Tests of a Signaling Hypothesis: The Choice between Fixed- and Adjustable-Rate Debt

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We develop a model wherein the choice between adjustable- and fixed-rate debt can serve as a signal of firm quality. The nature of the signal depends on expected inflation volatility relative to other risk parameters. Evidence from a matched sample of debt announcements over the period 1978 to 1986 shows a difference of -2.05 percent between stock price reactions to adjustable rate and fixed rate announcements when expected inflation volatility is above an estimated threshold Below this threshold, the difference is +0.98 percent. The evidence supports the hypothesis that the riskier debt choice serves as a favorable signal of firm quality.

The early 1980s witnessed the emergence of an active adjustable-rate corporate debt market in the United States that paralleled the acceleration of interest rate and inflation volatility. As a proportion of the total face value of newly registered issues, adjustable-rate debt rose from 1 percent in 1978 to a peak in 1983 of about 13 percent. Adjustables, however, quickly returned to obscurity, representing less than 4 percent of the market for new issues in 1986. By 1986, the high levels of interest rate and inflation volatility had also subsided.

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In this article, we use the growth and decline of the adjustable-rate debt market to motivate an investigation into how and whether corporate choices of debt type reveal private information about firm quality. We develop a model where costs of financial distress give managers an incentive to choose borrowing strategies that stabilize net income. The model also provides managers incentives to adopt high-risk strategies when gains from signaling favorable inside information about firm quality outweigh increases in distress costs. A signaling outcome exists if taking on risk in net income brings about a smaller increase in costs of distress for highquality firms than for their low-quality counterparts.

What is unusual about the possibility of signaling through the choice between fixed- and adjustable-rate debt is that either contract can be the more risky and, therefore, the more favorable signal depending upon the relative levels of expected real rate and inflation volatility. To clarify how, suppose the correlation between operating income (before interest expense) and the level of market interest rates is caused primarily by the correlation between operating income and inflation. The tendency for interest payments on adjustable-rate debt to track expected inflation helps hedge uncertainty in operating income because changes in expected inflation will be a component of the total inflation realized over multiple periods. Thus, where the volatility of expected inflation is high, adjustable-rate debt stabilizes net income (after interest expense) relative to the' impact on net income of a fixed-rate debt contract. On the other hand, changes in real market interest rates also cause changes in the interest paid on adjustable-rate debt. Where real volatility dominates expected inflation volatility, adjustable-rate debt destabilizes net income.

The data suggest that changes in expected inflation volatility were the primary cause of changes in interest rate volatility over the period in which the adjustable rate debt market gained prominence. Holding constant the other risk parameters in the model, we characterize a threshold in expected inflation volatility above which adjustable rate debt is the low-risk choice and below which it is the high-risk choice, We then test the hypothesis that, when signaling is taking place, adjustable-rate financing conveys bad news if expected inflation volatility is above the threshold. Below the threshold, adjustable-rate financing conveys good news.

Although we explore several different empirical specifications, the simplest is built around a constant threshold in expected inflation volatility at which there is a shift in the mean of the difference in stock reaction between the two types of announcements. Within the model, we estimate the average difference conditional upon being above or below an estimated threshold,

r_{ft} (

r_{ft} (

[1] Thakor (1991) uses Ross' framework in his review of game theoretical models in corporate finance.

where i_t represents the continuously compounded, one-period real interest rate between t and $t+1$, π_t is the continuously compounded inflation rate between t and $t+1$, and E_t is the expectation operator at date t. In addition, it *simplifies* the model to assume that *inflation* follows a random walk-through time, with possibly time-varying volatility.[2]

Assumption 2. *Investing: There is a firm at t = 0 having access to a positive NPV project that requires $1 of external financing. The project is indivisible and foregone if not undertaken at t = 0. Combined with &sets in place, the project generates a single liquidating nominal*

$$p^A = \exp\{r_{f0} + r_{f1} + \Delta^A\},$$

$$\Delta^A,$$

$$\Delta^F \geq 0$$

termined by the levels of the other three risk parameters identified in Proposition 1. Values of expected inflation volatility above the threshold make adjustable-rate debt the low default risk strategy, while values below the threshold make fixed-rate debt the low default risk, strategy.

Proof. See Appendix A. ■

The condition underlying Proposition 2 that fixed-rate debt has lower default risk and is, therefore, the low-cost strategy when inflation is nonstochastic, seems consistent with the evidence that very little long-term adjustable rate debt is sold during periods of stable prices. With this caveat, Proposition 2 states that a unique positive threshold separates two distinct regimes: a regime of high expected inflation risk where default is hedged by borrowing at an adjustable rate, and a regime of low expected inflation risk where default is hedged by borrowing at a fixed rate.

As we show later, the volatility of the expected real rate has small time-series variation relative to the variation in expected inflation volatility. In addition, we show that there is very high correlation between the volatility of expected inflation and the 'volatility of the nominal interest rate. Thus movements in the volatility of expected inflation seem to dominate the potential influence of any covariance between the expected real rate and expected inflation. Therefore, we conclude that it is appropriate to think of the threshold for a given firm as approximately constant over time and focus on the level of expected inflation volatility as the determinant of whether or not the firm is above or below its thresho1d.[9]

In general, thresholds are firm specific because they depend on the covariance between the expected real interest rate and the log of the real cash flows to the firm. A high covariance implies a low threshold Initially our empirical model assumes the same threshold for all firms, treating cross-sectional variation as a second-order effect. We then introduce a more general empirical specification.

1.3 Equilibrium analysis and testable implications

Appendix B contains an analysis of equilibrium borrowing strategies. The main conclusions are summarized in Proposition *3.*

Proposition 3. *Borrowing strategies have three possible sequential equilibria: a separating equilibrium wherein firms above a minimum*

[9] For a given covariance between the expected real rate and the log of real cash flow, sufficient conditions for the relative hedging quality to depend only on the level of expected inflation volatility are that both the volatility of the expected real rate and the correlation between the expected real rate and expected inflation are constant or at least nonstochastic functions of expected inflation volatility.

