces this problem to that analyzed in Section A.1. Thus,

$$J^*(W^*, 0, 1) = J^*(W^*, 1, 0) = H^*(W^*, 1), \tag{A.15}$$

where $H^*(W^*, 1)$ is given by (A.6) and (A.9), subject to the associated verification argument. As argued before, a joint-default event leads to a simple savings problem, and thus

$$J^*(W^*, 0, 0) = \frac{-\mathrm{e}^{-\alpha W^*}}{r\alpha}. \tag{A.16}$$

Consider the candidate solution to the Bellman equation (A.13) given by

$$J^*(W^*,1,1) \quad = \quad -\frac{e^{-\alpha r(W^*+h)}}{\alpha r}, \tag{A.17}$$

where h is a constant. For this conjectured value function, it is easy to verify that

$$\frac{\partial J^*(W^*,1,1)}{\partial W^*} = -\alpha r J^*(W^*,1,1). \tag{A.18}$$

Combining (A.18) with (A.14), the candidate consumption plan $a(\cdot)$ satisfies

$$a(t) = r(W^* + h). \tag{A.19}$$

We can identify the constant h from the Bellman equation (A.13). We insert (A.15), (A.16), (A.17), (A.18), and (A.19) into (A.13), rearrange terms, and obtain

$$\alpha r^2 h + e^{\alpha r h}\left(2\eta e^{-\alpha r k} + \eta_C\right) = 2(\alpha r + \eta) + \eta_C. \tag{A.20}$$

The left-hand-side of (A.20) is a bijection and the right-hand-side is a constant, so this equation has a unique solution in h (which we find numerically). Thus, (A.8) provides an $(\mathcal{F}_t)$-adapted candidate solution for the long-term investor's problem. Its optimality can be verified by confirming the transversality condition, which we do in A.6.

Subject to that verification argument, the long-term investor's indirect utility from holding one unit each of bond 1 and bond 2 is $J^*(W^*,1,1)$, where J^* is given by (A.17) and h is the unique solution to (A.20). The long-term investor's reservation value Z_2 for one unit of bond 2 is given by the unique solution (which does not depend on W^*) to

$$J^*(W^*,0,1) = J^*(W^* - Z_2, 1, 1). \tag{A.21}$$

Individually, the two bonds are identical. Thus, the investor's reservation value Z_1 for one unit of bond 1 is equal to Z_2, his reservation value for one unit of bond 2.

A.3. The Intermediary's Problem when One Bond has Defaulted

As in Section A.1, the single bond in this economy pays cumulative dividend $D(t)$ by time t, where $dD(t) = 1_{\{t<T\}}dt$ and T is an Poisson arrival time with intensity η. Consider an intermediary with $q(t)$ bond units, who trades away his inventory until it reaches the target s, as described in (1). Let $B(t)$ be the intermediary's wealth at time t, and $R(q(t))$ be his time-t reservation value for one unit of bond.

The trading price $V(t)$ of this bond is

$$V(t) = b\,R\big(q(t)\big) + (1-b)\,Z, \tag{A.22}$$

where b is a constant which represents the relative bargaining power of the intermediary.

Changes in the intermediary's wealth are driven by interest on current wealth, proceeds from bond sales, consumption at rate $c(t)$, and dividends. Thus,

$$\mathrm{d}B(t) \;=\; (rB(t) - \dot{q}(t)V(t) - c(t))\,\mathrm{d}t + q(t)\mathrm{d}D(t). \tag{A.23}$$

The intermediary chooses his consumption process, $\{c(t) : t \geq 0\}$, to solve

$$H\big(B(t), q(t)\big) = \sup_{\{c\}}\; E_t\left[\int_t^{\infty} \mathrm{e}^{-r(u-t)}\;\frac{-\mathrm{e}^{-\alpha c(u)}}{\alpha}\;\mathrm{d}u\right], \tag{A.24}$$

where E_t denotes expectation given the information set $\mathcal{F}_t = \{1_{\{T_1 \geq s\}}, 1_{\{T_2 \geq s\}}, s \leq t\}$. Consumption is required to be $(\mathcal{F}_t)$-adapted, and integrable. Suppressing from the notation the dependence of q, B, and V on t, the Bellman equation associated with this problem is

$$\sup_{x \in (-\infty,+\infty)} \left\{ \frac{\partial H(B,q)}{\partial B}(rB + q - \dot{q}V - x) \;+\; \frac{\partial H(B,q)}{\partial q}\dot{q} \right.$$
$$\left. +\; rH(B,q) \;-\; \frac{\mathrm{e}^{-\alpha x}}{\alpha} + \;\eta\,[H(B,0) - H(B,q)] \right\} \;=\; 0. \tag{A.25}$$

