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A Gap-Filling Theory of Corporate Debt Maturity Choice

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ABSTRACT

We argue that time variation in the maturity of corporate debt arises because firms behave as macro liquidity providers, absorbing the supply shocks associated with changes in the maturity structure of government debt. We document that when the government funds itself with more short-term debt, firms fill the resulting gap by issuing more long-term debt, and vice versa. This type of liquidity provision is undertaken more aggressively: (1) when the ratio of government debt to total debt is higher and (2) by firms with stronger balance sheets. Our theory sheds new light on market timing phenomena in corporate finance more generally.

THERE IS SUBSTANTIAL year-to-year variation in the average maturity of corporate debt issues. For example, using *Flow of Funds* data from the Federal Reserve, which cover all forms of borrowing, including both public and private, we estimate that in 1999, 24.7% of nonfinancial corporate debt issues were "long term"—defined as having a maturity of 1 year or more. This long-term share fell sharply to 19.9% in 2000, and then bounced back to a new peak of 30.1% in 2001.

What accounts for these movements? There are a number of prominent theories of debt maturity choice, but the majority of these theories focus on firm-level determinants and hence do not have clear-cut implications for aggregate time-series behavior. One familiar idea is that firms should attempt to match the maturities of their assets and liabilities (e.g., Myers (1977), Hart and Moore (1995)). Indeed, in Graham and Harvey's (2001) survey of financial managers, this emerges as the most highly cited factor in the debt maturity decision. However, unless there are sharp changes over time in economy-wide asset composition, maturity matching has little to say about the patterns described above. In related work, Diamond (1991) argues that firms decide on debt maturity by trading off the favorable signaling properties of short-term debt against an

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increased risk of inefficient liquidation (see also Flannery (1986)). But again, this model is more naturally suited to making cross-sectional, as opposed to time-series, predictions.

A smaller and almost entirely empirical literature seeks to explain the time series of corporate debt maturity by appealing to "market conditions," which include the general level of interest rates, the slope of the yield curve, etc. (e.g., Taggart (1977), Bosworth (1971), Marsh (1982)).[1] While this may seem like a more natural avenue to pursue, we lack a fully developed theory for *why* such market conditions should matter. One possibility is that managers are eager to increase short-term earnings, perhaps at the expense of long-run value (Stein (1989)). If so, they will tend to borrow at short maturities when the yield curve is steeply upward sloping, and vice versa, simply to keep their current interest expenses low (Faulkender (2005), Chernenko and Faulkender (2008)). This may be why survey respondents tell Graham and Harvey (2001, pp. 224–225) that they prefer to borrow at shorter maturities "when short-term interest rates are low compared to long-term rates." Note that this story can be told in a classical asset pricing setting where the expectations hypothesis of the term structure holds—there is no need to introduce predictability in the relative returns on bonds of different maturities.[2]

An alternative market conditions story, and one that relies on a violation of the expectations hypothesis, is put forward by Baker, Greenwood, and Wurgler (2003), hereafter BGW. They argue that managers time the maturity of their debt issues to exploit the predictability of bond market returns. That is, they issue short-term debt when the expected return on short-term debt is below the expected return on long-term debt, and vice versa.

BGW (2003) offer several pieces of evidence in support of their timing hypothesis. However, they do not explicitly spell out the root sources of bond market predictability, nor do they indicate why corporate issuers might be expected to have a comparative advantage—relative to other market participants—in recognizing or responding to temporary mispricings. Some critics have interpreted BGW as claiming that corporate issuers have a *forecasting advantage* over other players, a premise that these critics see as implausible. As Butler, Grullon, and Weston (2006, p. 1732) put it, "While it is provocative to think that corporate managers may be better able to predict interest rate movements than other market participants . . . most purchasers of corporate debt are sophisticated investors (for example, banks, insurance companies, and pension funds) who are unlikely to make naïve investment decisions."

In this paper, we develop a new theory to explain time variation in corporate maturity choice. As in BGW (2003), our theory allows for predictability in

[1] Several firm-level studies also control for market conditions. See Guedes and Opler (1996), Barclay and Smith (1995), and Stohs and Mauer (1996).

[2] Graham and Harvey (2001) also report that managers borrow short when they are "waiting for long-term interest rates to decline." Thus, if managers believe that the level of rates is slowly mean reverting, we might expect firms to borrow short when the level of interest rates is high. Evidence in Baker, Greenwood, and Wurgler (2003), replicated below, is consistent with this idea.

bond market returns and has the feature that corporate issuers tend to benefit from this predictability—that is, they use short-term debt more heavily when its expected returns are lower than the expected returns on long-term debt. Crucially, however, we do not assume any forecasting advantage for corporate issuers: They have no special ability to predict future returns, or to recognize sentiment shocks. Instead, the key comparative advantage that corporate issuers have relative to other players in our model is an advantage in *macro liquidity provision*.

More specifically, our theory has the following ingredients. First, the bond market is partially segmented, in that there are some important classes of investors who have a preference for investing at given maturities. These investors might include, for instance, pension funds, which, based on the structure of their liabilities, have a natural demand for long-term assets. Second, there are shocks to the supply of long- and short-term bonds that are large relative to the stock of available arbitrage capital. In our empirical work, we associate these supply shocks with changes in the maturity structure of U.S. government debt. And third, there are arbitrageurs (e.g., broker-dealers and, more recently, hedge funds) who attempt to enforce the expectations hypothesis, but—given limited capital and the undiversifiable nature of the required trade—do so incompletely, leaving behind some residual predictability in bond returns.

Taken together, these three ingredients imply that bond market predictability takes a particular form: When the supply of long-term Treasuries goes up relative to the supply of short-term Treasuries, long-term Treasuries must offer a greater expected return. This idea goes back to Modigliani and Sutch (1966a, 1966b) and is developed formally in recent work by Vayanos and Vila (2009), as well as by Greenwood and Vayanos (2008), who provide supporting evidence.[3] Building on these papers, we add one further ingredient to the story, namely, corporate issuers, who have to raise a fixed amount of total debt financing and who must choose whether to issue at short or long maturities. These corporate issuers have no forecasting edge over the arbitrageurs, since government-induced supply shocks are perfectly observable to both types of agents. Rather, what distinguishes the corporate issuers from the arbitrageurs is that they have a potentially greater capacity to absorb the supply shocks. In other words, corporate issuers have a comparative advantage in the provision of this particular kind of liquidity.

The source of this comparative advantage flows from the logic of the Modigliani–Miller (1958) theorem. To see why, imagine a world in which there are no taxes or costs of financial distress, so that firms are indifferent as to the maturity structure of their debt. If we now introduce into this world even tiny differences in the expected returns to short- and long-term debt, firms will respond very elastically by varying the maturity of what they issue. Indeed, in

[3] In related work, Krishnamurthy and Vissing-Jorgensen (2008) show that when the overall supply of Treasury securities goes up, Treasuries offer a greater expected return relative to corporate bonds.

the limit, they will do so until the point where any expected return differentials are eliminated.

In a more realistic setting, firms are likely to have well-defined preferences over their maturity structures, for the reasons alluded to above, and will view it as costly to deviate from their maturity targets. Nevertheless, to the extent that these costs are modest—that is, to the extent that the objective function is flat in the neighborhood of the target—patterns of corporate debt issuance will still respond elastically to differences in expected returns, though no longer to the point of completely eliminating these return differences.

In what follows, we develop this theory with a simple model that embeds the limited-arbitrage logic of Vayanos and Vila (2009) and Greenwood and Vayanos (2008), and that adds a rudimentary corporate sector. We then go on to test four broad implications of the theory:

A. Gap Filling by Corporate Issuers

First and foremost, our theory predicts that corporate issuance will fill in the supply gaps created by changes in government financing patterns. When the government issues more long-term debt, firms should respond by issuing more short-term debt, and vice versa. Consistent with this prediction, we document a strong negative correlation between the maturities of government and corporate debt. A rough estimate is that the corporate sector fills 30% to 40% of the gap created by a shock to government debt maturity. This result holds in a battery of specifications that: (1) use different measures of corporate debt issuance; (2) control for contemporaneous interest rate conditions, credit spreads, and macroeconomic variables; and (3) take into account the dynamics of corporate and government issuance.

One possible objection to our interpretation of these results is that, counter to the spirit of our model, government debt maturity is endogenous and may be influenced by some of the same forces as corporate debt maturity. To address this concern, we instrument for government debt maturity using the ratio of government debt to GDP. These two variables are strongly positively correlated: When the government's financing needs are greater, it tends to extend its offerings out to longer maturities. Moreover, it seems plausible that the ratio of government debt to GDP—essentially, a measure of past fiscal policy—is not itself correlated with the sort of omitted factors that might govern corporate maturity choice, and hence is likely to be a valid instrument. Reassuringly, the results from this instrumental variables approach are nearly identical to our baseline results.

B. Time-Series Variation in Gap Filling

If we allow for time-series variation in the relative sizes of the government and corporate debt markets, our theory makes an additional prediction: When the government's share of total debt is larger, gap-filling behavior by firms will be more pronounced, because larger supply shocks imply a larger reward for liquidity provision. This prediction is also borne out in the data.

C. The Cross-section of Gap Filling

At a micro level, our theory further implies that those firms with the smallest costs of deviating from their maturity targets will be the most aggressive gap fillers. To operationalize this hypothesis, we observe that a firm with a strong balance sheet (a firm that is relatively unconstrained in its investment behavior) is less likely to pay a price if it deviates from its maturity target, thereby taking on, for example, more interest rate or refinancing risk, than a firm with a weak balance sheet. Thus, we would expect firms with stronger balance sheets to have maturity choices that respond more elastically to changes in the structure of government debt.[4] Using a variety of measures of balance sheet strength, we confirm this prediction.

D. The Origin of Corporate Market Timing Ability

As noted above, BGW (2003) document that corporate maturity choices have forecasting power for bond returns, but they do not specify the mechanism that drives this relationship. Our theory suggests that corporate actions can be informative because they are a mirror of government supply shocks, which in turn are the primitive drivers of expected returns. Consistent with this, we find that the ability of corporate issuance to forecast bond returns is attenuated if government debt maturity is included in the regression. Nevertheless, we should stress that our model's implications for returns are neither as fundamental nor as robust as its implications for quantities. In the Modigliani–Miller limit where firms are indifferent as to the maturity of their debt, there will be strong quantitative gap-filling behavior, but all predictability in returns will be arbitraged away. Moving away from the limit, this suggests that any predictability we do find may be modest in nature, even when the mechanism in our model is key to understanding observed corporate debt maturity. Thus, while the predictions for expected returns are of some interest, they are not central for our purposes.

The remainder of the paper is organized as follows. Section I outlines our model of gap filling. Section II discusses the issues that arise in taking the model to the data. Section III describes our measures of corporate and government debt maturity. Sections IV through VII test the four sets of hypotheses described above. Section VIII concludes.

I. The Model

We consider a simple model with three dates labeled 0, 1, and 2. Short-term interest rates follow an exogenous process; one can think of them as determined either by monetary policy or by a stochastic short-term storage technology that is in perfectly elastic supply. In particular, the short-term rate from time 0 to 1, denoted r_1, is known at time 0. The short-term rate from time 1 to 2, denoted

[4] This prediction is similar to that of Hong, Wang, and Yu (2008), who argue that firms with strong balance sheets can act as liquidity providers in their own stocks by repurchasing shares when prices drop below fundamental value.

r_2, is random as of time 0, with mean $E[r_2]$ and variance $Var[r_2]$. There is also a default-free long-term bond that pays one unit of wealth at time 2, and that trades at the price P at time 0; P will be determined endogenously, as described below.