Table 1
Corporate debt and treasury bill information over the period 1978-1986

Year	Fixed rate	Adjustable rate	Proportion of adjustable	Avg. yield on 1-month T-bill	Volatility of $E_t(\pi_{t+1})$	Avg. monthly inflation
1978	19,466	200	1.0%	6.92%	0.61	8.64%
1979	23,093	803	3.4%	9.69%	0.70	12.51%
1980	63,230	669	1.0%	10.44%	2.17	11.71%
1981	39,017	676	1.7%	13.66%	2.01	8.57%
1982	90,425	7,889	8.0%	9.96%	1.81	3.77%
1983	61,556	9,158	13.0%	8.28%	1.16	3.72%
1984	105,344	12,440	10.6%	8.94%	0.91	3.97%
1985	142,483	7,635	5.1%	7.15%	0.99	3.73%
1986	236,167	9,233	3.8%	5.63%	0.82	1.19%
Sum	780,781	48,703	5.9%			
Avg.				8.96%	1.24	6.42%

Amounts of fixed- and adjustable-rate debt are in $million of principal as reported in the Registered Offering Statistics of the SEC. T-bill information taken from CRSP. Volatility is of the unanticipated change in expected inflation for the next month, $E_t(\pi_{t+1})$, stated as standard deviation based on an ARCH(12) variance process. Inflation is the average monthly change in seasonally adjusted CPI stated in percent per annum

Table 2
SIC breakdown of debt issues and selected summary statistics for the matched and raw bond samples

Statistic	Matched sample (adjustable rate, fixed rate)	Raw sample (adjustable rate, fixed rate)
No. observations	(47, 47)	(67, 179)
Average maturity (years)	(11, 13)	(11, 16)
Average principal (millions)	(128, 111)	(136, 100)
Average asset size of the issuing firm (billions)	(24.8, 16.0)	(26.9, 4.5)
S&P ratings		
75th percentile	(BBB,BBB)	(BBB,BB)
Median	(A,A)	(A,A)
25th percentile	(AA,AA)	(AA,AA)
Range	(B-AAA,B-AAA)	(B-AAA,CCC-AAA)
SIC classification		
Extractive	(1, 1)	(2, 12)
Manufacturing	(17, 17)	(21, 88)
Transportation	(0, 0)	(1, 3)
Communication	(1, 1)	(1, 2)
Electric, gas, water	(2, 2)	(4, 19)
Thrifts (depository) institutions)	(18, 18)	(28, 9)
Financial (excluding banks, bank holding cos., and thrifts)	(3, 3)	(5, 18)
Commercial and other	(5, 5)	(5, 28)

We selected from the ROS tape an initial sample of 250 fixed-rate issues, stratified across the years 1978 through 1986. Employing the same screens and search procedures resulted in a final sample of 179 fixed-rate announcements. The fixed- and adjustable-rate observations are combined into a raw sample of 246. The primary emphasis in our empirical test, however, revolves around a sample of matched pairs

in which each adjustable-rate debt announcement is matched with a suitable fixed-rate announcement and the stock price reactions differenced. A total of 47 pairs were obtained by employing the following matching criteria on the entire fixed-rate debt population of the ROS tape:

(i) Timing: a clean fixed-rate debt issue announcement' had to be found within a 6-month period centered around the adjustable-rate announcement.
(ii) Industry classification: the fixed-rate announcement had to be made by a firm in the same three-digit SIC industry.
(ii) Maturity: whenever more than one match satisfying (i) and (ii) could be found, the one with the maturity closest to that of the adjustable-rate issue was selected.

Of the original 67 adjustable-rate observations, 20 were lost because no fixed-rate observation could be found that satisfied (i) and (ii) simultaneously.. We did not screen on whether the fixed-rate issue contained a call provision since most fixed-rate debt issued over this period was callable.[10]

Table 2 reports selected summary statistics on the matched and raw, unmatched samples. As expected, differences in average maturity, average principal, and average asset size between the fixed- and adjustable-rate observations are smaller in the matched sample. Both samples show similar ratings for fixed- and adjustable-rate debt. Within the SIC classifications represented, the most noteworthy feature is the relatively large number of depository institutions issuing adjustable-rate debt. This reflects the acknowledged importance of banks in the adjustable rate note market. The lack of industry control in the raw sample thus leads to a larger asymmetry in the relative proportions of banks and manufacturing firms among the adjustable- and fixed-rate observations in that sample. We break out the banking subsample during our discussion of the matched sample results.

The motivation for placing emphasis on results from the matched sample is threefold. First, return differencing provides a control for the average unconditional announcement effect of debt issues and helps eliminate potential dependencies in announcement returns that may exist for reasons outside our model. Second, matching provides a screen against pooling equilibria. Recall that the matched sample contains only adjustable-rate observations for which there exist fixed-rate matches in the same industry issued at the same time

[10] Of the 47 fixed-rate issues used in the matched sample, 40 contain call provisions. For these 40, the average the to first call **is** 5.9 years with an average call premium of 2.9 percent over face value.

and thus drawn from the same volatility environment. Similarly, the matched sample only contains fixed-rate observations with adjustable-rate matches. Thus, to the extent that similar firms facing the same market environment at the same point in time also face the same equilibrium, matched observations must come from separating equilibria. Finally, since matched pairs control for calendar time and thus volatility, matching provides a natural solution to the following prior anticipation problem.

In measuring the average impact of a signal, we would like to estimate the expected difference between the market value of a firm that gives one signal and the market value of a firm that gives the opposite signal under the same market conditions. Using the vernacular of Acharya (1988) and Lanen and Thompson (1988), the difference in average reactions within raw or unmatched samples of firms announcing adjustable- and fixed-rate debt will tend to be an attenuated estimate of the average impact of the signal. To see why, consider a signaling equilibrium within a range of high volatility where adjustable-rate debt is the low-risk alternative. With a continuum of firm types, the average quality and overall percentage of firms choosing adjustable-rate debt will tend to increase with volatility. Raw sampling of adjustable-rate announcements, by tending to overrepresent periods of high volatility, attenuates toward zero the average announcement effect relative to an average taken evenly across the spectrum of volatilities. A similar attenuation occurs within a raw sample of fixed-rate announcements because periods of low volatility will tend to be overrepresented.

On the other hand, by pairing adjustable- and fixed-rate announcements from the same volatility, the difference in announcement effects is an unbiased estimate of the impact of the difference in signals. Matching controls for the prior probability of a signal in much the same way as the estimation of the prior probability function described by Acharya (1988). It should be emphasized, however, that attenuation caused by an incomplete control for differences in prior probability serves only to reduce the average difference in announcement effects without altering the sign.

3. Empirical Model

3.1 Preliminaries

Our model postulates a relationship between the level of volatility of expected inflation and the difference in stock price reaction to the announcement of adjustable- and fixed-rate debt. To give a sense for the functional form, consider initially the following simplifications:

(i) the volatility. of expected real rates and the correlation between the expected rate and expected inflation are constant,

(ii) the covariance between real cash flow and the expected real interest rate is the same for all firms, and
(iii) investors attach a uniform prior distribution of firm quality over the range *B* to G for all firms, with quality measured in present value.