The first-order condition is

$$\frac{\partial H(B,q)}{\partial B} = \mathrm{e}^{-\alpha x}. \tag{A.26}$$

After the bond defaults, the intermediary faces a simple savings problem. Thus, as in (A.5),

$$H(B,0) = \frac{-\mathrm{e}^{-\alpha B}}{\alpha r}. \tag{A.27}$$

Consider the candidate solution to the Bellman equation given by

$$H(B,q) \;=\; -\frac{\mathrm{e}^{-\alpha r(B + k(t(q)))}}{\alpha r}, \tag{A.28}$$

where $t(\cdot)$ is the inverse function of $q(\cdot)$. (This inverse exists and is unique.) Here, $k(\cdot)$ is a differentiable function, to be determined. It is easy to verify that

$$\frac{\partial H(B,q)}{\partial B} - \alpha r H(B,q) \tag{A.29}$$

and

$$\frac{\partial H(B,q)}{\partial q}\dot{q} = -\alpha r H(B,q) k'(t(q)). \tag{A.30}$$

Combining (A.29) with the first-order condition (A.26), the candidate consumption plan is

$$c(t) = r(B(t) + k(t)). \tag{A.31}$$

The intermediary's reservation value for this bond, $R(q(t))$, is determined by equating marginal rates of substitution,

$$\frac{\partial H\big(B(t),q(t)\big)}{\partial q(t)} = R(q(t))\ \frac{\partial H\big(B(t),q(t)\big)}{\partial B}, \tag{A.32}$$

and the trading price V is then given by (A.22).

The function $k(\cdot)$ is identified by the Bellman equation (A.25). We insert (A.27), (A.28), (A.29), (A.30), (A.31), and (A.32) into (A.25), rearrange terms and obtain the ordinary differential equation

$$k'(t(q)) = \frac{-\alpha r^2 k(t(q)) + \eta\left(\mathrm{e}^{\alpha r k(t(q))} - 1\right) + \alpha r q\big(1 + \lambda b Z\big)}{\alpha r \lambda\left(q + (q-s)(1-b)\right)}. \tag{A.33}$$

The right-hand-side of (A.33) is Lipschitz and continuous in k, thus a unique solution exists. We find this solution numerically, using a fourth-order Runge-Kutta method. Thus, (A.28) is a candidate solution for the long-term investor's problem. The candidate consumption plan given in (A.31) and (A.33) is $(\mathcal{F}_t)$-adapted. In Section A.6, we verify the solution's optimality by checking the transversality and integrability conditions.

A.4. The Intermediary's Problem when Neither Bond has Defaulted

This problem is described in Section 1. The intermediary's inventories of the two types of bonds, $q_1(t)$ and $q_2(t)$, are given in (1) and (2). His wealth, $W(t)$, is given by (5). The intermediary chooses his consumption process, $\{c(t) : t \geq 0\}$, to maximize the indirect utility of the bonds in his inventory, $J(W(t), q_1(t), q_2(t))$. Formally, his portfolio consumption problem is (6).

The Bellman equation is (8), with first-order condition

$$\frac{\partial J(W,q_1,q_2)}{\partial W} = \mathrm{e}^{-\alpha x}. \tag{A.34}$$

A default of a single bond reduces the intermediary's problem to that analyzed in Section A.3. Thus,

$$J(W,q_1,0) = H(W,q_1), \tag{A.35}$$

and

$$J(W,0,q_2) = H(W,q_2), \tag{A.36}$$

where $H(W,q)$ is given in (A.28) and (A.33), subject to the associated verification argument.

If both bonds default simultaneously, the intermediary faces a simple savings problem. Thus,

$$J(W,0,0) = \frac{-e^{-\alpha W}}{\alpha r}. \tag{A.37}$$

Consider the candidate solution to the Bellman equation given by

$$J(W,q_1,q_2) \quad = \quad -\frac{e^{-\alpha r(W+G(t(q_2)))}}{\alpha r}, \tag{A.38}$$

where $t(\cdot)$ is the inverse of $q_2(\cdot)$. (The inverse exists and is unique.)

For this conjectured value function, it is easy to verify that

$$\frac{\partial J(W,q_1,q_2)}{\partial W} \quad = \quad -\alpha r J(W,q_1,q_2) \tag{A.39}$$

$$\frac{\partial J(W,q_1,q_2)}{\partial q_1}\dot{q}_1(t) \quad = \quad -\alpha r J(W,q_1,q_2)G'(t). \tag{A.40}$$

$$\frac{\partial J(W,q_1,q_2)}{\partial q_2}\dot{q}_2(t) \quad = \quad -\alpha r J(W,q_1,q_2)G'(t). \tag{A.41}$$