There are four types of actors in our model: preferred-habitat investors, the government, arbitrageurs, and corporations. The preferred-habitat investors can be taken to represent pension funds, life insurance companies, endowments, or others who have a natural demand for long-duration assets. These investors inelastically demand a dollar quantity L of long-term bonds at time 0. At the same time, the government issues a dollar quantity G of long-term bonds. In what follows, we only need to keep track of $g = G - L$, which measures the time 0 excess supply of long-term government bonds relative to preferred-habitat investor demand. The quantity g, which is exogenous in our model, can be either positive or negative.

Next, we add risk-averse arbitrageurs who have zero initial wealth. In equilibrium, they buy a dollar amount h of long bonds at time 0, and finance this by borrowing short term. Note that h can also be negative, in which case the arbitrageurs buy short-term bonds financed with long-term borrowing. Terminal arbitrageur wealth is simply $w = h[P^{-1} - (1+r_1)(1+r_2)]$. We assume that arbitrageurs have mean-variance preferences with risk tolerance γ, choosing h to maximize $E[w] - (2\gamma)^{-1}Var[w]$. Given these assumptions, it is easy to show that arbitrageurs' time 0 demand for long-term bonds is given by

$$h^*(P) = \gamma \frac{[P^{-1} - (1+r_1)(1+E[r_2])]}{(1+r_1)^2 Var[r_2]}. \tag{1}$$

As in Vayanos and Vila (2009), arbitrageurs borrow short and invest long when long-term bonds offer an expected return premium over short-term bonds. Conversely, when the return premium is negative, they borrow long and invest at the short rate.

Suppose for the moment that we leave out corporate issuers. The market clearing condition is $h^*(P) = g$, which implies

$$P^{*-1} - (1+r_1)(1+E[r_2]) = \frac{(1+r_1)^2 Var[r_2]}{\gamma} g. \tag{2}$$

Thus, the expectations hypothesis holds, that is, $P^{*-1} = (1+r_1)(1+E[r_2])$, if any of the following hold: (1) $g = 0$, so that government supply matches preferred-habitat investor demand for long-term bonds; (2) $Var[r_2] = 0$, so that arbitrageurs face no interest rate risk; or (3) γ is infinite, so that arbitrageurs are risk neutral. Otherwise, an increase in the supply of long-term government bonds raises their expected return premium.

As a quantitative matter, equation (2) implies that supply shocks have the potential to generate economically interesting effects to the extent that g is large relative to γ, or in other words, to the extent that the shocks are large compared to the risk tolerance of the arbitrageurs. To get a sense of the magnitudes involved, note that in our sample, the standard deviation of the long-term share of government debt is 9% (around a mean of 59%). The total amount of outstanding government debt at the end of 2005 was \$4.7 trillion.[5] These numbers imply that, in order to absorb a one-standard deviation increase in the maturity of government debt, the arbitrage sector would have to go long \$423 billion of long-term bonds, funding this position at the short-term rate. The annualized standard deviation of excess bond returns is 10%, which implies that this trade has a 1% value-at-risk (VaR) of approximately \$98 billion, assuming normally distributed returns. This \$98 billion VaR figure can be compared to the total assets of macro- and fixed-income arbitrage hedge funds, which were \$118 billion and \$28 billion, respectively, in 2005 according to Hedge Fund Research, Inc. Thus, it seems likely that the limits of arbitrage identified by Shleifer and Vishny (1997) would loom large in this context, especially given that the risk in question is a macro one that cannot be diversified away easily.

The last set of players in our model is a group of operating firms. We assume that these firms collectively need to borrow a total dollar amount C; as will become clear, the parameter C effectively indexes the size of the corporate sector relative to the government sector. Firms raise a fraction f (and hence a dollar amount fC) of their needs from long-term debt, and the remaining $(1-f)$ from short-term debt. Timing considerations aside, their target optimal capital structure involves having a fraction z of long-term debt. If they stray from this target in either direction, they incur quadratic costs (in total dollar terms) of $\theta C(f-z)^2/2$. These costs might reflect interest rate exposure or refinancing risk, either of which could lead to a tightening of financial constraints and ultimately to a reduction in value-creating investment. In this context, the parameter θ can be thought of as a measure of balance sheet strength. In the limit where $\theta = 0$, the firm in question has a balance sheet that is so strong that it is financially unconstrained in all states of the world and it is therefore indifferent as to the maturity structure of its debt. At the other extreme, where θ is large, the firm has tightly binding financial constraints so that any increase in, say, interest rate risk has the potential to be very costly.

In the spirit of Stein (1996), the firm's objective function is to minimize the sum of expected interest costs plus the costs associated with financial constraints. That is, firms solve

$$\min_f \left[C\left((1-f)(1+r_1)(1+E(r_2)) + \frac{f}{P} + \theta\frac{(f-z)^2}{2} \right) \right], \tag{3}$$

which has the solution

$$f^*(P) = z - \frac{P^{-1} - (1+r_1)(1+E[r_2])}{\theta}. \tag{4}$$

[5] This figure refers to publicly held federal debt. The higher figure that one sometimes hears, \$8.2 trillion as of year-end 2005, includes intergovernmental holdings, for example, holdings by the Social Security Administration.

The intuition is that when long-term debt is expensive, that is, when $P^{-1} - (1+r_1)(1+E[r_2])$ is high, firms deviate from their target debt mix and issue less long-term debt ($f < z$).

Once we add the corporate sector to the model, the market clearing condition for long-term bonds becomes $h^*(P) = g + Cf^*(P)$, which implies

$$P^{*-1} - (1+r_1)(1+E[r_2]) = \left[\frac{\theta(1+r_1)^2 Var[r_2]}{\gamma\theta + C(1+r_1)^2 Var[r_2]}\right](g + Cz). \tag{5}$$

We can solve for the equilibrium fraction of long-term corporate debt by substituting equation (5) into equation (4), which yields

$$f^* = z - \left[\frac{(1+r_1)^2 Var[r_2]}{\gamma\theta + C(1+r_1)^2 Var[r_2]}\right](g + Cz). \tag{6}$$

As above, the expectation hypothesis holds as γ tends to infinity or as $Var[r_2]$ goes to zero, since in either case arbitrageurs take arbitrarily large long (short) positions in long-term bonds if they deliver higher (lower) expected returns than short-term bonds. In addition, as θ tends to zero, so that there are no costs of deviating from the target maturity z, firms completely absorb any changes in government supply ($Cf^* = -g$), and the expectations hypothesis holds irrespective of arbitrageur risk tolerance. In such a world, firms respond aggressively to government supply shocks, even though these shocks have no effect on equilibrium prices.

In the limiting case where $\gamma = 0$, so that there are no arbitrageurs, the expected return premium on long-term bonds is given by $(\theta/C)(g + zC)$. This is because there is a net excess supply of long-term bonds of $(g + Cz)$ if firms stick to their target debt mix, while θ/C measures the (lack of) willingness of the corporate sector to absorb this excess supply.

The following four propositions, which follow immediately from equations (4) through (6), provide the basis for our empirical work below.

PROPOSITION 1 (Gap Filling): *It is apparent from equation (6) that* $\partial f^*/\partial g < 0$. *When the government issues more long-term debt, firms respond by tilting their debt issuance away from long-term debt.*

PROPOSITION 2 (Time Variation in Gap Filling): *Equation (6) also implies that* $\partial^2 f^*/\partial g \partial C > 0$. *Gap-filling behavior is more pronounced when the stock of government debt is large relative to the stock of corporate debt.*

One simple intuition for the result in Proposition 2 is that gap filling is fundamentally a dollars-for-dollars phenomenon. When C is small (i.e., there is relatively more government debt), it takes a larger change in the fractional composition f of corporate debt to absorb a given dollar shock to supply. Although the dollars-for-dollars nature of Proposition 2 makes it sound mundane, it is actually a sharply differentiating prediction of our theory. To see why, consider an alternative explanation for gap filling. One might argue, for example, that government debt maturity is itself endogenous, and responds to the same unobserved factors that drive corporate maturity decisions, albeit with the opposite sign. Perhaps the government tends to shorten the duration of its debt when it perceives future economic conditions to be deteriorating, while the corporate sector does just the reverse. This could generate $\partial f^*/\partial g < 0$, as in Proposition 1. But it would *not* generate $\partial^2 f^*/\partial g \partial C > 0$, as in Proposition 2, since in this alternative story, all that is relevant about government financing choices is their informational content, not their raw scale.

PROPOSITION 3 (The Cross-Section of Gap Filling): *Another implication that follows from equation (6) is that* $\partial^2 f^*/\partial g \partial \theta > 0$. *Loosely speaking, firms with stronger balance sheets (those for which* θ *is closer to zero) will exhibit more aggressive gap-filling behavior.*

PROPOSITION 4 (The Origin of Corporate Market Timing Ability): *In our model, corporate maturity choices forecast bond returns, so long as we are not in the limiting case where* $\theta = 0$. *In particular, when* f^* *is high, so that firms are tilting toward long-term debt, expected returns on long-term bonds are lower, and vice versa. However, the ability of* f^* *to forecast returns in this way arises because* f^* *endogenously responds to changes in the supply* g *of long-term government bonds, with* g *being the exogenous factor that drives variation in expected returns.*

One implication of Proposition 4 is that we would expect the forecasting power of corporate maturity choices for bond returns to be diminished if we also include a measure of government debt maturity in the forecasting regression. Indeed, if changes in g are the *only* source of variation in expected returns, the two variables f^* and g are completely colinear. More generally, if there are other sources of variation (e.g., shocks to target corporate maturity z, or to arbitrageur risk tolerance γ), then f^* may retain some incremental predictive power for bond returns, even controlling for g. The details of this more elaborate case are in the Internet Appendix.[6]

II. Taking the Model to the Data

In our baseline tests, we proxy for the variables g and f^* with data on the maturity of federal debt and all nonfinancial corporate debt, respectively. However, this mapping from the theory to the data raises a number of issues that merit further discussion.

A. Isolating Supply Effects in Government Debt Maturity

The model assumes that changes in government debt maturity represent exogenous supply shifts. In reality, it may be the case that both the government and firms respond endogenously to some other factor, such as changes in

[6] The Internet Appendix is available at http://www.afajof.org/supplements.asp.

investor demands. One such episode occurred in the United Kingdom, where the 2005 Pension Reform Act forced pension funds to mark their liabilities to market by discounting them at the yield on long-term bonds. This significantly increased their hedging demand for long-term securities. Greenwood and Vayanos (2010) describe how this demand flattened the U.K. yield curve:

> The effects of pension-fund demand on the shape of the term structure were immediate and dramatic. [...] The inversion appeared even more strongly on the 2055 bond, which was yielding 0.48%, an extremely low rate relative to the historical average of 3% of long real rates in the U.K.

In this case, the demand shock led both the government and firms to lengthen their maturities so as to exploit low long-term rates, thereby inducing a *positive* correlation between government and corporate debt maturity.[7] To the extent that such shocks are present more generally, they will tend to obscure the negative correlation suggested by our model. For example, the yield curve may steepen either because habitat investors demand fewer long-term bonds, or because the government's desired maturity has increased.