Assumptions (i) and (ii) ensure that every firm faces the same time-independent threshold in expected inflation risk above which a fixed-rate announcement conveys good news relative to an adjustable rate announcement and below which the converse is true. Assumption (iii) implies that the difference in stock price reaction will be a simple step function in expected inflation risk. The logic of this last implication is as follows.

At volatility levels above the threshold, each manager having a debt offering must decide whether the benefit of issuing fixed-rate debt and pooling with higherquality firms compensates for the extra default cost. There will be a function; say Q(b), relating the true quality of the marginal firm (Q) to volatility *(h)* and this function is known by the market. A firm announcing adjustable-rate debt at volatility level h has a stock price change equal to *(B +-Q(h))/2 - (B + G)/2* because the announcement reveals that firm quality is between *B* and . . Q(h). Since Q(b) lies between *B* and G, this change is negative. Similarly, a firm announcing fixed-rate debt at volatility *h* has a stock price change equal to *(G+Q(h))/2- (B+ G))/2,* which is positive. The difference in stock price changes between firms announcing adjustable- and fixed-rate debt is -(G - *B)/2,* a negative constant independent of the function *Q(b)* and hence independent of volatility.[11] Parallel reasoning implies that the same relation exists below the threshold, but with the opposite sign. Stated as a rate of return on the initial stock price, we could define the (percentage) stock price reaction as *-(G - B)/(B + G).* For the difference in stock price reactions, we are led to the implication that the function has two values, positive below the threshold and negative above it. Extensions of this model are examined in Section *3.5.*

For the matched sample, the empirical model attempts to identify the parameters of a step function in (abnormal) rates. of return. Let R_j represent the difference in the stock price reaction to the announce-

[11] While the prior distribution over quality type affects the functional form of the differential reaction, the uniform case, which implies a constant differential, captures the essence of the process. The uniform case is attractive because the function Q(h) drops out of the differential stork reaction, avoiding the need to characterize Q(h). With a nonuniform prior, the magnitude **of the** differential stock reaction will depend on Q(h) and could be quite complex. To compute the differential stock response, Q(b) would have to be derived explicitly, which would require further parameterization of the model (e.g., the relation between variances and probabilities of default). We show in our empirical work that a constant differential conforms well to the data.

ment of adjustable-rate debt for firm j and the stock price reaction to the announcement of fixed rate debt for a suitable control firm. To simplify the exposition call R_j the abnormal return to firm j's announcement. The model predicts

$$\begin{aligned} E(R_j) &< 0 \quad b_j > b^* \\ E(R_j) &> 0 \quad b_j < b^* \end{aligned} \tag{4}$$

where E is the expectations operator, b_j is the volatility in expected inflation existing at the time of announcement for firm j, and h^* is the threshold level of volatility above which a fixed-rate issue is a favorable signal and below which adjustable-rate debt is a favorable signal.

We proceed by positing a switch in means for the abnormal return process and simultaneously estimating the switch point, h^*, and the two means. Let

$$\begin{aligned} R_j &= \alpha_1 + e_j \quad b_j < b^* \\ R_j &= \alpha_2 + e_j \quad b_j > b^* \end{aligned} \tag{5}$$

where R_j, b_j, and h^* are as defined in Equation (4), α_1 and α_2 are the mean abnormal returns in the two regions of h, and e_j is a mean zero error term. The null hypothesis that there is no signaling region implies that $\alpha_1 = \alpha_2 = 0$. The model with signaling in both regions predicts $\alpha_1 > 0$, $\alpha_2 < 0$. Indeed, in the uniform case where all observations come from signaling regions and both regions exist in the data, the model 'implies that the two means are of equal magnitude but **of** opposite sign. If, however, good firms are able to signal with adjustable rate debt throughout the range of volatility in the sample, it is possible that no threshold will be detected and both sample means will be positive.[12]

Empirical work for the raw sample proceeds along similar lines except the abnormal return for each announcement is defined as the difference between the stock price reaction to the announcement and the mean stock price reaction for the entire sample. This provides a crude control for the unconditional effect of a debt issue announcement. Where the data come from a signaling region, the model predicts an abnormal return for fixed-rate announcements that has the opposite sign as the abnormal return for adjustable-rate announcements. Combining observations, we reverse the sign of the fixed-rate abnormal returns. The pattern predicted by the model for the raw

[12] Because many observations come from low-volatility periods where fixed rate is the predominant choice, we view the hypothesis that adjustable rate is the low-cost debt choice (and as a result, fixed rate is the favorable signal) over the entire range of data as implausible. A negative average difference throughout the range of observations would seen inconsistent with the model.

sample is that the combined observations should tend to be negative at high levels of volatility and positive at low levels.

3.2 Abnormal return calculations

The stock price reaction to a debt announcement is estimated for every firm' in the sample using standard event study methodology. The stock price reaction for firm j on day t, AR_{jt}, is defined as

$$AR_{jt} = r_{jt} - a_j - b_j r_{mt} \tag{6}$$

where r_{jt} is the return on the common stock of firm j on day t, r_{mt} is the return for the CRSP equal-weighted index on day t, and a_j and b_j are the ordinary least squares coefficients of the market model estimated over the period [-140, -21] preceding announcement. Missing observations and those directly following them are omitted prior to estimation. The stock price reaction to the announcement is defined as the sum of the reactions on the announcement day (day 0) and the day immediately preceding the announcement day (day -1).

The average stock price reaction across the raw sample of 246 fixed- and adjustable-rate announcements is -0.47 percent. The fixed rate announcements have an average stock price reaction of -0.42 percent while the adjustable-rate announcements have an average of -0.61 percent. The average for the fixed-rate sample is consistent with the results reported by Mikkelson and Partch (1986) and Eckbo (1986), who document insignificant or marginally significant negative average stock price reactions to nonconvertible fixed rate announcements.

For the matched sample, the abnormal stock return for each adjustable-rate announcement is the difference between the stock price reactions to the adjustable announcement and the fixed-rate control. The average abnormal return for the matched sample of -0.97 percent is larger in magnitude than the difference 'in means between the adjustable and fixed-rate subsets of the raw sample. The increase in magnitude presumably reflects the fact that matching helps eliminate pooling observations and helps control for prior anticipation. A plot of the abnormal returns against time is shown in Figure 1 for the matched sample. The figure shows a discernable U-shaped pattern with more negative abnormal returns bunched in the middle of the time period.