Combining (A.39) with (A.34), the candidate optimal consumption policy $c(t)$ satisfies

$$c(t) = r(W+G(t)). \tag{A.42}$$

The function $G(\cdot)$ is identified from the Bellman equation (8). We insert (3), (9), (A.35), (A.36), (A.37), (A.38), (A.39), (A.40), (A.41), and (A.42) into (8), rearrange terms, and obtain

$$G'(t) = \frac{-rG(t) + \frac{e^{\alpha r G(t)}}{\alpha r}\left(\frac{\eta}{e^{\alpha r k_1(t)}} + \frac{\eta}{e^{\alpha r k_2(t)}} + \eta_C\right) - \Big[q_2(t) + q_1 + \lambda b Z_2(q_2(t)-s)\Big] - \frac{2\eta+\eta_C}{\alpha r}}{\lambda\Big[q_2(t) + (q_2(t)-s)(1-b)\Big]} \tag{A.43}$$

The right-hand-side of (A.43) is Lipschitz and continuous in G, and thus has a unique solution. We find this solution numerically, using a fourth-order Runge-Kutta method. Thus, (A.38) and (A.43) provide a candidate solution for the intermediary's problem, and the associated candidate consumption policy, (A.42), is $(\mathcal{F}_t)$-adapted. In Section A.6, we verify the solution's optimality.

A.5. Pricing Bond 1 for $t < 0$

For $t < 0$, the intermediary solves the finite-horizon portfolio-consumption problem (7). We already know the value function at $t = 0$, $J(B(0), q(0), S+s)$, from solving the intermediary's problem for $t > 0$. Thus, (7) is analogous to the infinite-horizon portfolio-consumption problem

$$H(B(t), q(t), t) = \sup_{\{c\}} \; E_t \left[\int_0^{\infty} \mathrm{e}^{-r(u-t)} \; \frac{-\mathrm{e}^{-\alpha c(u)}}{\alpha} \; \mathrm{d}u \right], \tag{A.44}$$

with the boundary condition $H(B(0), q(0), 0) = J(B(0), q(0), S+s)$. This problem is identical to that analyzed in Section A.3.

A.6. Verification of the Conjectures in A.3, A.4, A.1, and A.2.

The verification argument is identical in all four cases, so we adopt a unifying notation to avoid repetition. Let $W^c(t)$ denote the wealth associated with consumption policy c, from initial wealth w. We take $\Gamma(w)$, the set of admissible consumption policies, to be those satisfying, for $c \in \Gamma(w)$,

$$U(c) \equiv E_0 \left[\int_0^{\infty} \mathrm{e}^{-rs} \frac{-\mathrm{e}^{-\alpha\, c(s)}}{\alpha} ds \right] > -\infty, \tag{A.45}$$

and, from a condition suggested by DeMarzo and Yan (2003),

$$c(t) \leq rW^c(t) + rF, \tag{A.46}$$

for some arbitrary fixed $F > 0$, where r is the risk-free rate. Because a no-savings policy achieves a finite utility, $\Gamma(w)$ is not empty. As will be shown, the DeMarzo and Yan (2003) condition limits the rate of consumption relative to wealth in a manner sufficient to obtain a transversality condition. In what follows, we refer to a generic agent, who is the long-term investor from Sections A.1 and A.2, and the intermediary from Sections A.3 and A.4. Let $X(t)$ be the time-t state variables, other than wealth. The agent chooses his consumption policy, $\{c(t)\}$, from the set of feasible consumptions, $\Gamma(w)$, to maximize $U(c)$, subject to $c \in \Gamma(w)$.

Integrating the Bellman equations (A.3), (A.13), (A.25), and (8) from t to any time $T > t$, and taking expections (assuming that they exist), we get that, for an arbitrary policy $c \in \Gamma(w)$,

$$E_0 \left[\int_0^{T} \mathrm{e}^{-rs} \frac{-\mathrm{e}^{-\alpha c(s)}}{\alpha} \mathrm{d}s \right] \leq J\left(W^c(t), X(t)\right) - E_0 \left[\mathrm{e}^{-rT} J\left(W(T), X(T)\right) \right], \tag{A.47}$$

with an equality for the candidate consumption plan c^* obtained from solving those Bellman equations. Then,

$$\limsup_{T\to\infty} E_0\left[\int_0^T \mathrm{e}^{-rs}\frac{-\mathrm{e}^{-\alpha c(s)}}{\alpha}\mathrm{d}s\right] \le J\left(W^c(t),X(t)\right) - \limsup_{T\to\infty} E_0\left[\mathrm{e}^{-rT}J\left(W^c(T),X(T)\right)\right]. \tag{A.48}$$

The consumption plan c^* obtained from solving the Bellman equation is optimal if, for $X(0) = x$,

$$J(w,x) \ge U(c) = E_0\left[\int_0^\infty \mathrm{e}^{-rs}\frac{-\mathrm{e}^{-\alpha c(s)}}{\alpha}\mathrm{d}s\right], \qquad c\in\Gamma(w), \tag{A.49}$$

with equality c^*. A sufficient condition for (A.49) is the transversality condition

$$\limsup_{T\to\infty} E_0\Big[\mathrm{e}^{-rT}J\left[W(T),X(T)\right]\Big] = 0, \tag{A.50}$$

since in that case (A.48) directly implies (A.49). We are going to show that (A.50) holds, and that c^* satisfies (A.49) with equality.