Ideally, therefore, we would like to have an instrument for government debt maturity that is uncorrelated with demand-side factors. Empirically, there is a powerful association between government maturity and the ratio of government debt to GDP: In our sample period, a univariate regression of the former on the latter yields an R^2 of 0.74. Thus, when the government's financing needs are greater, it extends its offerings out to longer maturities. This relationship leads Greenwood and Vayanos (2008) to use the debt-to-GDP ratio as an instrument for government maturity in a setting similar to ours, and we follow this approach below. Before doing so, however, it is useful to pause and ask *why* one might expect to see such a strong empirical connection between government debt maturity and the debt-to-GDP ratio; as far as we know, this connection is not clearly predicted by any existing formal theory.[8]

One informal hypothesis goes as follows. On the one hand, by financing on a short-term basis, the government can economize on the historically positive term premium. On the other hand, short-term financing requires more frequent rollovers. As the size of the government's debt increases, so too do the risks associated with larger and more frequent refinancings—for example, the possibility that a temporary dislocation in markets causes unexpectedly low investor turnout at an auction. An aversion to such risks would lead the government to extend its maturities as the stock of its debt goes up.

Former Treasury secretary Lawrence Summers describes government financing behavior along just these lines: "I think the right theory is that one tries

[7] According to data from the 2008 Blue Book (Table 3.1.9), the long-term corporate share in the United Kingdom peaked in 2005. The U.K. government reacted in the same direction, arguing: "Treasury concluded that there appeared to be an ongoing structural demand for such instruments and that it would be possible to issue ultra-long gilts at a favourable cost to the Government, given the inversion at the long-end of the gilt yield curve and the shortage of alternative instruments in this sector of the market." (Debt Management Office Annual Review 2005–2006).

[8] For theories of optimal government debt maturity, see, e.g., Roley (1979), Barro (1995), Angeletos (2002), and Guibaud, Nosbusch, and Vayanos (2007).

to [borrow] short to save money but not [so much as] to be imprudent with respect to rollover risk. Hence there is certain tolerance for [short-term] debt but marginal debt once [total] debt goes up has to be more long term."[9] If this account is on target, using the debt-to-GDP ratio as an instrument for government debt maturity would appear to be a well-motivated exercise, grounded in a specific model of government financing policy. And as we show below, this instrumental variables (IV) approach yields results that are very close to those obtained from an OLS specification.

B. Mortgages

Taken together, federal debt and nonfinancial corporate debt comprise an average of 40% of all credit market debt over our 1963 to 2005 sample period, according to the definition in Table L.1 of the *Flow of Funds*. Another important category is mortgage loans, which make up an additional 21% of credit market debt, and which our model does not speak to directly. How should we treat mortgages empirically?

One possibility is to add the stock of mortgages—or at least, those mortgages that are securitized and hence publicly traded—to the stock of long-term government debt, and to use this figure in computing a broader measure of supply shocks. This approach makes most sense to the extent that one thinks of the stock of mortgages as being: (1) predominantly long term in duration and (2) determined by factors orthogonal to those that influence corporate debt maturity, such as the relative costs of short-term and long-term borrowing. Although it is not clear that these assumptions are defensible, we experiment with the broader supply measure that aggregates mortgage debt and long-term government debt. Our results are robust to this variation.

Alternatively, one might hypothesize that there is an element of endogenous gap-filling behavior in the mortgage market, much like in the market for corporate debt—that is, mortgage borrowers might reduce the duration of their loans (by switching from fixed rate loans to adjustable rate loans) when they expect this to lower their borrowing costs. This hypothesis receives support in recent work by Koijen, Van Hemert, and Van Niewerburgh (2009), but we do not pursue it here as our focus is on understanding the determinants of corporate debt maturity.

C. Interest Rate Swaps

In our model, maturity matters in so far as it affects the interest rate sensitivity of debt claims, that is, their duration. Thus, if interest rate swaps alter the duration of corporate debt, they should in principle be counted in our maturity

[9] Private email correspondence, April 28, 2008. Also relevant is Garbade (2007), who emphasizes the Treasury's desire to minimize the uncertainties associated with the auction process. He notes that after 1975, Treasury officials explicitly renounced the concept of "tactical issuance" and replaced it with a policy of "regular and predictable" note and bond offerings. According to Garbade, "the move to regular and predictable issuance was widely credited with reducing market uncertainty, facilitating investor planning and lowering the Treasury's borrowing costs (p. 53)."

measures. Crucially, however, we care about the extent to which the corporate sector is a *net* buyer of swaps—if swaps only shift interest rate exposure around *within* the sector, they are irrelevant for our aggregate maturity measures.

Based on data compiled by Chernenko and Faulkender (2008), we are able to construct a net swap series for the interval 1993 to 2002. Although this does not cover our full sample period, it captures much of the relevant action since the interest rate swap market only came into existence in 1982 and remained small throughout the 1980s. As we show below, a swap-augmented version of our corporate debt maturity variable yields results very similar to those obtained when swaps are ignored. This is perhaps not too surprising in light of the fact that, even over 1993 to 2002, the net fraction of total corporate debt that was swapped from fixed to floating (or vice versa) was usually less than 2% in either direction and never more than 4%.

D. Financial Firms

In our baseline specifications, we focus on the maturity choices of nonfinancial firms. We do so for two reasons. First, this aligns us with previous work on the topic: BGW (2003), Butler et al. (2006), and Chernenko and Faulkender (2008) all study nonfinancial firms. Second, our prior is that asset-liability matching is likely to be more important for highly leveraged financial firms. Put in the language of the model, this means that financial firms have higher costs θ of deviating from their target capital structures, leading us to expect less aggressive gap-filling behavior among this group. Nevertheless, in our robustness tests we look at the debt maturity choices of financial firms. We find evidence of significant gap filling here too.

E. International Capital Market Integration

Our model envisions the U.S. capital market as segmented from the rest of the world, while in reality, markets are becoming increasingly integrated. There are two relevant dimensions of integration. The first is that foreigners can buy U.S. debt. This is not particularly problematic for our purposes—there is nothing in our model that requires the preferred-habitat investors to be U.S. based. Of course, foreigners may have specific patterns of demand that differ from those of local investors; for example, they may have a greater appetite for longer-duration debt. But as the U.K. pension fund example above suggests, such demand shifts can equally well arise locally, so this is something we have to contend with either way.

The second dimension of integration is that U.S. investors can buy foreign debt. This creates a form of measurement error for our regressions, to the extent that the relevant supply shocks are not changes in the maturity of U.S. government debt, but rather changes in the aggregate maturity of, say, the debt of all G7 countries. We do not attempt to fix this measurement error problem, and simply note that its effect will be to bias our estimates toward zero.

III. The Maturity of Corporate and Government Debt

In this section, we describe our proxies for corporate and government debt maturity. For government debt, we use the bond database from the Center for Research on Security Prices (CRSP). For corporate debt, we rely on two sources: the Federal Reserve *Flow of Funds* and Compustat. Because Compustat is available starting in 1963 and since many bond market studies (e.g., Fama and Bliss (1987), Cochrane and Piazzesi (2005)), start their forecasting in 1963 or 1964, we use 1963 to 2005 as our main period of study. However, the *Flow of Funds* data are available earlier, and thus many of our tests can be replicated on a longer sample; where applicable we mention these results.[10]

A. *Flow of Funds Data on Corporate Debt Maturity*

The Federal Reserve *Flow of Funds* tracks financial flows throughout the U.S. economy. We use annual data from the credit market liabilities of the nonfarm, nonfinancial corporate business sector (Table L. 102). This sector comprises all private domestic corporations except corporate farms, S-corporations, and real estate management corporations.

We follow BGW (2003) and define short-term corporate debt as the sum of "commercial paper," "bank loans not elsewhere classified," and "other loans and advances," all of which generally have a maturity of less than 1 year. Assuming these debts have a maturity of 1 year or less, they must have been issued in that year, and thus short-term debt issues ($d^C_{S,t}$) are the same thing as short-term debt outstanding ($D^C_{S,t}$). Throughout the paper, we follow the convention of level variables being denoted in upper case, and issue variables being denoted in lower case.

Long-term corporate debt ($D^C_{L,t}$) is the sum of "industrial revenue bonds," "corporate bonds," and "mortgages." BGW (2003) provide a detailed description of each of these items, as well as their shares in total long-term debt. Our first corporate debt maturity measure, the *long-term corporate level share*, is simply long-term corporate debt over total corporate debt ($D^C_{L,t}/D^C_t$).[11] As can be seen in the summary statistics in Table I, the level share based on *Flow of Funds* data is quite persistent, with a first-order autocorrelation of 0.85.

In the context of our static model, the most obvious way to test Proposition 1 is to regress the corporate level share on the analogous construct for government

[10] The *Flow of Funds* data are available as early as 1945. However, reliable estimates of government debt maturity based on CRSP cannot be constructed until the early 1950s. Furthermore, most studies focus on the period following the 1951 Fed–Treasury accord, prior to which interest rates were partially pegged. When we work with this longer sample, we follow BGW (2003) and begin in 1953.

[11] Alternatively, one might measure time variation in duration within a specific instrument category. Available data sources do not allow this, with the possible exception of long-term bonds. In any case, focusing on any individual category would miss broader shifts between, for example, long-term bonds and commercial paper that are well captured by our data. The term structure literature suggests that these broader shifts are more likely to be relevant from a market timing perspective (e.g., Piazzesi (2004)).

Table I

Summary Statistics

This table presents means, medians, standard deviations, extreme values, and time-series autocorrelations (denoted ρ) of variables between 1963 and 2005. Panel A shows the corporate long-term level share, and the corporate long-term issue share, based on *Flow of Funds* (FOF) data. All FOF short-term debt is assumed to be new short-term issues. FOF long-term issues are defined as the change in FOF long-term debt plus one-tenth of lagged FOF long-term debt. Panel B shows the corresponding levels measure from Compustat. Compustat debt is the sum of long-term debt and debt in current liabilities. Long-term debt is the sum of all long-term borrowings, plus debt that was originally issued long term but that is about to retire. Panel C summarizes measures of public debt maturity, estimated using the CRSP government bond database. D^G_L/D^G is the fraction of principal and coupon payments that is due in more than 1 year. M is the face value–weighted maturity of government bonds. Panel D summarizes interest rate conditions: y_{St} is the log yield on 1-year Treasuries, $y_{Lt} - y_{St}$ is the spread between the log yields of the 20-year Treasury bond and the 1-year Treasury bond, $R_{Lt+1} - y_{St}$ is the log 1-year forward excess bond return, and *Credit spread* is the Moody's Baa yield minus the average yield on long-term Treasuries. Panel E summarizes the ratio of government debt to GDP; the ratio of government debt to total credit market debt; annual GDP growth; a recession dummy based on NBER dating conventions; the ratio of government investment to GDP; and net private investment, calculated as the percentage change in net private property, plant, and equipment. The last two variables are calculated using data from the Bureau of Economic Analysis. All variables, except for M and the *Recession dummy*, are expressed in percentage terms.