3.3 Inflation volatility and other macroeconomic variables

To provide a thorough investigation into the announcement effects of fixed versus adjustable-rate debt, we estimate volatility several different ways and incorporate other macroeconomic variables that might

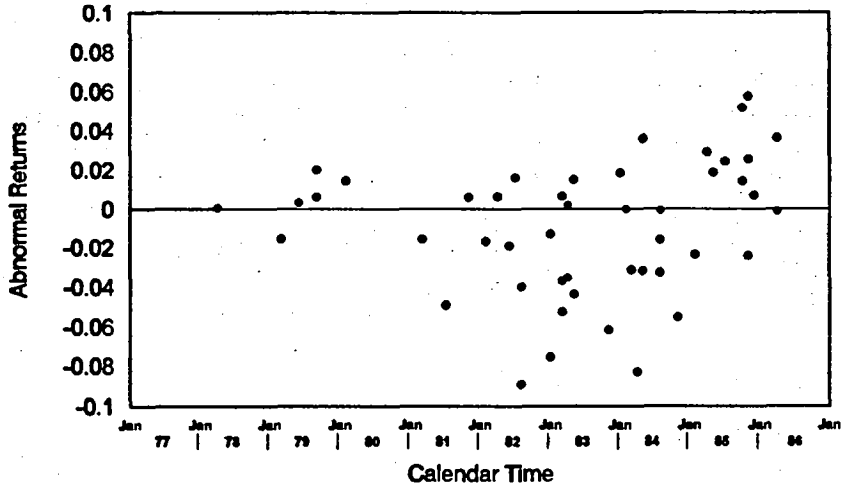

Figure 1
Plot of abnormal returns against time for a matched sample of 47 adjustable rate debt announcements over the period 1978-1986
Abnormal return is the 2-day stock price reaction to an adjustable rate debt announcement minus the reaction to a control firm announcing a fixed rate issue. The plot indicates a slight U-shaped pattern but reveals little correlation between abnormal return and calendar time.

influence the debt decision of corporations. Our investigation works with four macroeconomic series. Data on short-term Treasury bills are taken from the Fama files in the CRSP tape over the period 1970 through 1989. Monthly rates of inflation are defined as the percentage change in the CPI, where the level of the CPI is seasonally adjusted as reported by DRI. The term structure slope is defined as the difference between the 5- and l-year spot rates on Treasury securities as provided in the Fama and Bliss files on CRSP. Monthly default spreads are defined as the difference in yield on BBB and AAA bonds as reported in Standard and Poor's.

Most adjustable-rate debt contracts are written on the 3-month Treasury bill yield or 3-month LIBOR rate, suggesting a short-term perspective for volatility. We build our volatility measures by adopting the naive interest rate model described in Fama and Gibbons (1984, Section 2.3), which captures the essence of a full ARIMA specification of real rate and inflation expectation formation. First, a time-varying measure of the expected real rate of interest is calculated by averaging the realized real rate over the previous 12 months. Second, the expected rate of inflation for a given month is defined as the difference between the l-month T-bill rate at the beginning of the month

and the expected real rate for the month. We create a time series of expected real interest rates and expected inflation by applying these procedures over the 20-year period from 1970 through 1989. Finally, time-varying volatilities of the expected inflation and the real rate are estimated for the series using an ARCH methodology.[13]

The expected inflation series is close to a random walk because it is constructed as the difference between a T-bill rate and a slow-moving average of past realized real rates. Although a simple characterization, let the first differences in expected inflation be white noise with time-varying volatility conforming to an ARCH(12) as adopted by Engle, Lilien, and Robins (1987) in their study of monthly T-bill return premia on two month bills[14]

$$\begin{aligned} y_t &= y_{t-1} + \varepsilon_t \\ h_t &= \gamma_0 + \gamma_1 \sum_{j=1}^{12} w_j \varepsilon_{t-j}^2 \\ w_j &= (12 + 1 - j)/78 \\ \varepsilon_t &\sim N(0, h_t) \end{aligned} \tag{7}$$

where y_t is expected inflation in month t, h_t is the conditional variance of the change in expected inflation from t - 1 to t, and ε_t is the change in expected inflation from t - 1 to t. The parameters of the process, γ_0, and γ_1, are estimated over the entire 20-year period with iterative scoring as suggested in Engle (1982). Results do not appear sensitive to the choice of starting values.

The point estimates of γ_0, γ_1 are 0.169 and 0.972,[15] indicating, consistent with previous research on interest rates, a high autoregressive component to the variance series. From these estimates and the realizations of u, a volatility measure in each month is calculated from Equation (7). The relevant measure of volatility for the debt announcements in a given month t pertains to the expected inflation for month

[13] Within our expected inflation specification. longer-term measures of volatility are monotonic in the current short-term measure. As an exercise, we developed a 3-month expected inflation forecast based on the naive interest rate model as described in the text. The correlation between the expected volatility of the sum of the next 3 months of expected inflation and the volatility of next month's expected inflation exceeds 0.95 over our time period.

[14] Gray (1994) provides a more complete structure for short-term interest t-ate dynamics. We remove any first-order autocorrelation In the **first** differences of the expected inflation series before applying the ARCH methodology.

[15] The corresponding *t -ratios* for these estimates are 4.36 and 15.10, suggesting significant explanatory power for the model. Standard errors are asymptotic as discussed in Engle (1982). Expected inflation is stated as an annual percentage rate (i.e., the monthly rate times 1200).

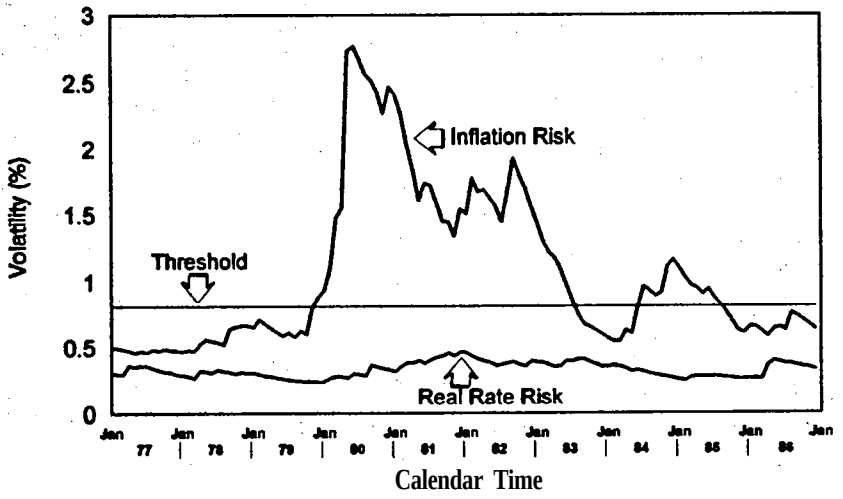

Figure 2
Volatilites of expected inflation and the expected real rate of interest based on ARCH(12) models
The expected real rate and expected inflation series are constructed **the** the naive interest rate model described in Fama and Gibbons (1984) and Section 3.3 in the text. Volatilities are standard deviations of annualized percentage rates. The plot shows high volatility of expected inflation during several time periods and generally more constant volatility for the expected real rate. The threshold indicated in the figure is the estimated switch point in inflation volatility above which fixed rate debt is the favorable signal of firm quality. Below the threshold, adjustable rate debt appears to be the more favorable signal.