In each of Sections A.1, A.3, A.2, and A.4, we conjectured that indirect utilities are of the form

$$J\left(W(t),X(t)\right) = \frac{-\mathrm{e}^{-\alpha r(W(t)+N(X(t)))}}{\alpha r}, \tag{A.51}$$

for some $N(\cdot)$. In (A.9) and (A.20), $N(X(t))$ is given by implicit equations. In (A.33) and (A.43), $N(X(t))$ is given by ordinary differential equations. In all cases

$$N(X(t)) < F, \tag{A.52}$$

for some finite constant F. Therefore, substituting (A.52) into the candidate consumption policies (A.8), (A.19), (A.31), and (A.42),

$$c^*(t) \le rW^{c^*}(t) + rF, \tag{A.53}$$

so (A.46) is satisfied. Moreover, a bounded convergence argument implies that these candidate solutions, which satisfy (A.47) in equality, also satisfy (A.49) in equality. A similar argument shows that these candidate solutions also satisfy (A.45).

Using (A.51) and (A.52), the transversality condition (A.50) to be shown is

$$\liminf_{T\to\infty} E_0\left[\mathrm{e}^{-r(T+\alpha W(T))}\right] = 0. \tag{A.54}$$

Using (A.45) and (A.53), we have, for any c in $\Gamma(w)$,

$$-\infty < E_0 \left[\int_0^\infty e^{-rs} \frac{-e^{-\alpha c(s)}}{\alpha} ds \right] \leq \frac{-e^{-\alpha rF}}{\alpha} E_0 \left[\int_0^\infty e^{-r(s+\alpha W(s))} ds \right]. \tag{A.55}$$

Since the integrand in the last term is positive, we use Fubini's Theorem to the change the order of integration, yielding

$$\frac{-e^{-\alpha rF}}{\alpha} E_0 \left[\int_0^\infty e^{-r(s+\alpha W(s))} ds \right] = \frac{-e^{-\alpha rF}}{\alpha} \int_0^\infty E_0 \left[e^{-r(s+\alpha W(s))} \right] ds. \tag{A.56}$$

We conclude that, as $T \to \infty$, $E_0 \left[e^{-r(T+\alpha W(T))} \right]$ cannot be bounded away from zero (otherwise the integral will not be finite), and therefore (A.54) must hold. Thus, the conjectured solutions that we provided in Sections A.1, A.3, A.2, and A.4 are indeed optimal. $\square$

B. An Economy with Two Gaussian Assets

An explicit solution to the price impact of issuance is available if we replace the bonds with assets that have Gaussian payoffs, and if the long-term investors are assumed to have all of the bargaining power when trading with the intermediary (that is, $b = 1$). We offer this solution in order to more easily illustrate the intuition behind the model.

Suppose that the cumulative dividend processes satisfy

$$d \begin{bmatrix} D_1(t) \\ D_2(t) \end{bmatrix} = \begin{bmatrix} \mu \\ \mu \end{bmatrix} dt + \begin{bmatrix} \sigma & 0 \\ \sigma\rho & \sigma\sqrt{1-\rho^2} \end{bmatrix} dB_t, \tag{B.1}$$

where B ia a standard two-dimensional Brownian motion, μ and σ are the mean and volatility of the assets' dividend processes, and the constant $\rho \in [-1,1]$ is a correlation coefficient. The intermediary chooses a consumption process $\{c(t)\}$ that solves (6) (subject to the wealth process in (5) and the usual transversality condition), where E_t denotes expectation given $\{B_s : s \leq t\}$. The Bellman equation, its solution [inspired by DeMarzo and Uroševic (2003)], and verification of optimality are provided below.