	Mean	Median	*SD*	Min	Max	ρ
Panel A: FOF Corporate Debt Maturity %						
Levels: D^C_L/D^C	61.51	60.93	4.97	53.46	73.12	0.85
Issues: d^C_L/d^C	21.57	21.28	4.14	14.75	30.13	0.58
Panel B: Compustat Corporate Debt Maturity %						
Levels: D^C_L/D^C	83.41	83.69	3.36	77.00	89.75	0.76
Panel C: Gov. Debt Maturity %						
D^G_L/D^G	59.09	58.78	8.94	41.74	72.48	0.95
M (years)	4.51	4.57	0.90	2.82	5.75	0.96
Panel D: Short Rate, Term Spread, Subsequent Bond Returns, and Credit Spread (%)						
y_{St}	6.01	5.41	2.99	0.96	16.86	0.74
$y_{Lt} - y_{St}$	0.87	0.73	1.41	−1.60	3.75	0.63
$R_{Lt+1} - y_{St}$	0.98	0.22	9.81	−15.21	21.01	−0.10
Credit spread	2.01	1.84	0.80	0.59	3.85	0.69
Panel E: Other Controls (%)						
D^G/GDP	34.63	34.08	7.73	22.46	48.67	0.96
D^G/D	19.62	19.73	3.11	15.20	26.96	0.85
$\Delta Log(GDP)$	3.25	3.48	1.99	−1.95	6.94	0.25
Recession dummy	0.26	0.00	0.44	0.00	1.00	0.38
GovInv/GDP	1.42	1.27	0.47	0.89	2.60	0.86
$CorpInv/PPE_{t-1}$	7.29	6.30	3.49	2.13	17.80	0.76

bonds—the fraction of government debt due in more than 1 year. While this is where we begin, two considerations lead us to also examine the maturity of corporate *issues*. First, in a more realistic dynamic setting, where adjustment costs prevent firms from recasting their balance sheets overnight, equilibrium involves a partial adjustment mechanism, whereby it is corporate *issuance* that responds at the margin to the expected return differentials induced by the relative stocks of long-term and short-term government debt.[12] Second, looking at issuance helps to resolve some of the econometric concerns associated with the high degree of persistence in the levels variable.

Accordingly, we construct long-term debt *issues* ($d^C_{L,t}$) as the change in the level of long-term corporate debt outstanding ($D^C_{L,t}$), plus one-tenth the level of long-term debt in the previous year. That is, we have

$$d^C_{L,t} = \left(D^C_{L,t} - D^C_{L,t-1}\right) + 0.1 \times D^C_{L,t-1}. \tag{7}$$

This amounts to assuming that long-term debt has an average maturity of 10 years, which roughly corresponds to the figure in Guedes and Opler (1996). Our results are not at all sensitive to this assumption, however.

Total corporate debt issues, d^C_t, is the sum of long- and short-term issues. Our second corporate maturity measure, the *long-term corporate issue share*, is the ratio of long-term issues to total issues ($d^C_{L,t}/d^C_t$). Not surprisingly, the issue share closely tracks the level share, with a time-series correlation of 0.75. Nevertheless, the issue share is substantially less persistent, with a first-order autocorrelation of 0.58, compared to 0.85 for the level share.

B. Compustat Data on Corporate Debt Maturity

Compustat is a second source of data for corporate debt maturity. The advantage of the Compustat data is that it can be disaggregated; this makes it indispensable for our cross-sectional tests of Proposition 3. However, it also has an important limitation. Because it focuses only on public firms, time variation in a Compustat-based measure of aggregate debt maturity will be influenced by compositional effects.[13]

Since compositional effects are likely to be especially problematic for higher frequency movements, when working with Compustat we restrict attention to a levels measure of debt maturity, and do not attempt to construct an issues measure. For the sake of comparability, we construct our Compustat levels measure to correspond as closely as possible to the *Flow of Funds*

[12] Several recent papers emphasize the importance of adjustment costs for firms' capital structure decisions. See, e.g., Leary and Roberts (2005) and Strebulaev (2007).

[13] For example, suppose that in year t there are 100 private firms with zero long-term debt, and 100 public firms (of the same size) with 50% long-term debt. Suppose further that no firm alters its capital structure in year $t + 1$ (so that a *Flow of Funds*–type measure remains constant) but that 10 of the private firms go public. The measured long-term debt share based on public firm data would drop from 50% to 50/110 = 45%. According to Fama and French (2004), between 1980 and 2001 an average of 10% of public firms were new lists in a given year. So, compositional effects of this sort have the potential to be quantitatively significant.

long-term level share. Aggregating across all nonfinancial firms, we define long-term debt as the sum of all long-term borrowings (item 9) plus debt that was originally issued long term but that is about to retire (item 44). We define short-term debt as total debt (item 9 plus item 34) minus long-term debt. Our convention of counting the current portion of long-term debt as long term is meant to replicate the procedure used in the *Flow of Funds*, whereby corporate bonds are classified as long-term instruments even though some portion of these bonds may, at any point in time, have a short remaining duration.

Over the 1963 to 2005 period, the Compustat long-term level share is generally higher than the corresponding *Flow of Funds* series; the means of the two series are 83.4% and 61.5%, respectively. We suspect that this is because Compustat firms, which are public, have better access to longer-term financing instruments—an observation that reinforces the above concern about compositional effects. At an annual frequency, the two variables have a correlation of 0.41. This correlation is generally higher later in the sample period, and higher still if one nets out a time trend in the Compustat measure.

C. CRSP Data on Government Debt Maturity

The available data on government bonds allow for a much finer characterization of debt maturity structure than we are able to obtain for firms. Nevertheless, we stick with a simple measure that matches our corporate maturity variable: the fraction of government debt with a maturity of 1 year or more, hereafter the *long-term government share*.

To construct the long-term government share, we follow Greenwood and Vayanos (2008). The CRSP U.S. Treasury Database reports detailed information on every Treasury security that was outstanding between 1925 and 2006. For each security, CRSP reports a number of characteristics, including the issue date, final maturity, and callability features. CRSP also provides monthly readings of the dollar face value of each instrument. Changes in face value reflect repurchases, as well as follow-on offerings (or "reopenings") of an existing issue.

We decompose the payment stream of each outstanding issue into a series of principal and coupon repayments. In each month, these series are adjusted for variation in the face value outstanding. Every month, we aggregate payments due in the subsequent n periods, across all issues that are still outstanding. The government long-term share ($D^G_{L,t}/D^G_t$) is then defined as total payments due in more than 1 year, divided by total payments in all future periods.

To ensure robustness, we also rerun some of our basic specifications with a second measure of government debt maturity: the dollar-weighted average maturity of principal payments, which we denote by M. As can be seen in Table I, both of these variables are highly persistent, with first-order autocorrelations on the order of 0.95.

D. Other Variables

Our tests below use several other variables, also summarized in Table I: the short-term (1-year) Treasury yield y_{St}; the spread between the long-term (20-year) Treasury yield and the short-term yield, $(y_{Lt} - y_{St})$; the 1-year excess log return on long-term Treasuries, $(R_{Lt+1} - y_{St})$; the credit spread, defined as the Moody's Baa yield minus the yield on long-term Treasuries; the ratio of government debt to GDP; the ratio of government debt to total credit market debt; annual GDP growth, $(\Delta\text{Log(GDP)})$; a recession dummy based on NBER dating conventions; the ratio of government investment to GDP; and net private investment, calculated as the percentage change in net private property, plant, and equipment.

IV. Proposition 1: Gap Filling

A. Univariate Tests

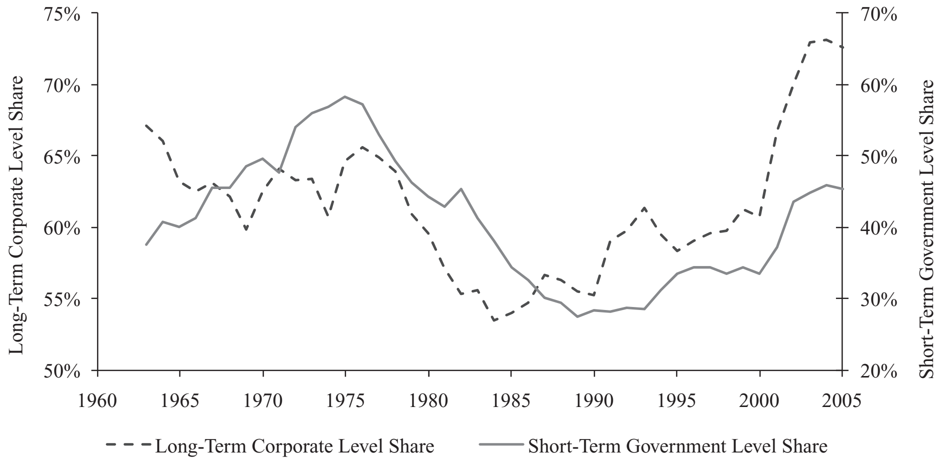
The primary prediction of our theory, Proposition 1, is that when the government lengthens the maturity profile of its debt, firms respond by doing the opposite. Panels A to C of Figure 1 present a first look at this prediction. In Panel A, we plot the *Flow of Funds* long-term corporate level share against one minus the government long-term share. Given this transformation of the government share variable, our hypothesis is that the two series in the figure should be positively correlated. In Panels B and C, we replace the *Flow of Funds* level share with the *Flow of Funds* issue share and the Compustat level share, respectively. In all three cases, the correlation between corporate and government debt maturity is readily apparent.

Table II presents a set of univariate OLS regressions corresponding to Figure 1. We separately regress each of our three measures of corporate debt maturity against either: (1) the government long-term share or (2) the weighted maturity M of government debt. In these regressions, we do not invert the government variables, so we expect to see negative correlations. Since all of the underlying series are persistent, we report Newey-West (1987) standard errors, which are robust to serial correlation at up to two lags.

In all six regressions, we obtain the predicted negative coefficients. The results for both the *Flow of Funds* level share and the *Flow of Funds* issue share are strongly statistically significant, with t-statistics ranging from 2.64 to 4.21. The results for the Compustat level share are statistically marginal, with t-statistics of 1.83 and 1.67.

In terms of economic magnitudes, the regression coefficients in the first and third columns of Table II (-0.262 and -0.249) imply that when the fraction of U.S. Treasury debt longer than 1 year rises by 10%, the long-term corporate share based on *Flow of Funds* falls by about 2.5%; this holds in both levels and issues. To understand what this means for total gap filling, we can multiply this by the average ratio of corporate debt to government debt during the sample period of 1.09, which yields 2.7%. This suggests that on a dollar-for-dollar basis, firms fill 27% of the gap created by variation in government debt maturity.