$t + 1$, and thus h_{t+1} is used as the volatility measure for announcements in month t.[16] The volatility measures are plotted over the time period covered in our study (1977-1986) in Figure 2. The volatilities are represented by their square roots (i.e., the standard deviation) based on monthly rates stated in percent per year.

As can be seen, expected inflation volatility is high in the middle of the time period from the third quarter in 1979 to the second quarter of 1983. A second plot shows the estimated volatility of expected real rates. This series is estimated using the same procedures as shown in Equation (7) with y_t equal to the expected real rate. The estimated volatility of expected real rates is a much smoother series, supporting our assumption that real rate risk is essentially constant over the period of interest. The figure also contains a reference line that identifies the

[16] **Where the fixed-rate announcement for the control firm occurs in a different month from the adjustable-rate announcement, the volatility estimates are averaged to arrive at the volatility measure used for that pair of firms.**

Table 3
Correlation matrix of macroeconomic variables over the period 1977 through 1986

	Volatility of exp. inflation	Volatility of T-bill yield (1 mo.)	Term str. slope	Default spread	T-bill yield (1 yr)
Volatility of exp. inflation	1.881				
Volatility of T-bill yield (1 mo.)	0.945	1.754			
Term structure slope	−0.132	−0.134	1.57		
Default spread	0.34	0.346	0.006	0.336	
T-bill yield (1 yr)	0.504	0.462	−0.586	0.390	2.758

Volatilities are measured with ARCH(12) on annualized rates. Expected inflation is as defined in the text. Volatility of T-bill yield is for 1-month T-bills. Term slope is 5-year spot rate minus 1-year T-bill rate. Default spread is BBB yield minus AAA yield. All rates are stated as annual percentages. Diagonal elements are standard deviations.

threshold level of inflation volatility estimated and discussed in the next section.

As a final check on the data, volatility measures for 1-month T-bill returns are calculated using the same stochastic process as described in Equation (7) with y_t equal to the 1-month T-bill rate for month t. This series is highly correlated with the expected inflation volatility series. Table 3 shows a correlation matrix of the macroeconomic variables used in this study where the correlation between expected inflation volatility and T-bill rate volatility is 0.945 over the 10-year period 1977 through 1986. Table 3 also shows that the T-bill yield is positively correlated with both inflation volatility and nominal interest rate volatility as expected. It is also positively correlated with the default spread and negatively correlated with the term structure slope. The default spread is positively correlated with both volatilities. The correlations between the term structure slope and the volatilities, while negative, are not significant at conventional levels.[17]

3.4 Estimating the volatility threshold and the abnormal return means

The estimation of models having discrete parameter variation has been studied by Goldfeld and Quandt (1972), Quandt (1960), and Hinkley

[17] A standard test for the significance of a correlation coefficient is to compute the ratio $t = \rho[(N-2)/(1-\rho^2)]^{1/2}$ where N is the number of observations and ρ is the correlation coefficient, and t is approximately unit normal under the null hypothesis. With 120 observations, the 5 percent level is about 0.2 in magnitude.

(1970).[18] Our initial approach follows most closely that of Hinkley (1970). He considers a discrete time process with independent observations in which a change in means occurs at an unknown point in the process. For the case where the error terms are normally distributed, Hinkley (1970) shows that. the method of maximum likelihood involves finding the point in time where dividing the sample and estimating both means generates the smallest sum of squared errors. Although the model is nonlinear, the mean in each regime typically conforms well to its asymptotic distribution, with variance equal to the variance of the residuals from the fitted model divided by the number of observations in the regime.[19]

In our application, the switch in regimes is determined by expected inflation volatility and not a time subscript. Our first step is to rank the observations on volatility and find the switch that produces the minimum residual variance from the fitted model with two means.[20,21]

Several results are shown in Table 4. The first panel shows the average abnormal return for the complete matched sample and for the banking subsample. For the complete sample, the average attains a value of –0.0095 with a *t*- statistic of -1.92. The banking subsample has an average abnormal return of -0.019 with a t-statistic of -2.24. This panel provides marginally significant evidence of a negative overall difference in stock price reaction to adjustable-rate debt announcements. The negative impact is over twice as large for the banking subsample, but the difference in means between the banking subsample and the rest of the sample is not significant at conventional levels. The difference is -0.0157 with a t-ratio of -1.56 based on the

[18] Kon and Jen (1978) and Sanders and Unal (1988) use Goldfeld and Quandt (1972) switching regression to estimate regimes for, respectively, market model beta and short-term interest rate volatility. Gray (1994) formulates short-term interest rate dynamics where the probability that the process is in a specific regime at time t depends on the regime at time $t-1$. In our model, the regime is determined solely by the level of volatility at time t, simplifying the estimation problem.

[19] See Hinkley (1970). Simulation evidence in related contexts such as those investigated by Goldfeld and Quandt (1972) and Quandt (1960) indicates that, where there are sufficient observations in each regime, the sampling distributions of the processes in each regime conform well to their distributions under the assumption that the switch point is known.

[20] We assume that the estimation errors in the volatility series are sufficiently small that they do not affect the ranking of the volatilities, at least in the neighborhood of the switch in regimes.

[21] If the model is correct, the two sample means are consistent estimates of the true means in the two regimes, with standard errors approximated by $\sigma/\sqrt{t_1}$ and $\sigma/\sqrt{(T-t_1)}$, where T and t_1 are the total number of observations and the number in regime one, and σ is the standard deviation of the residuals approximated by $\sqrt{(SSE/(T-3))}$. Consistency of the sample means can be shown for a limit as the number of observations in each regime increases. We could not reject the hypothesis of homoskedasticity across the two regimes.