Proposition B.1. *The equilibrium prices of assets 1 and 2 are*

$$V_i(t) = \int_t^\infty e^{-r(u-t)} [\mu - p_i(u)] du, \qquad i \in \{1,2\}, \tag{B.2}$$

where

$$p_1(t) = r\alpha\sigma^2 (q_1(t) + q_2(t)\rho), \tag{B.3}$$

$$p_2(t) = r\alpha\sigma^2 (q_2(t) + q_1(t)\rho). \tag{B.4}$$

The equilibrium asset price $V_i(t)$ is the present value of the asset's expected future dividends, at rate μ_i, adjusted by the risk premium $p_i(t)$. The risk premium $p_i(t)$ is proportional to the derivative of the instantaneous variance of the intermediary's dividend process, $\sigma^2(q_1^2(t) + 2\rho q_1(t)q_2(t) + q_2^2(t))$, with respect to the inventory q_i of that asset.[31]

Combining Proposition B.1 with the intermediary's inventory, we obtain the price paths:

$$V_1(t) = \frac{\mu}{r} - \alpha s\sigma^2(1+\rho) - \frac{\sigma^2\rho}{(\lambda+r)} S e^{-\lambda t}, \qquad t \geq 0; \tag{B.5}$$

$$V_2(t) = \begin{cases} \dfrac{\mu}{r} - \alpha s\sigma^2(1+\rho) - \dfrac{\sigma^2}{(\lambda+r)} S e^{-rt}, & t < 0, \\[2ex] \dfrac{\mu}{r} - \alpha s\sigma^2(1+\rho) - \dfrac{\sigma^2}{(\lambda+r)} S e^{-\lambda t}, & t \geq 0. \end{cases} \tag{B.6}$$

The first term, μ/r, in (B.5) and (B.6) is the reservation value of the asset for a risk-neutral investor. The second and third terms are compensation for risk bearing by the intermediary. The second term reflects the intermediary's market-making inventory level s. The third terms in (B.5) and (B.6) are reductions for the temporary excess inventory $q_2(t) - s$ that the intermediary holds as he unloads the new issues, and are proportional to that inventory.

The intermediary's inventory peaks at issuance, as does its effect on prices. As the intermediary trades away his excess inventory, the inventory discount decreases and prices revert toward their original levels. The prices of the issued asset and the correlated asset are related via the common dependence of the intermediary's reservation values for these assets on his inventory position. Prior to issuance date, the prices of both assets reflect anticipated future risk premia. As a result, the issuance places price pressure on *both* assets.

B.1. Proof of Proposition B.1

This proof extends the one-asset result of DeMarzo and Uroševic (2003). We solve for the intermediary's optimal inventory as a function of asset prices, and then impose market clearing.

Let $V_1(t)$ and $V_2(t)$ be the prices of assets 1 and 2, respectively, and let Σ be the covariance matrix in (B.1). Assume that the intermediary's holdings of assets 1 and 2 are $\beta_1(t)$ and $\beta_2(t)$, respectively, and that his consumption policy is c_t. In what follows, we sometimes suppress the

[31]This is comparable to a standard result in a static CARA-Normal framework. Consider an asset whose payoff is normally distributed with mean μ and variance σ^2. A CARA agent with coefficient γ is indifferent between holding q shares of this asset and the certainty equivalent $q\mu - \frac{\gamma}{2}q\sigma^2$. The marginal price this agent would pay is $\mu - \gamma q\sigma^2$.

dependency of β_1, β_2, V_1, and V_2 on time. Then the intermediary's riskless-asset holding, M_t, is given by

$$dM_t = (rM_t + \beta_1\mu + \beta_2\mu - c_t)dt - V_1 d\beta_1 - V_2 d\beta_2 + \begin{bmatrix}\beta_1 & \beta_2\end{bmatrix} \cdot \Sigma \cdot dB_t. \tag{B.7}$$

The intermediary's wealth is $W_t = M_t + \beta_1 V_1 + \beta_2 V_2$. It follows that

$$dW_t = (rM_t + \beta_1\mu + \beta_2\mu - c_t)dt + \beta_1 dV_1 + \beta_2 dV_2 + \begin{bmatrix}\beta_1 & \beta_2\end{bmatrix} \cdot \Sigma \cdot dB_t. \tag{B.8}$$

We define the risk premia on the two assets, $p_1(t)$ and $p_2(t)$, by

$$rV_i + p_i = \frac{dV_i}{dt} + \mu, \qquad i \in \{1,2\}. \tag{B.9}$$

We insert (B.9) into (B.8), and obtain

$$dW_t = (rW_t + \beta_1 p_1 + \beta_2 p_2 - c_t)dt + \begin{bmatrix}\beta_1 & \beta_2\end{bmatrix} \cdot \Sigma \cdot dB_t. \tag{B.10}$$

The Bellman equation for the value function $J(W,t)$ is

$$\sup_{\beta_1,\beta_2,x} \left\{ J_W(rW + \beta_1 p_1 + \beta_2 p_2 - x) + J_t - rJ + \sigma^2 \left(\frac{\beta_1^2 + \beta_2^2}{2} + \beta_1\beta_1\rho \right) J_{WW} + u(x) \right\} = 0, \tag{B.11}$$

with the first order condition (for x)

$$J_W = u'(x). \tag{B.12}$$

Consider the candidate solution to the Bellman equation

$$J(W,t) = \frac{u\big(r(W + k_G(t))\big)}{r} = \frac{-e^{-\alpha r(W + k_G(t))}}{\alpha r}, \tag{B.13}$$

where

$$k_G(t) = \int_t^\infty e^{-r(u-t)} r\alpha\sigma^2 \left(\frac{\beta_1(t)^2 + \beta_2(t)^2}{2} + \beta_1(t)\beta_2(t)\rho \right) du. \tag{B.14}$$