Panel A. *Flow of Funds* Levels



Panel B. *Flow of Funds* Issues

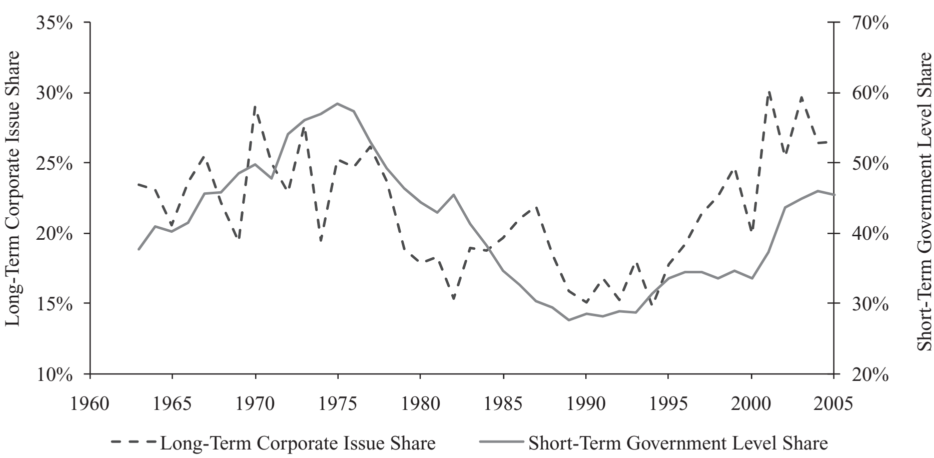


Figure 1. Corporate and government debt maturity, 1963–2005. The dashed line, plotted on the left axis, is the share of long-term corporate debt as a fraction of total debt. The solid line, plotted on the right axis, is the share of government debt with maturity of 1 year or less. Panel A shows the corporate long-term level share based on *Flow of Funds* data. Panel B shows the corporate long-term issue share based on *Flow of Funds* data. Panel C shows the corporate long-term level share based on Compustat data.

B. Multivariate Tests

In Table III, we take the univariate regressions from Table II and add a set of further controls: (1) the short-term Treasury yield y_{St}; (2) the term spread $(y_{Lt} - y_{St})$; and (3) a linear time trend. Why might these controls be useful? A simple story is that for reasons outside the model, firms' debt maturity choices

Panel C. *Compustat* Levels

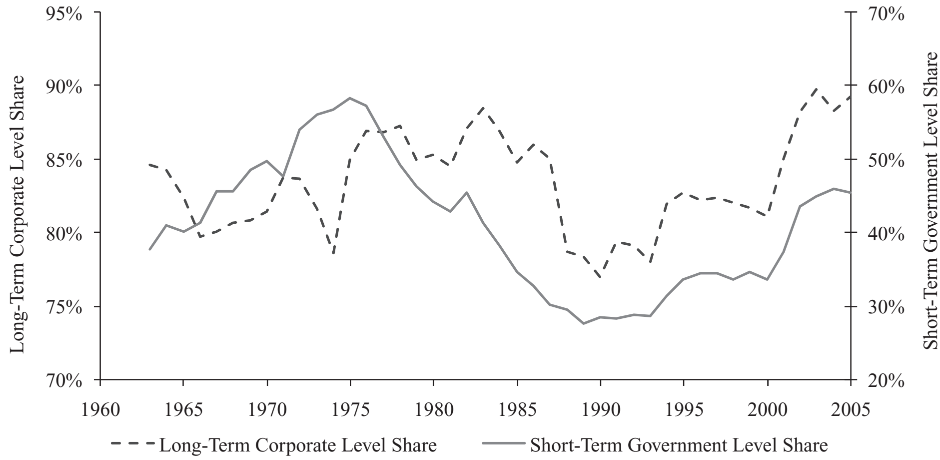


Figure 1. Continued

Table II

The Maturity of Corporate and Government Debt, 1963–2005: Univariate Regressions

This table presents OLS regressions of the maturity of corporate debt on the maturity of government debt. The dependent variable is alternately the *Flow of Funds* (FOF) corporate long-term level share, the FOF corporate long-term issue share, or the Compustat corporate long-term level share. The maturity of government debt is defined as either the share of government debt and coupon payments with maturity of 1 year or more (D_L^G/D^G), or the dollar-weighted maturity of principal payments (M). t-statistics, in brackets, are based on Newey-West (1987) standard errors allowing for 2 years of lags.

	FOF: Levels		FOF: Issues		Compustat: Levels	
	(1)	(2)	(3)	(4)	(5)	(6)
D_L^G/D^G	−0.262		−0.249		−0.147	
	[−3.64]		[−4.21]		[−1.83]	
M		−1.804		−1.949		−1.272
		[−2.64]		[−2.85]		[−1.67]
Constant	76.971	69.656	36.286	30.372	92.090	89.148
	[16.58]	[22.08]	[10.16]	[10.42]	[18.55]	[24.62]
R^2	0.22	0.11	0.29	0.18	0.15	0.12

respond not only to the expected term premium, but also to movements in the level of rates and the shape of the term structure that are orthogonal to the expected term premium.[14] This could reflect accounting-driven considerations about current interest costs, as in Faulkender (2005) and Chernenko and

[14] Similarly, the inclusion of a time trend could capture secular shifts in firms' ability to access the short-term debt markets that are unrelated to term premia.

Table III

The Maturity of Corporate and Government Debts, 1963–2005: Multivariate Regressions

This table presents OLS regressions of the maturity of corporate debt on the maturity of government debt, controlling for the short-term rate, the term spread, and a time trend. The dependent variable is alternately the *Flow of Funds* (FOF) corporate long-term level share, the FOF corporate long-term issue share, or the Compustat long-term level share. The maturity of government debt is defined as either the share of government debt and coupon payments with maturity of 1 year or more (D_L^G/D^G), or the dollar-weighted maturity of principal payments (*M*). *t*-statistics, in brackets, are based on Newey-West (1987) standard errors allowing for 2 years of lags.

	FOF: Levels				FOF: Issues				Compustat: Levels			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
D_L^G/D^G	−0.296	−0.387			−0.278	−0.318			−0.169	−0.228		
	[−5.14]	[−5.45]			[−5.00]	[−5.77]			[−1.96]	[−2.33]		
M			−2.540	−3.488			−2.526	−2.939			−1.474	−2.094
			[−4.31]	[−4.03]			[−4.68]	[−5.64]			[−1.80]	[−2.05]
y_{St}	−1.214	−1.263	−1.317	−1.404	−0.815	−0.836	−0.919	−0.957	0.155	0.123	0.095	0.038
	[−2.93]	[−3.55]	[−2.87]	[−3.43]	[−5.02]	[−5.81]	[−4.97]	[−5.45]	[0.60]	[0.48]	[0.35]	[0.14]
$y_{Lt} - y_{St}$	−0.613	−1.257	−0.781	−1.436	−0.207	−0.486	−0.355	−0.641	0.919	0.504	0.824	0.395
	[−1.11]	[−2.72]	[−1.30]	[−2.94]	[−0.48]	[−1.08]	[−0.74]	[−1.31]	[1.77]	[0.88]	[1.48]	[0.66]
Trend		0.160		0.154		0.069		0.067		0.103		0.101
		[2.26]		[1.78]		[2.07]		[1.71]		[1.62]		[1.38]
Constant	86.825	89.736	81.581	83.723	43.109	44.369	38.807	39.742	91.679	93.550	88.765	90.168
	[17.14]	[22.70]	[16.76]	[21.30]	[13.74]	[14.83]	[15.25]	[17.11]	[16.14]	[17.26]	[18.93]	[20.60]
R^2	0.63	0.73	0.55	0.64	0.59	0.61	0.52	0.54	0.25	0.34	0.20	0.29

Faulkender (2008). To take a concrete example, suppose that the expectations hypothesis always holds exactly—so that the term premium is always zero—but that the yield curve gets steeper, reflecting expectations of future short rate increases. A firm concerned with maximizing current earnings may be inclined to shift its borrowing to shorter maturities because this allows it to have lower stated interest expense. In this setting, adding interest rate controls absorbs some of the unexplained variation in corporate behavior, thereby improving the precision of our estimates.

As can be seen in the table, the addition of these controls makes the coefficient on government debt maturity more statistically significant in all cases. For example, in the regression of the *Flow of Funds* level share against the government level share, the t-statistic is 5.14, as compared to its value of 3.64 in the univariate specification. The regression coefficient in this specification is virtually unchanged (−0.296, as compared to −0.262 in the univariate specification). Adding a time trend, the coefficient increases in absolute magnitude to −0.387 (t-statistic of 5.45). The controls have a similar effect in the regressions explaining the *Flow of Funds* issue share and the Compustat level share.

C. Robustness

Table IV presents a number of robustness checks on the multivariate results of Table III. There are three columns, corresponding to our three measures of corporate debt maturity. In the first row, we reproduce our baseline estimates from Table III using the government level share as the key explanatory variable and including the full set of controls. (These baseline estimates correspond to columns 2, 6, and 10 of Table III.) In the second and third rows, we display subsample estimates. As can be seen, the results are generally stronger, both economically and statistically, in the second half of the sample, which runs from 1984 to 2005. The differences across sample periods are relatively modest with the two *Flow of Funds* measures of corporate debt maturity, but are striking with the Compustat measure; in this case the point estimate is very large and significant in the post-1984 period (−0.787, with a t-statistic of 11.68), but actually goes the wrong way in the first half of the sample. We suspect that this divergence may have something to do with the fact that Compustat offers less complete coverage of the entire (public plus private) universe during the earlier period.

In the fourth row of Table IV, we extend the sample for the *Flow of Funds* measures further back in time, so that it covers 1953 to 2005. (Again, we are unable to go back further than 1963 with the Compustat data.) The results are qualitatively similar to those from our baseline sample period of 1963 to 2005, albeit a bit smaller in absolute magnitude.

In the fifth through seventh rows of Table IV, we add a number of further controls for general economic and credit market conditions. The motivation is that some firms may find it difficult to issue long-term debt during periods of weak economic growth, or when credit spreads are high. In the fifth row, we add two leads and two lags of an NBER recession dummy to the previous

Table IV

Robustness Checks

This table presents regressions of the maturity of corporate debt on the maturity of government debt. We vary the basic specification in row (1) by: (2) using only the first half of the sample period; (3) using only the second half of the sample period; (4) extending the *Flow of Funds* (FOF) data back to 1953; (5) controlling for two leads and two lags of the NBER recession dummy; (6) controlling for two leads and two lags of changes in log GDP; (7) controlling for the credit spread, defined as the Moody's Baa yield minus the average yield on long-term Treasuries; (8) controlling for government investment scaled by total government fixed capital; (9) controlling for aggregate corporate investment; (10) adjusting our corporate maturity measure for marketable mortgages, which we count as long-term debt; (11) adjusting our corporate maturity measure for net swap activity; and (12) using Compustat financial firms only. In the last two rows, we instrument for government debt maturity using the ratio of government debt to GDP where (13) presents instrumental variable estimates for our baseline specification, and (14) adds the credit spread as a control. For each regression, we report the slope coefficient on the long-term government share, b, and its associated t-statistic. All regressions include a constant term and controls for the short-term rate, the term spread, and a time trend. t-statistics are based on Newey-West (1987) standard errors allowing for 2 years of lags.