Table 4
Estimates of the announcement effect of adjustable rate debt: matched sample of 47 pairs from 1978–1986

Panel 1: Mean abnormal returns.

	Mean	t-ratio	Sum of squared errors
All firms	−0.95	−1.92	0.0543
Banks	−1.92	−2.24	
Banks - Others	−1.57	−1.56	

Panel 2: Switch in regimes based on Equation (5) in the text. Ranking variable is expected inflation volatility.

	No.	α_1	t-ratio	α_2	t-ratio	LR test	DW
All firms	17/30	0.984	1.30	−2.05	−3.59	9.80	1.97
Banks	7/11	−0.556	−0.40	−2.79	−2.51		2.48

Panel 3: Switch in regimes based on Equation (5) in the text. Ranking variable is T-bill rate volatility.

	No.	α_1	t-ratio	α_2	t-ratio	LR test	DW
All firms	18/29	0.794	1.06	−2.04	−3.46	8.63	2.05
Banks	7/11	−0.556	−0.40	−2.79	−2.51		2.47

Panel 4: Regression of abnormal return on inflation volatility.

	Int.	t-ratio	Slope	t-ratio	LR test	DW
All firms	0.48	0.58	−1.03	−2.13	4.54	1.83
Banks	−0.70	−0.46	−0.87	−0.98		2.61

Abnormal return, stated as a percentage, is defined as the 2-day stock price reaction for adjustable-rate announcements minus the 2-day stock price reaction for a control fixed-rate announcement. Volatilities are variances of monthly rates stated in percentage per year. Banks are the subsample of 18 pairs with SIC code 602. Panel 1 contains simple means for the entire time period. Banks - Others is the difference in means for the two subsamples. The t-ratio is based on a common estimate of variance and a difference in means test applied. Panels 2 and 3 contain estimated means from two regimes as in Equation (5) of Section 3.1. For Panel 2, MLE of switch point is at variance 0.654. Linear regression is for comparison purposes. Likelihood ratio is $-2\log\lambda$ where λ is the likelihood ratio (the ratio of the sum of squared errors from Panel 1 divided by the sum of squared errors from the model in the respective panel, all raised to the power of 47/2). The statistic conforms closely to a χ^2_3 [see Hinkley (1970)]. The statistic in Panel 4 would conform asymptotically to a χ^2_1 under the null hypothesis. The statistic in Panel 4 is reported for comparison with other panels. LR test reported for total sample only. DW = Durbin-Watson statistic with data ordered by volatility. No. = number of firms in first regime/number of firms in second regime.

assumption that the true but unknown standard deviation is the same for both subsamples.

Panel 2 shows the two means and associated t-ratios when a threshold for volatility is estimated. The mean of the first regime is 0.98 percent and insignificant (t = 1.3, while the mean of the second regime is -2.05 percent with a t-ratio, of -3.59. The likelihood ratio (LR) for the model suggests a significant overall fit at about the 2 percent level.[23] The volatility threshold is at 0.654, which splits the sample into

[22] Throughout the empirical work, we base our standard errors on the cross-sectional dispersion of the abnormal return observations around the various fitted models.

[23] The data suggest that our inferences can be based on applicitions of normal theory. A Pearson goodness cd fit test with eight groupings of equal probability applied to the residuals of the fitted

17 announcements below and 30 announcements above the threshold, The estimated threshold is shown graphically in Figure 2 as a horizontal line at standard deviation 0.81.

The results support the signaling model in several ways. The first regime has a positive mean and the second has a negative. The LR statistic indicates significance for the difference in means and the *t*-ratio shows that the mean of the second regime is also significant. The absolute values of the two means are close in that it is not possible to reject the hypothesis that the sum of the means equals zero based on an asymptotic *z* test for the sum of normal, independent random variables. That the mean of the first regime is not sign&ant at conventional levels appears to reflect primarily the small sample size as evidenced by the tests on the raw sample discussed later.

Results for the banking subsample are also broken out separately in Panel 2. The means and *t*- ratios are based on the estimated threshold from the complete matched sample. The 7 banks in the low volatility regime have a mean of -0.556 percent, which is insignificant, while the 11 banks in the high volatility regime have a mean of -2.79, percent, with a t-ratio of -2.51. The sample size does not allow rejection of the equality of means for the banking subsample, either across the two regimes identified with the complete matched sample, or across the two regimes chosen as the maximum likelihood partition of the banking subsample. As a result, the data do not allow us to draw conclusions about the possibility that the volatility threshold for banks is below the threshold for the entire matched sample. A lower threshold is predicted if the degree of covariability between real returns and the real interest rate is more positive for banks.

Panel 3 shows results for the model based on T-bill rate volatility as a surrogate for the volatility of expected inflation. As mentioned, the two volatility series are highly correlated. If the variation in T-bill rate volatility is driven by variation in the volatility of expected inflation, it serves as a proxy for expected inflation volatility that avoids the explicit modeling of inflation expectations. For the complete matched sample, one firm changes from the high regime to the low regime causing both means to move slightly toward zero. The mean in the second regime rises from -2.05 to -2.04. The t-ratio for this mean

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model yielded a test statistic of 2.5, which is to be compared with a χ^2_5. The median of the residuals is -000.71, with 25 negative residuals and 22 positive.

For models with a single switch in means, Hinkley (1970) suggests comparing the LR statistic to a χ^2_3. Our value of 9.8 attains the 2 percent level of significance under **this** comparison. Since the switch model is nonlinear, however, statistical inference can be sharpened by simulating an empirical distribution of test statistics under the null hypothesis. We searched for a switch point and two means with our algorithm in 500 samples of randomly generated normal errors paired with the same volatility measures used in the study. There were 3 *t*- ratios exceeding 3.588 in absolute value (none negative) and 13 cases of LR statistics exceeding 9.8.

rises to -3.46. The LR statistic falls to 8.93, which is still above the 5 percent cut point for a x_3^2. Results for the banking subsample are unchanged from Panel 2.