For this conjectured value function,

$$J_W = -\alpha r J, \tag{B.15}$$

$$J_{WW} = (\alpha r)^2 J, \tag{B.16}$$

$$J_t = -\alpha r J k_G'(t). \tag{B.17}$$

We insert (B.15) into (B.12) to obtain the candidate consumption policy

$$c_t = r(W + k_G(t)). \tag{B.18}$$

Maximizing over β_1 and β_2, and using (B.15), (B.16), and (B.17), we obtain

$$p_1(t) = r\alpha\sigma^2(\beta_1(t) + \beta_2(t)\rho), \tag{B.19}$$

$$p_2(t) = r\alpha\sigma^2(\beta_2(t) + \beta_1(t)\rho). \tag{B.20}$$

We check that the value function (B.13) and (B.14) solves the Bellman equation (B.11). We rewrite (B.11) using (B.13), (B.15), (B.16), (B.18), (B.17), (B.19), and (B.20), and obtain the ordinary differential equation

$$k'_G(t) = rk_G(t) - r\alpha\left(\frac{\beta_1(t)^2\sigma^2 + \beta_2(t)^2\sigma^2}{2} + \beta_1(t)\beta_2(t)\sigma^2\rho\right), \tag{B.21}$$

whose solution is given in (B.14). The verification of the solution's optimality, subject to the same constraints that we imposed in A.6 on the feasible consumption set Γ, is identical to that of DeMarzo and Yan (2003). The specialist's inventory of asset 1 is $q_1(t) = s$, and his inventory of asset 2, $q_2(t)$, is given by (2). Market clearing implies that $q_1(t) = \beta_1(t)$ and $q_2(t) = \beta_2(t)$. Thus, we obtain (B.3) and (B.4) from (B.17) and (B.19); (B.2) follows from (B.9). □

C. Risk-Free Zero-Coupon Curves

We use swap yields, rather than treasuries, since government bond prices are often "contaminated" by the presence of tax, repo special, and other regulatory considerations that are irrelevant to our study. We collect daily currency-specific swap yields for one- to thirty-year maturities from DataStream. The t-period swap rate is taken to be the yield to maturity on a semi-annual coupon bond maturing at time t and trading at par. We estimate the implied zero-coupon yield curve by fitting Svensson's (1995) extension of the Nelson and Siegel (1987) model to observed swap rates. Under this model, the zero-coupon yield to maturity t periods in the future is

$$y(t) = a_0 + (a_1 + a_2)\left(\frac{1 - e^{-\lambda_1 t}}{\lambda_1 t}\right) - a_2 e^{-\lambda_1 t} + a_3\left(\frac{1 - e^{-\lambda_2 t}}{\lambda_2 t} - e^{-\lambda_2 t}\right). \tag{C.1}$$

For each day in our sample period, we choose $a_0, a_1, a_2, a_3, \lambda_1$, and λ_2 to minimize the weighted sum of the absolute pricing errors of all available swaps. We replace the estimated yields for maturities of less than one year, which tended to be volatile and even negative, with currency-specific, zero-coupon, money-market LIBOR rates.[32] This tractable parameterization of the zero-coupon yield curve is able to capture a wide variety of term-structure shapes.

D. Default Probabilities

Default probabilities for the firms in our sample are required in the construction of default-risk-adjusted issuance measures. We model each firm's credit rating as a stationary, continuous-time Markov chain over Moody's eight major rating classes[33] Aaa, Aa, A, Baa, Ba, B, Caa, and D, where D is default. Transition from rating class i to rating class j has constant intensity λ_{ij} [see, for example, Resnick (1992)]. For example, the probability that this transition occurs over a short interval of length Δt is approximately $\lambda_{ij}\Delta t$. Default is an absorbing state; the probability of transition from default to any other state is treated as zero.

Consider $P(\Delta)$, the transition matrix between Moody's rating classes over a time horizon of length Δ , where the (i,j) component, $P_{ij}(t)$, is the probability that an i-rated firm will be a j-rated fitm Δ units of time hence. We know that

$$P(\Delta) = e^{A\Delta}, \tag{D.1}$$

where A is the generator of the Markov chain associated with $(\lambda_{i,j})$. Let $\widehat{P}(1)$ be the empirical one-year transition matrix obtained from *www.creditmetrics.com*. (This data set is constructed from 26 years of data on primarily U.S. firms. This sample suffers from a "peso-problem" since some transitions, such as one-year transitions from Aaa to D, were not recorded.) We found that $\widehat{P}(1)$ has a matrix logarithm, denoted $\widehat{A}$. We reset negative off-diagonal entries in $\hat{A}$ to zero, adjust the diagonal entries accordingly, and use the assumed transition matrix, P^* obtained from the adjusted $\hat{A}$ and (D.1). (An alternative approach is given by Lando and Skødeberg (2002).) The resulting time-t estimate of the default probability for firm f within Δ units of time is the entry that corresponds to the firm's time-t credit rating in the D-th column of $P^*(\Delta)$.