	FOF: Levels			FOF: Issues			Compustat: Levels		
	b	$[t]$	R^2	b	$[t]$	R^2	b	$[t]$	R^2
OLS:									
(1) Baseline	−0.387	[−5.45]	0.73	−0.318	[−5.77]	0.61	−0.228	[−2.33]	0.34
(2) First half (1963–1983)	−0.203	[−1.84]	0.58	−0.266	[−3.34]	0.52	0.163	[1.88]	0.58
(3) Second half (1984–2005)	−0.371	[−3.63]	0.95	−0.477	[−2.97]	0.71	−0.787	[−11.68]	0.92
(4) Long sample (1953–2005)	−0.303	[−3.99]	0.70	−0.269	[−4.72]	0.60	–	–	–
(5) Business cycle leads and lags	−0.383	[−4.64]	0.74	−0.381	[−8.13]	0.77	−0.305	[−4.02]	0.62
(6) ΔLog(GDP) leads and lags	−0.373	[−5.20]	0.74	−0.383	[−6.30]	0.73	−0.280	[−3.37]	0.57
(7) Control for credit spread	−0.466	[−7.52]	0.76	−0.328	[−5.80]	0.61	−0.308	[−3.34]	0.41
(8) Control for Gov. investment	−0.532	[−5.12]	0.76	−0.418	[−5.11]	0.64	−0.312	[−3.23]	0.37
(9) Control for corp. investment	−0.367	[−4.37]	0.73	−0.338	[−4.41]	0.61	−0.332	[−2.58]	0.38
(10) Adjust for mortgages	−0.509	[−4.00]	0.61	−0.434	[−5.00]	0.52	−0.293	[−1.87]	0.25
(11) Adjust for swaps	−0.309	[−5.83]	0.73	–	–	–	−0.149	[−1.70]	0.21
(12) Compustat financial firms	–	–	–	–	–	–	−0.298	[−2.06]	0.63
Instrumental Variables:									
(13) Baseline	−0.395	[−4.89]	0.73	−0.402	[−5.33]	0.59	−0.242	[−2.59]	0.34
(14) Control for credit spread	−0.539	[−6.41]	0.75	−0.488	[−4.87]	0.55	−0.392	[−4.35]	0.39

regression. In the sixth row, we control for leads and lags of GDP growth. In the seventh row, we control for the credit spread, defined as the Moody's Baa yield minus the average yield on long-term Treasuries.[15] In all cases, our results are either similar to those from the baseline specification or somewhat stronger.

In the eighth row, we control for the ratio of government investment to government capital stock. We do so in order to address the following concern. Suppose that when public investment is high, the government funds itself with long-term debt. If public investment crowds out long-term private investment, then our results may be due to changes in the composition of corporate investment (that is, changes in the mix of plant and equipment investment relative to inventories), rather than to the effects described in our model. However, it turns out that public investment is uncorrelated with corporate debt maturity and, when included in our baseline regression, does not materially affect the correlation between corporate and government maturities (see Internet Appendix). Following the same motivation, in the ninth row we control for net corporate investment in property, plant, and equipment. Again, our results are little changed. We also interact government investment with proxies for business cycle conditions, under the theory that investment crowding out would be more pronounced in booms; again, this additional control does not change our conclusions.

In the tenth row, we adjust our government maturity variable by adding marketable mortgage debt (i.e., debt and mortgage securities issued or backed by government-sponsored enterprises) to long-term government debt. As mentioned earlier, one may perhaps want to combine mortgages with government debt to form a broader measure of supply shocks. If anything, this leads to slightly elevated point estimates relative to our baseline specification.

The eleventh row shows results that adjust the corporate maturity measures for net swap activity. We obtain data on the net fraction of long-term corporate debt swapped to floating from Chernenko and Faulkender (2008).[16] This adjustment attenuates our results, but not by much.

In the twelfth row, we compute the Compustat long-term corporate share for financial firms, defined as those having an SIC code between 6000 and 7000. The results suggest that, like their nonfinancial counterparts, financial firms also engage in significant gap filling.

Finally, we attempt to address concerns about the potential endogeneity of government debt maturity. As described above, one approach to doing so is to use the ratio of government debt to GDP as an instrument for government debt maturity. Accordingly, in the thirteenth row of Table IV, we return to the baseline specification of the first row, but we estimate the regression by IV

[15] Kessel (1965) discusses the relationship between credit market conditions and economic growth.

[16] Chernenko and Faulkender (2008) collect swap data for 1993 to 2002. We assume that net corporate swap activity was zero from 1963 to 1992 and that it remained at its 2002 level from 2003 to 2005. The adjusted series are computed by subtracting this estimate of the net swap activity of the corporate sector from our baseline series.

instead of by OLS. As can be seen, this produces estimates that are very close to those from the corresponding OLS specifications. For example, with the *Flow of Funds* level share as the dependent variable, IV yields a point estimate of -0.395 (t-statistic of 4.89), as compared to the OLS estimate of -0.387.

One concern with the IV approach is whether our instrument, the debt-to-GDP ratio, satisfies the exclusion restriction. For example, one might worry that high values of the debt-to-GDP ratio are associated with adverse credit market conditions, and hence influence corporate debt maturity through another channel. In an effort to address this issue, we run an augmented version of the IV specification that adds a control for the credit spread. The results, shown in the last row of Table IV, are somewhat stronger than those without the credit spread control.

D. Differenced and GLS Specifications

As emphasized above, our measures of corporate and government debt maturity are highly persistent. One way to address this persistence is to compute adjusted standard errors that take it into account, as we have been doing throughout. Alternatively, the classic prescriptions for persistence are either to estimate the regression in first differences, or to use a generalized least squares (GLS) estimator that addresses the serial correlation of regression residuals. We try both of these techniques below. In each case, however, we have to be mindful of the risk of over-differencing. Specifically, in a world where issuance costs and other frictions create lags in the adjustment process, it might be unrealistic to expect an innovation in government debt maturity in year t to be met with the full response of corporate debt maturity in the same year t—rather, it might take a few years for the adjustment process to play itself all the way out.

In the left-most panel of Table V, we use the *Flow of Funds* issue share to estimate specifications of the form

$$d^C_{L,t}/d^C_t = a + b \cdot \Delta_k\left(D^G_{L,t}/D^G_t\right) + \varepsilon_t, \tag{8}$$

where $\Delta_k(D^G_{L,t}/D^G_t)$ represents the *cumulative* change in the long-term government share variable over the past k years, for $k = 1, 2, 3, 4,$ and 5. Thus, these specifications explore how corporate issues respond not to the *level* of the government share, but instead to *recent changes* in the government share. This is one simple approach to differencing. When the differencing window is only 1 year, the results are statistically weaker than when the government share is entered in levels form. However, as we broaden the differencing window out to 2 years and beyond, the results again become strongly significant. By the time the window reaches 5 years, the estimated value of b is -0.289, with a t-statistic of 5.10. Thus, while the response of corporate issues to changes in government debt maturity is not entirely contemporaneous, it appears that our earlier results reflect something more than the juxtaposition of very low-frequency trends in the two series.

Table V

Differenced Regressions

This table presents regressions of the following form:

$$d^C_{L,t}/d^C_t = a + b \cdot \Delta_k\left(D^G_{L,t}/D^G_t\right) + \varepsilon_t$$

$$\Delta_k\left(D^C_{L,t}/D^C_t\right) = a + b \cdot \Delta_k\left(D^G_{L,t}/D^G_t\right) + \varepsilon_t.$$

The dependent variable is alternately $d^C_{L,t}/d_t$, the *Flow of Funds* (FOF) long-term corporate issue share, or $\Delta_k(D^C_{L,t}/D^C_t)$, the change in either the FOF or the Compustat long-term corporate level share over a k-year window. The independent variable is $\Delta_k(D^G_{L,t}/D^G_t)$, the change in the long-term government share over a k-year window. t-statistics are based on Newey-West (1987) standard errors allowing for k years of lags in the regressions on $\Delta_k(D^G_{L,t}/D^G_t)$.

	FOF Issues			Changes in FOF Levels			Changes in Compustat Levels		
	b	$[t]$	R^2	b	$[t]$	R^2	b	$[t]$	R^2
$k = 1$ lag	−0.309	[−1.33]	0.04	−0.179	[−1.30]	0.06	−0.211	[−1.84]	0.07
$k = 2$ lags	−0.331	[−2.26]	0.12	−0.265	[−1.64]	0.13	−0.273	[−2.10]	0.13
$k = 3$ lags	−0.287	[−2.65]	0.16	−0.282	[−1.63]	0.16	−0.237	[−1.78]	0.13
$k = 4$ lags	−0.285	[−3.92]	0.25	−0.308	[−1.88]	0.21	−0.228	[−1.78]	0.16
$k = 5$ lags	−0.289	[−5.10]	0.33	−0.325	[−1.94]	0.24	−0.230	[−1.76]	0.19

In the second and third panels of Table V, we alternately use the *Flow of Funds* and Compustat level shares to estimate specifications of the form

$$\Delta_k\left(D^C_{L,t}/D_t\right) = a + b \cdot \Delta_k\left(D^G_{L,t}/D^G_t\right) + \varepsilon_t. \tag{9}$$

This is just a differenced version of our baseline levels specification, with the differencing window again varying from 1 to 5 years. For the *Flow of Funds* level share the results are statistically weak when using a 1-year window, but grow progressively stronger as the window is widened. With a 5-year window, the estimate of b is −0.325, with a t-statistic of 1.94. By contrast, the results for the Compustat level share are of roughly similar significance for all values of k.

In untabulated regressions (see the Internet Appendix), we also explore the lead-lag properties of the relation between government and corporate maturities. Specifically, in bivariate vector autoregressions we find a negative and significant relation between the current corporate issue share (or changes in the corporate level share) and lagged changes in the government level share. However, there is no significant relation between current changes in the government level share and the lagged corporate issue share. That is, changes in government maturities appear to Granger-cause changes in corporate maturities. This lead-lag asymmetry further alleviates possible concerns about reverse causation.

In Table VI, we report Prais-Winsten (1954) GLS estimates of the univariate and multivariate specifications from Tables II and III under the assumption

Table VI

GLS Regressions

This table presents Prais–Winsten GLS regressions of the maturity of corporate debt on the maturity of government debt, controlling for the short-term rate, the term spread, and a time trend. The Prais-Winsten GLS procedure is based on the assumption that regression residuals follow an AR(1). The dependent variable is alternately the *Flow of Funds* (FOF) corporate long-term level share, the FOF corporate long-term issue share, or the Compustat long-term level share. The maturity of government debt is defined as the share of government debt and coupon payments with maturity of 1 year or more (D_L^G/D^G). The t-statistics, in brackets, are computed using Prais–Winsten GLS standard errors that are robust to heteroskedasticity (equivalent to robust OLS standard errors for the transformed regression). We also report the estimated first-order autocorrelation of the residuals, denoted by ρ.

	FOF: Levels			FOF: Issues			Compustat: Levels		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
D_L^G/D^G	−0.187	−0.130	−0.130	−0.238	−0.276	−0.316	−0.212	−0.194	−0.209
	[−1.44]	[−1.20]	[−1.21]	[−2.60]	[−4.70]	[−6.16]	[−1.73]	[−1.85]	[−2.05]
y_{St}		−0.290	−0.300		−0.780	−0.826		−0.274	−0.259
		[−1.14]	[−1.15]		[−3.78]	[−4.60]		[−1.14]	[−1.04]
$y_{Lt} - y_{St}$		0.299	0.284		−0.100	−0.437		0.155	0.145
		[0.79]	[0.73]		[−0.21]	[−1.02]		[0.40]	[0.36]
Trend			0.101			0.066			0.087
			[0.65]			[1.95]			[1.16]
Constant	77.087	75.492	71.445	35.753	42.698	43.478	96.499	96.886	94.831
	[7.79]	[9.72]	[9.22]	[6.18]	[11.97]	[14.06]	[12.51]	[14.32]	[14.33]
R^2	0.62	0.66	0.73	0.25	0.53	0.59	0.87	0.88	0.90
ρ	0.96	0.97	0.96	0.43	0.16	0.05	0.80	0.82	0.77

that the regression residuals follow an AR(1). The middle panel presents results for the *Flow of Funds* issue share. These results are almost identical to those obtained using OLS. For example, with the full set of controls, we obtain a GLS estimate of −0.316 for the coefficient on the government share, with a t-statistic of 6.16; this compares with an OLS estimate of −0.318 (t-statistic of 5.77) for the corresponding regression in Table III. Thus, for the *Flow of Funds* issue variable, our results are entirely robust to using GLS.