Table 5 shows results for the raw sample, which corroborate the conclusion of a switch in means. Moreover, the estimates of the means are attenuated relative to the estimates from the matched sample, consistent with the prediction that a raw sample will contain pooling observations -and suffer from a 'bias created by failing to control for differences in prior anticipation across the sample. In Panel 1, the mean abnormal return is an insignificant -0.075 percent. If the raw sample had an equal number of fixed and adjustable rate announcements, the mean would equal zero by construction. When volatility is based on the expected inflation series, Panel 2 reveals that the mean of the first regime is 0.537 percent with a t-ratio of 2.13. The second regime has a mean of -0.413 with a t-ratio of -2.21. The LR statistic for the overall model is slightly lower than for the matched sample at 9.09. This value is significant at about the 4 percent level. There are 88 firms in the low volatility regime and 158 in the high volatility regime. Panel 3 shows results for the model based on T-bill rate volatility. The explanatory power of the model falls as judged by the drop in the LR statistic from 9.09 to 6.58. The mean in the high volatility regime is not changed appreciably, although there are 40 fewer firms represented. The mean is -0.48 percent with a Z-ratio of -2.19.

3.5 Extensions of the basic empirical model

In this section, the empirical tests are broadened to account for several extensions. The first is to estimate a smooth relation between abnormal return and volatility level. Such a process can be motivated on two grounds. First, there is the possibility that different firms have different volatility thresholds but that the frequency of firms above their own threshold increases in volatility. Second, suppose firms have different degrees of asymmetric information cross-sectionally. Firms with larger informational asymmetries will find themselves unable to signal when the difference in default risk between the two strategies is low, because it pays for the lowerquality types to mimic. On the other hand, firms facing better-informed outsiders and, thus, having less to gain from communicating favorable information to the market, will not want to signal if the difference in default risk between debt strategies is large. This logic implies that more information is revealed on average by debt announcements that are made when the difference in default risk between the two strategies is relatively large. Then, in the neighborhood of the threshold, one would expect the abnormal return on a matched pair to increase in magnitude as volatility departs from the threshold.

Table 5
Estimates of the announcement effect of debt issues: raw sample of 246 fixed and adjustable rate announcements over the period 1978-1986

Panel 1:	**Mean abnormal returns.**						
	Mean	***t*-ratio**					
	−0.075	**−0.32**					
Panel 2:	**Switch in regimes based on Equation (5) in the text. Ranking variable is expected inflation volatility.**						
	No.	**α_1**	***t*-ratio**	**α_2**	***t*-ratio**	**LR test**	**DW**
	88/158	**0.537**	**2.13**	**−0.41**	**−2.21**	**9.09**	**2.19**
Panel 3:	**Switch in regimes based on Equation (5) in the text. Ranking variable is T-bill rate volatility.**						
	No.	**α_1**	***t*-ratio**	**α_2**	***t*-ratio**	**LR test**	**DW**
	129/118	**0.299**	**1.43**	**−0.48**	**−2.19**	**6.58**	**2.20**
Panel 4:	**Regression of abnormal return on inflation volatility.**						
	Int.	***t*-ratio**	**Slope**	***t*-ratio**	**LR test**	**DW**	
	0.15	**0.69**	**−0.10**	**−1.43**	**2.05**	**2.12**	

Abnormal return is defined as the 2-day stock price reaction minus the avenge for the raw sample. Abnormal returns for fixed-rate announcements are entered with a minus sign for the analysis in this table. Volatilities are variances of monthly rates stated in percentage per year. Panel 1 contains the mean for entire time period. Panels 2 and 3 contain estimated means from two regimes. For Panel 2, MLE of switch point is at variance 0.892. LR test is $-2\log\lambda$ where λ is the likelihood ratio (the ratio of the sum of squared errors from Panel 1 divided by the sum of squared errors from the model in the respective panel, all raised to the power of 246/2). The statistic conforms closely to a χ_3^2 [see Hinkley (1970)]. The statistic in Panel 4 would conform asympromatically to a χ_1^2 under the null hypothesis. The statistic in Panel 4 is reported for comparison with other panels. DW = Durbin-Watson statistic with data ordered by volatility. No. = number of firms in firsts regime/number of firms in second regime.

We consider first a linear regression of abnormal return on volatility. Regression results are reported in the last panel of Tables 4 and 5. For the complete matched sample, the slope is negative and significant with a t-ratio of –2.14. For the raw sample, the regression slope is not significant. Our primary concern, however, is not the regression slopes, which we expect to be negative given the evidence from the switch in means model, but the fit of the linear model. The likelihood ratios indicate that regressions fit the data less well than the switch in means model, supporting the prediction of an abrupt transition.

To gain further insight into model specification, the final columns in Tables 4 and 5 report Durbin-Watson statistics for all of the models with the observations ranked by volatility.[24] None of the statistics are significant at conventional levels. The closest is the regression for the

[24]If data were generated by a hear model and a switching regimes model erroneously fit, the Durbin-Watson statistic might be able to pick up a model misspecification because the residuals would be dumped in groups according to the independent variable, which is ordered Similarly, data generated by a switching regimes model might yield a large Durbin-Watson statistic if a linear regression were fit. The power of this test for detecting such model misspecifications has not, to our knowledge, been studied

banking subsample. Unfortunately there is little difference between the Durbin-Watson statistics across the models, so it is not possible to draw strong conclusions about model specification based on these comparisons.'

Obviously, linear regression assumes an extreme degree of smoothness in the relation between abnormal return and volatility. A more flexible functional relation is provided by a nonlinear model of the form

$$R_j = \alpha_1 + (\alpha_2 - \alpha_1)D_j + e_j \qquad (8)$$

where D_j is a function of the volatility, b_j, with the property that $D_j = 0$ at low h_j and $D_j = 1$ at high h_j. We use the function $D_j = \{1+\exp[k - m\ln(h_j)]\}^{-1}$ where k and m are estimated parameters and $m > 0$.[25] Estimating the model through the method of linearization yields very similar results to those reported above. For the matched sample, the transition is abrupt with $D = 0.5$ at $h = 1.13$.[26] The estimate of α_1 is 0.436 percent (t-ratio = 1.06) while α_2 is -2.37 percent (t-ratio = -5.3). The overall fit of the model is close to the simple, two-regime model, yielding an LR statistic of 9.0.

A final extension is to address the possible nonmonotonicity of abnormal returns in the signaling regions by working with the sign of each abnormal return. A general model would seem to predict that the frequency of firms signaling favorable information with fixed-rate debt and unfavorable information with adjustable increases in volatility, as long as the data are matched. A test based on the percentage negative for the matched sample sheds light on this implication of the theory, in addition to providing an indication of the extent to which outliers are important in our previous tests.

We constructed a contingency table of positive and negative abnormal returns in the two regimes. Partitioning the observations for the matched sample yields 3 of 17 negative in the first regime and 19 of 30 negative in the second regime. The Pearson chi-squared statistic (with Yates' correction) is 7.35, rejecting independence at beyond the 1 percent level.