The assumption that ratings transition intensities are time-invariant is dubious,[34] as is the assumption that credit ratings are timely measures of credit risk.[35] Nevertheless, this methodology

[32] Daily LIBOR rates for maturities of one to twelve months were obtained from the British Bankers' Association.

[33] Similar techniques are employed by Jarrow and Turnbull (1995), Jarrow, Lando, and Turnbull (1997), Lando (1998), and Elton, Gruber, Agrawal, and Mann (2001). See also Israel, Rosenthal, and Wei (2001) for a more general discussion of using empirical transition matrices to find generators for Markov chains.

[34] Lando and Skødeberg (2002) and Carty and Fons (1994) find evidence of "momentum" effects in ratings transitions. Blume, Lim, and MacKinlay (1998) demonstrate that rating agencies are becoming more conservative in assigning ratings. Nickell, Perraudin, and Varotto (2000) document business cycle effects in ratings transitions.

[35] See footnote 25.

is reasonable for our purposes. For example, (i) Firms that had lower ratings are assigned higher default probabilities, and (ii) The probability of default before maturity increases with time till maturity.

E. Asset Volatility

Asset volatility is required to calculate the distance to default of the firms in our sample. For the purpose of this calculation, we assume that the firm's asset process, $A_{f,t}$, satisfies

$$\log\left(\frac{A_{f,t}-\Omega_{f,t}}{A_{f,t-1}-\Omega_{f,t-t}}\right)=\gamma_0+\gamma_1\log\left(\frac{A_{f,t-1}-\Omega_{f,t-1}}{A_{f,t-2}-\Omega_{f,t-2}}\right)+u_{f,t}. \tag{E.1}$$

where $\Omega_{f,t}$ is the time-t change in the firm f's assets due to debt issuance or retirement.[36] The error terms $u_{f,t}$ are taken to be mean zero with GARCH(2,2) [Bollerslev (1986)] variances, $\sigma_{f,t}$,

$$\sigma^2_{f,t}=\alpha+\gamma_1 u^2_{f,t-1}+\gamma_2 u^2_{f,t-2}+\gamma_3\sigma^2_{f,t-1}+\gamma_4\sigma^2_{f,t-2}, \tag{E.2}$$

for coefficients α, γ_1, γ_2, γ_3, and γ_4 that we estimate. We interpret the variance process $\sigma^2_{f,t}$ as the volatility of the firm's assets.

F. The Impact of Debt Issuance on Bid-Ask Spreads

This appendix examines the impact of issuance on the bid-ask spreads of the bonds in our sample. Amihud and Mendelson (1980) and Ho and Stoll (1981) propose inventory-based models that predict the behavior of bid-ask spreads. We have daily bid and ask prices for these bonds [subject to exclusion criteria that are similar to those that we used in estimating (12)], but not the market depth that these bids and asks represent. Let $BA_{n,f,t}$ be the difference between the yield spreads implied by the bid price and the ask price of bond n, issued by firm f, on week t. We estimate the model:

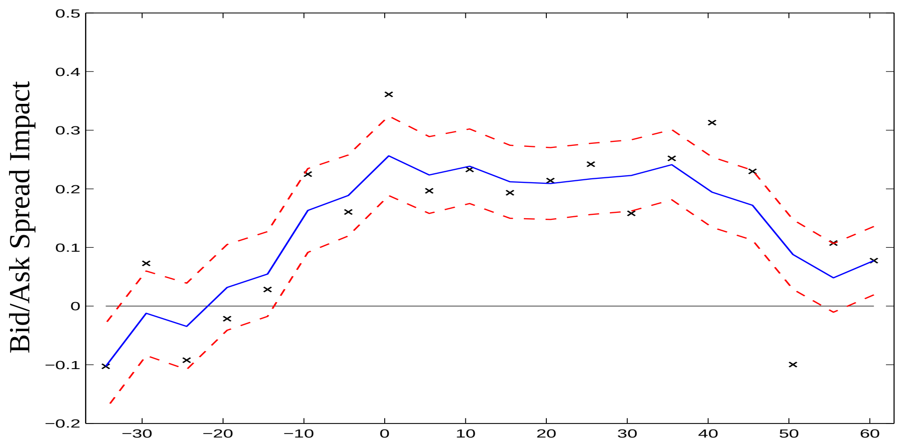
$$BA_{n,f,t}=\alpha_f\;+\;\beta\log(SIZE_n)\;+\;\psi(L)\log\left(\overline{I_t}\right)\;+\;\varepsilon_{n,f,t}, \tag{F.1}$$

where $SIZE_n$ is the face value of bond n in Euros, and $\psi(L)$ is a lead-lag polynomial similar to the one in (13). The "error" terms $\varepsilon_{n,f,t}$ are taken to be of mean zero and uncorrelated with current and lagged regressors. We correct for possible heteroskedasticity and up to ten lags of autocorrelation with a Newey and West (1987) covariance estimator.