The GLS procedure makes less sense when using the *Flow of Funds* level share. This is illustrated in the left-hand panel of Table VI. As can be seen, the high persistence of the levels variable leads to an estimated value of ρ on the order of 0.96 in the GLS procedure. Hence, in this case, GLS is essentially identical to first-differencing the data. And, as seen in Table V, running the *Flow of Funds* levels regressions in first differences leads to insignificant results, for the reasons developed above. Given that ρ is estimated to be almost 1, the GLS results for *Flow of Funds* levels in Table VI amount to no more than a restatement of this prior finding. Note that GLS is *not* redundant in the same way when the dependent variable is the *Flow of Funds* issue share; in this case, the estimated value of ρ ranges from 0.05 to 0.43, so GLS is quite distinct from first differences.

Finally, the GLS results for the Compustat level share, shown in the right-hand panel of Table VI, represent an intermediate case between those for the

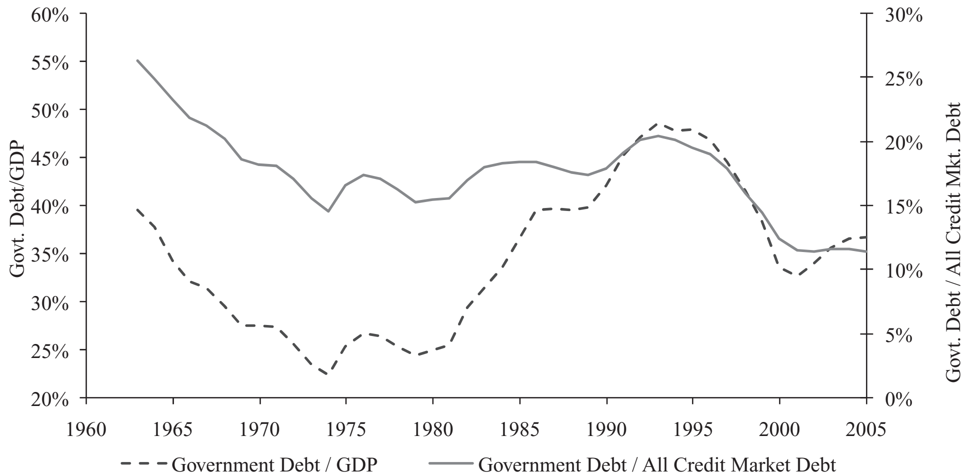


Figure 2. Debt market size, 1963–2005. The dashed line shows the ratio of government debt to GDP. The solid line shows the ratio of government debt to total credit market debt. Total outstanding Treasury securities and total credit market debt are from Table L.4 of the *Flow of Funds*. In addition to Treasury securities, credit market debt includes open market paper, municipal securities, corporate and foreign bonds, bank loans not elsewhere classified, other loans and advances, mortgages, and consumer credit. Data on nominal GDP are from the Bureau of Economic Analysis.

two *Flow of Funds* variables. In this specification, we estimate ρ to be 0.80, so that while there is a good deal of persistence, GLS is not literally the same thing as first-differencing the data. And as can be seen, the GLS results for the Compustat level share look very similar, in both magnitude and statistical significance, to their OLS counterparts in Tables II and III.

V. Proposition 2: Time Variation in Gap Filling

Our model predicts that when government debt supply is large, gap filling by firms will be quantitatively stronger. To test this hypothesis, we consider two proxies for the size of the government bond market. The first is the ratio of government debt to GDP, and the second is the ratio of government debt to total credit market debt; these two variables are plotted in Figure 2. In each case, we use the *Flow of Funds* long-term corporate issue share as our dependent variable, and we run regressions of the form

$$\begin{aligned} d^C_{L,t}/d^C_t &= a + b \cdot \left(D^G_{L,t}/D^G_t\right) + c \cdot Scale_t + d \cdot \left(Scale_t \times D^G_{L,t}/D^G_t\right) \\ &\quad + e \cdot time + f \cdot \left(time \times D^G_{L,t}/D^G_t\right) + \boldsymbol{\theta}'\mathbf{x}_t + \varepsilon_t, \end{aligned} \tag{10}$$

where $Scale_t$ denotes one of our two measures of the size of the government bond market described above, *time* is a linear trend, and $\mathbf{x}_t$ is a set of controls for debt market conditions (yield spread and the short-term bond yield). The coefficient

Table VII

The Effect of Government Bond Market Size on Gap-Filling Intensity

This table presents regressions of the following form:

$$d^C_{L,t}/d^C_t = a + b \cdot \left(D^G_{L,t}/D^G_t\right) + c \cdot Scale_t + d \cdot \left(Scale_t \times D^G_{L,t}/D^G_t\right) + e \cdot time + f \cdot \left(time \times D^G_{L,t}/D^G_t\right) + \boldsymbol{\theta}'\mathbf{x}_t + \varepsilon_t.$$

The dependent variable is the *Flow of Funds* corporate issue share. $Scale_t$ is either the ratio of government debt to GDP, or the ratio of government debt to total credit market debt. t-statistics, in brackets, are based on Newey-West (1987) standard errors allowing for 2 years of lags.

	Scale = Gov. Debt to GDP		Scale = Gov. Debt to Total Debt	
	(1)	(2)	(3)	(4)
D^G_L/D^G	0.640	0.351	1.272	0.768
	[2.79]	[1.49]	[3.07]	[2.15]
Scale	2.906	0.957	5.427	2.495
	[4.41]	[1.32]	[3.85]	[1.89]
$(Scale) \times (D^G_L/D^G)$	−4.400	−1.941	−8.397	−4.794
	[−4.49]	[−1.77]	[−3.83]	[−2.37]
Time	−0.916	−0.268	0.503	0.200
	[−1.69]	[−0.68]	[1.19]	[0.80]
$Time \times (D^G_L/D^G)$	0.017	0.006	−0.005	−0.003
	[1.77]	[0.87]	[−0.71]	[−0.77]
y_{St}		−0.920		−0.904
		[−4.08]		[−4.66]
$y_{Lt} - y_{St}$		−0.293		−0.322
		[−0.58]		[−0.70]
Constant	−26.504	12.442	−65.239	−10.706
	[−2.09]	[0.76]	[−2.55]	[0.44]
R^2	0.52	0.71	0.49	0.71

of interest, d, is that on the interaction between $Scale_t$ and government debt maturity. If, as predicted in Proposition 2, gap filling is stronger when $Scale_t$ is high, then we should find $d < 0$.

Note that this specification also allows for an interaction between a time trend and government debt maturity. We are thus asking whether there is an independent effect of $Scale_t$ on gap-filling behavior above and beyond the existence of a simple time trend in the intensity of gap filling. This relatively stringent test is motivated by an earlier observation from Table IV, namely, that gap filling appears to be more pronounced in the latter half of our sample period.

The results of these regressions are shown in Table VII. There are four columns, corresponding to the two measures of the size of the government bond market and to versions of (10) with and without the further controls y_{St} and $(y_{Lt} - y_{St})$. In each of the four cases, the key coefficient d is estimated to be negative, as predicted. The results are statistically significant in the first, third, and fourth columns, and marginally significant (t-statistic of 1.77) in the second. Thus, the evidence is generally supportive of Proposition 2.[17]

VI. Proposition 3: The Cross-section of Gap Filling

The model predicts that gap filling should be more pronounced among firms with lower costs of deviating from their target debt maturities. To test this proposition, we use the Compustat data to create disaggregated versions of the long-term corporate level share for various subsamples of firms. We can then ask whether this share responds more sensitively to the long-term government share among firms that appear to have more financial flexibility.

We use six proxies for financial flexibility. The first is a firm's market capitalization. The other five are motivated by the work of Kaplan and Zingales (1997), who show that the following firm-level characteristics are associated with a lessening of financial constraints: high dividends, high cash flow to assets, high cash balances to assets, low Tobin's Q, and low book leverage. For all of the variables except dividends, each year we assign those nonfinancial firms below the 30th percentile to the "low" category, and those above the 70th percentile to the "high" category. For dividends, we separate the payers and the nonpayers. Again, our predictions are that the coefficient on the long-term government share should be more strongly negative for firms that rank "high" in terms of market cap, cash flow, and cash balances; for firms that rank "low" in terms of Q and leverage; and for firms that are dividend payers.

Table VIII reports the results of these tests. But for the disaggregation, the baseline specification is identical to that in column 10 of Table III, including as additional controls y_{St}, $(y_{Lt} - y_{St})$, and a time trend. The first row of Table VIII just repeats the coefficient estimate on the government long-term share from the full Compustat nonfinancial sample: -0.228, with a t-statistic of 2.33.

In the second row, we see that the coefficient for large firms is -0.286, while that for small firms is 0.024; the t-statistic on the difference between these two coefficients is 2.18. These findings with respect to market cap echo the survey results of Graham and Harvey (2001): Managers of larger firms are more likely to say that they attempt to time movements in Treasury rates. Similarly, the third row shows that the coefficient for dividend payers is -0.263, while that for nonpayers is -0.043, with a t-statistic on the difference of 1.91. These two sample splits are illustrated in Panels A and B of Figure 3.

The fourth and fifth rows document that firms with high cash flows and cash balances also have more negative coefficients on the government long-term share, though only the former comparison is statistically significant (t-statistics of 1.94 and 1.07, respectively). The sixth row shows that low-Q firms have a coefficient of -0.318, while high-Q firms have a coefficient of -0.063, with a t-statistic on the difference of 1.97. Thus, for five characteristics—size, dividends, cash flow, cash balances, and Q—each of the subsample comparisons

[17] The full effect of an increase in the government long-term share is given by $b + d \cdot Scale_t + f \cdot time$, which explains why we do not focus on the fact that $b > 0$.

Table VIII
Disaggregated Results by Firm Type, 1963–2005

This table presents OLS regressions of the Compustat long-term level share on the long-term government share, disaggregated by firm type. Each year, nonfinancial firms are classified as low (below 30th percentile) or high (greater than 70th percentile) with respect to: market capitalization, cash flow over assets, cash balances over assets, Tobin's Q, and book leverage. Firms are also classified as either dividend payers or nonpayers. These separate maturity measures for low and high firms are regressed on the long-term government share. In the first two columns of each row, we report the slope coefficient on the long-term government share, b, and its associated t-statistic from the separate low firm and high firm regressions. In the final column, we regress the difference between the level shares for the high and low groups on the government level share. The t-statistic from this regression tests the equality of coefficients between the high and low groups. All regressions include a constant term and controls for the short-term rate, the term spread, and a time trend. t-statistics are based on Newey-West (1987) standard errors allowing for 2 years of lags.

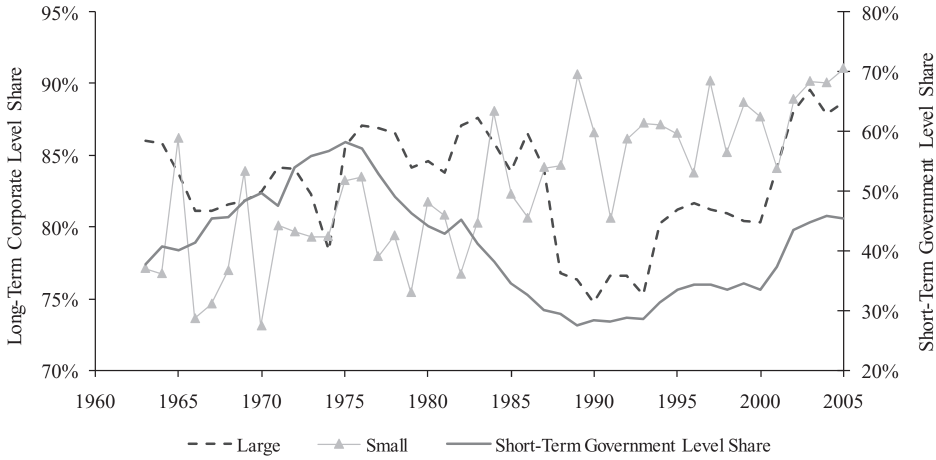
	Low		High		High − Low	
	b	$[t]$	b	$[t]$	$b^{\text{High}} - b^{\text{Low}}$	$[t]$
All Compustat nonfinancial	−0.228	[−2.33]				
Market capitalization	0.024	[0.43]	−0.286	[−2.50]	−0.310	[−2.18]
Nonpayers ("low"); payers ("high")	−0.043	[−0.83]	−0.263	[−2.30]	−0.220	[−1.91]
Cash flow/assets	0.073	[1.35]	−0.125	[−1.42]	−0.198	[−1.94]
Cash/assets	−0.059	[−0.39]	−0.215	[−2.53]	−0.156	[−1.07]
Tobin's Q	−0.318	[−3.09]	−0.063	[−0.69]	0.255	[1.97]
Leverage	−0.375	[−3.19]	−0.367	[−2.88]	0.008	[0.06]

goes in the direction predicted by the theory, albeit not significantly in the case of cash balances.

The one sample split that yields no meaningful differential is book leverage: The coefficients for high- and low-leverage firms are almost the same, at −0.367 and −0.375, respectively.[18] Nevertheless, the overall picture that emerges from Table VIII is that, according to most measures, it does appear that increased flexibility is associated with more aggressive gap filling. Thus, the evidence is largely, though not entirely, consistent with Proposition 3. While we interpret our results as reflecting differences in debt supply elasticities across firms, they would also be consistent with investors viewing the debt of corporations with strong balance sheets as a closer substitute for government debt than the debt of financially constrained firms.

[18] One possible explanation for this nonresult is that, by definition, high-leverage firms enjoy greater dollar benefits from timing the debt market. Hence, if there are any fixed costs associated with having an activist debt management policy, high-leverage firms will be more inclined to bear this fixed cost and thus to engage in gap filling. This would create an effect that runs counter to the financial flexibility effect envisioned in our model.

Panel A. Small and Large Firms



Panel B. Payers and Nonpayers

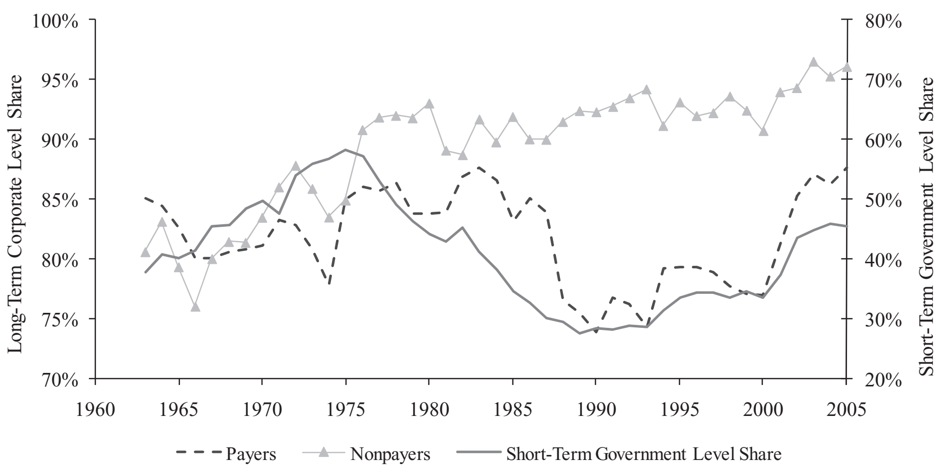


Figure 3. Long-term debt share, Compustat splits, 1963–2005. The solid line, plotted on the right axis, is the share of government debt with maturity of 1 year or less. In Panel A, the dashed and hatched lines plot the long-term corporate share for large capitalization and small capitalization firms, respectively. In Panel B, the dashed and hatched lines plot the long-term corporate share for dividend payers and nonpayers, respectively.

VII. Proposition 4: Gap Filling and Excess Bond Returns

Our final analysis, in Table IX, examines return predictability in the Treasury bond market. Here, we use a longer sample period of 1953 to 2005 to allow for comparison with BGW (2003), Butler et al. (2006), and Greenwood and Vayanos (2008). There are three blocks in the table, corresponding to future excess returns over 1-, 2-, and 3-year horizons. The first column in each block reproduces the baseline findings of Greenwood and Vayanos (2008), using the

Table IX

Corporate Debt Maturity, Government Debt Maturity, and Excess Bond Returns, 1953–2005

Annual regressions of future log excess bond returns over 1-, 2-, and 3-year horizons on combinations of the long-term government share, the *Flow of Funds* (FOF) long-term corporate level share, and the FOF long-term corporate issue share. In the 1-year regressions, the t-statistics, in brackets, are based on heteroskedasticity-robust standard errors. In the 2- and 3-year regressions, t-statistics are computed using Newey-West (1987) standard errors allowing for autocorrelation at up to 1 and 2 years, respectively.

	One-Year Excess Returns (%)					Two-Year Excess Returns (%)					Three-Year Excess Returns (%)				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)
$D^G_{L,t}/D^G_t$	0.225		0.185		0.170	0.523		0.394		0.388	0.824		0.580		0.576
	[1.57]		[1.09]		[1.10]	[2.59]		[1.60]		[1.71]	[3.06]		[1.82]		[1.95]
$d^C_{L,t}/d^C_t$		−0.337	−0.163				−0.912	−0.542				−1.588	−1.045		
		[−1.00]	[−0.42]				[−2.36]	[−1.16]				[−2.74]	[−1.60]		
$D^C_{L,t}/D^C_t$				−0.312	−0.206				−0.778	−0.531				−1.408	−1.034
				[−1.41]	[−0.84]				[−2.28]	[−1.33]				[−3.08]	[−1.99]
Constant	−12.583	8.190	−6.630	20.280	3.565	−29.532	21.591	−9.876	50.063	11.695	−46.773	37.122	−9.173	89.926	32.510
	[−1.52]	[1.03]	[−0.40]	[1.40]	[0.17]	[−2.31]	[2.54]	[−0.42]	[2.34]	[0.34]	[−2.71]	[3.16]	[−0.30]	[3.26]	[0.73]
R^2	0.04	0.02	0.05	0.03	0.05	0.12	0.09	0.15	0.11	0.17	0.19	0.17	0.25	0.20	0.28

long-term government share to forecast returns. When government maturity is high, subsequent returns on long-term bonds are high as well—hence, the motive for firms to shift toward short-term debt. The magnitudes are economically interesting: When the government share goes up by one percentage point, excess bond returns rise by 22.5, 52.3, and 82.4 basis points at the 1-, 2-, and 3-year horizons, respectively.

The second and fourth columns of each block present univariate regressions similar to those in BGW (2003). The long-term corporate level share and the long-term corporate issue share (both based on *Flow of Funds* data) are used one at a time to forecast excess returns. Both variables have significant predictive power at the 2- and 3-year horizons, though with the opposite sign as the government long-term share. It should be noted that while the qualitative picture is similar to that in BGW, the statistical significance of our results is somewhat weaker than that reported by BGW for the 1953–2000 period; this divergence is caused by the year 2001, when both corporate debt maturity and excess bond returns were high.

The above results are not new. However, our theory does make the following novel prediction, embodied in Proposition 4: To the extent that corporate debt maturity predicts bond returns, some of this predictability arises simply because corporate debt maturity serves as a mirror of government debt maturity, and hence of the supply shocks that are the ultimate driver of returns. Thus, once government maturity is included in the regression, the predictive power of corporate maturity—measured in either levels or in issues—should be diminished. These bivariate horse races are shown in the third and fifth columns of each block. And, as can be seen, they provide consistent support for this aspect of our theory. Consider, for example, the case where the long-term corporate issue share is used to forecast future excess returns over a 3-year horizon. When used as a univariate predictor, this variable attracts a coefficient of -1.588, with a t-statistic of 2.74. However, when it is entered with the long-term government share, the coefficient falls to -1.045, with a t-statistic of -1.60—that is, it shrinks by about one-third of its original value.

VIII. Conclusions

The survey evidence in Graham and Harvey (2001) suggests that at least some of the time-series variation in corporate debt maturity reflects an active effort by managers to time the debt market, that is, to issue at the cheapest point on the yield curve. Such attempts at market timing are difficult to understand if one thinks in terms of access to information or forecasting capabilities: It is hard to see why the managers of nonfinancial firms should have any advantage—relative to say, hedge fund managers—at predicting future bond market returns.

This paper argues that debt market timing by firms makes more sense when viewed through the lens of liquidity provision. Even if operating firms have access to the same information as hedge funds, and hence make the same forecasts of excess returns, they do bring to the table significant additional risk absorption capacity. This extra capacity is of particular value when movements in excess returns are driven by quantitatively large and undiversifiable supply shocks, as is the case in the Treasury bond market.

A similar logic can be used to think about other forms of market timing. For example, it has been documented that firms exhibit timing behavior with respect to both the firm-specific and aggregate components of stock prices.[19] While a theory based on private information may shed light on how individual firms manage to issue equity in advance of low idiosyncratic returns, this approach is less well suited to explaining why high values of aggregate issuance forecast low market-wide returns, as in Baker and Wurgler (2000). We suspect that here, too, thinking about firms as macro liquidity providers is likely to be fruitful. A clean illustration of this point comes from the crash of October 1987. In the wake of the crash, many firms announced repurchase programs. Given that the crash itself was common knowledge, it is hard to believe that it created a private advantage in forecasting ability. However, given the stresses on arbitrage capital caused by the crash, it seems likely that operating firms, especially those with strong balance sheets, were advantaged in terms of risk absorption capacity.

The hypothesis that firms behave as activist macro arbitrageurs may strike many as being far from the dictates of textbook corporate finance theory, which is often interpreted as saying that, absent adjustment costs, firms should stick close to an optimally chosen target capital structure. However, it should be emphasized that our theory is based on the single most fundamental concept in corporate finance, namely, the Modigliani and Miller (1958) irrelevance proposition. To the extent that Modigliani–Miller provides an accurate description of reality—that is, to the extent that firms are otherwise approximately indifferent to variations in capital structure in the neighborhood of their target optima—their comparative advantage over other capital market players in the realm of macro arbitrage is all the more pronounced.

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[19] See, e.g., Loughran and Ritter (1995) and Ikenberry, Lakonishok, and Vermaelen (1995) for evidence at the firm level, and Baker and Wurgler (2000) and Lamont and Stein (2006) for evidence at the market level.

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