[36] Issues and retirements of debt are capital structure changes, not changes in the underlying asset volatility per se.

Our sample contained 2223 weekly observations. Table VI presents the estimated coefficients and the associated t statistics. The regression R^2 is 25.0%. Figure 9 plots the cumulative impact of issuance on bid-ask spreads, as implied by the estimated $\psi(L)$ coefficients.

Figure 9. The Impact of New Debt Issuance on Bid-Ask Spreads



The estimated cumulative impact of issuance on other-bond bid-ask spreads, scaled to reflect the impact of Deutsche Telekom debt issuance in June of 2000. The $\times$'s mark the regression estimates from (F.1). The solid line is an exponential smoothing of those estimates. The dashed lines are an exponential smoothing of two-standard-error confidence bands around these estimates.

We note a mild upward trend in bid-ask spreads near debt issuance. This estimated impact peaks on weeks of issuance at 0.25 basis point, maintains approximately that level for seven more weeks, and is fully reversed by ten weeks after issuance. The mean bid-ask spread in our sample is 2.61 basis points, with a standard deviation of 1.82 basis points. The estimated impact of issuance of bid-ask spreads is almost negligible in comparison to the mean bid-ask spread.

Table VI. Estimated Coefficients for (F.1), The Bid-Ask Spreads Regression

Panel A: Control Variables

Firm	α	t-Stat
British Telecom	0.4893	$8.0828^{\ddagger}$
Deutsche Telekom	0.5033	$7.7732^{\ddagger}$
France Telecom	0.5300	$8.5517^{\ddagger}$
Telecom Italia	0.5050	$7.6968^{\ddagger}$
Royal KPN	0.4766	$7.5050^{\ddagger}$
log(SIZE$_n$)	-0.0243	$7.5346^{\ddagger}$

Panel B: Lead-Lag Polynomial Coefficients

Variable	Estimate	t-Stat	Variable	Estimate	t-Stat
ψ_{-12}	−1.2034e-04	0.71	ψ_{-2}	2.2662e-04	1.58
ψ_{-11}	8.5410e-05	0.54	ψ_{-1}	2.5089e-04	$1.74^{\S}$
ψ_{-10}	−1.0821e-04	0.62	ψ_{0}	2.8376e-04	$2.07^{\dagger}$
ψ_{-9}	−2.5139e-05	0.15	ψ_{1}	1.8546e-04	1.29
ψ_{-8}	3.3131e-05	0.20	ψ_{2}	2.9547e-04	$2.20^{\dagger}$
ψ_{-7}	2.6394e-04	1.56	ψ_{3}	3.6680e-04	$2.70^{\ddagger}$
ψ_{-6}	1.8813e-04	1.27	ψ_{4}	2.6933e-04	$1.91^{\S}$
ψ_{-5}	4.2346e-04	$2.58^{\ddagger}$	ψ_{5}	−1.1682e-04	0.85
ψ_{-4}	2.3060e-04	1.47	ψ_{6}	1.2602e-04	0.94
ψ_{-3}	2.7319e-04	$2.03^{\dagger}$	ψ_{7}	9.0832e-05	0.68

Ordinary-least-squares coefficient estimates from the regression

$$BA_{n,f,t} = \alpha_f + \beta \log(SIZE_n) + \psi(L) \log\left(\overline{I}_t\right) + \varepsilon_{n,f,t},$$

where $BA_{n,f,t}$ is the bid-ask spread (in basis points) on bond n, issued by firm f, at time t; $SIZE_n$ is the face value of bond n in Euros, L is the lag operator, the lead-lag polynomial $\psi(L)$ captures seven leads and twelve lags of our measures of other-bond risk-adjusted issuance, $\overline{I}_t$, and the residuals $\varepsilon_{n,f,t}$ are taken to be of mean zero and uncorrelated with current and lagged regressors.

Our sample contained 2223 non-overlapping weekly observations. Standard errors are computed using the Newey West (1987) covariance estimator with ten lags. The signs §, †, and ‡ represent significance at the 10%, 5%, and 1% levels, respectively. The regression R^2 was 25.0%

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