3.6 Multivariate regression tests

To study the influence of the broader market environment on the information conveyed by debt type, we regressed abnormal returns on

[25]This approach is fashioned after the discussion in Goldfeld and Quandt (1972) where the focus is on a transition that is based on time rather than a time-varying explanatory variable. The logic is parallel.

[26]The estimated values are $k = 13.44$ and $m = 106.5$. If alternative levels of h are substituted into the function, one an see that the transition is abrupt since h = 1.08 implies $D = 0.005$ and $h = 1.25$ implies $D = 0.999$. The data contain three variations between these values of h.

several contemporaneous macrovariables that are potentially related to debt choice.. Multiple regressions are reported in Table 6 along with the single variable regressions for comparison. The explanatory variables, as described in Section 3.3, are expected inflation volatility, the slope of the term structure, the default spread, and our expected inflation measure. The motivation for including the slope of the term structure is that a naive model of debt issue choice might build on such a variable. If financial markets are imperfect and managers wish to minimize expected interest payments, managers might issue adjustable-rate debt when the term structure is rising relatively more quickly and issue fixed-rate debt when the term structure is rising at a relatively slow rate. To the extent that investors view this behavior as clever, the market reaction to adjustable-rate announcements might be less negative when the term structure is steep. As will be shown shortly, however, the data do not support this prediction.

The default spread variable is included because there may be differences in the default risk of fixed- and adjustable-rate issues for reasons other than those discussed in our model. A naive model of debt choice might motivate including a default spread variable. Sup pose adjustable-rate debt has lower default risk for all firms but firms like to issue fixed-rate debt for other reasons. The benefit of issuing adjustable-rate debt would be higher when the default spread is high and vice versa. If problems of rational prior anticipation by investors are ignored, one might predict that the difference in market reaction between adjustable- and fixed-rate debt announcements should increase with default spreads.

We include expected inflation because the levels of expected inflation and interest rates seem to be potentially interesting variables that might be correlated with the market reaction to debt announcements in a more complex model of debt choice. Our model predicts that the level of expected inflation should not influence the relative default risk of the borrowing strategies. Empirical results for the level of interest rates are similar and not reported.[27]

Two dependent variables are considered in the regressions. Panel A shows the regression results when abnormal return is used. For the matched sample, the only coefficients in the regressions that exceed twice their standard errors are the coefficients on expected inflation volatility. None of the coefficients are significant for the raw sample. Panel B shows results when the residuals from the switch in means

[27]The level of interest rates is not significant when used in place of expected inflation. The correlation between interest rates and inflation volatility, however, causes volatility to lose significance in multiple regressions. If interest rates are first projected on volatility and the residuals used in the multiple regression, the residuals contribute almost nothing and volatility remains significant.

Table 6
Regression relationships between the announcement effects of debt issues and various microvariables over the period 1978–1986

Independent variables:	Intercept	Inflation volatility	Term slope	Default spread	Expected inflation
Panel A:	Dependent variable is abnormal return at announcement.				
	Sample: Matched sample of 47 pairs of announcements.				
Coefficient	1.74	−1.12	−0.51	0.30	−0.07
(*t*-ratio)	(0.77)	(−2.06)	(−0.76)	(0.18)	(−0.27)
	0.48	−1.03	—	—	—
	(0.58)	(−2.13)			
	Sample: Raw sample of 246 fixed and adjustable rate announcements.				
Coefficient	0.63	−0.06	−0.01	−0.31	−0.03
(*t*-ratio)	(0.86)	(−0.68)	(−0.09)	(−0.92)	(−0.51)
	0.15	−0.1	—	—	—
	(0.69)	(−1.43)			
Panel B:	Dependent variable is the abnormal return residual from the switch in regimes model of Equation (5) in the text.				
	Sample: Matched sample of 47 pairs of announcements.				
Coefficient	0.57	0.15	0.63	−1.50	0.07
(*t*-ratio)	(0.27)	(0.30)	(0.28)	(−0.92)	(0.30)
	0.50	−0.04	—	—	—
	(0.06)	(−0.08)			
	Sample: Raw sample of 246 fixed and adjustable rate announcements.				
Coefficient	0.40	0.08	−0.00	−0.22	−0.05
(*t*-ratio)	(0.56)	(0.96)	(−0.05)	(−0.49)	(−0.85)
	−0.06	0.2	—	—	—
	(−0.28)	(0.37)			

Abnormal returns are as defined in the text and in Tables 4 and 5. Abnormal return residuals are abnormal returns minus the mean in the appropriate regime as shown in Panel 2 of Tables 4 and 5. Inflation volatility is the variance of monthly rates stated in percent per year. Term slope is the difference in yield between 5-year and 1-year treasuries. Default spread is the difference in yield between BBB-rated and AAA-rated corporate bonds. Expected inflation is as defined in the text, stated in percentage per year. Univariate results with abnormal return as dependent variable are repeated from Tables 4 and 5.

model of Panel 2 in Tables 4 and 5 are used as dependent variable. Although no correction has been made for the fact that the dependent variable is a residual from a first pass through the data, the results suggest that the linear models contribute little beyond the switch model. There are no coefficients with magnitudes in excess of their standard errors. In total the multiple regressions reinforce the conclusion that the market reaction to debt choice is driven primarily by inflation volatility with a threshold in the data.

4. Summary and Conclusions

The data support the notion that firms were able to signal quality with their debt choice within regions of inflation volatility over the 1978 through 1986 period. Our attempt to exploit a threshold in the market

reaction to debt announcements allows us to infer that firms choosing the more risky strategy signaled favorable information. Matching on the basis of industry and volatility strengthens the results by eliminating pooling observations and reducing attenuation in the estimate of the signal's impact. We find all of these results encouraging.

An interesting side issue not addressed in our model concerns the potential use of interest-rate swaps to uncouple the interest payment pattern from the initial debt type chosen. The swap market began in 1982 and has gained popularity over the ensuing decade. One might conjecture that firms would attempt to use swaps, following a debt issue, to neutralize the cost of the initial debt signal. If such *ex post* swap transactions were costless to managers, the signaling equilibrium based on debt type would unravel. Yet our empirical results suggest that swaps failed to undermine the credibility of the signal sent by debt choice.

A limitation of our experiment is that, in reality, firms face 'a continuum of debt choices involving differential exposure to expected inflation risk. Among the possible features that affect this exposure are maturity, call and conversion provisions, and adjustable-rate caps and floors. A more general model would allow firms to signal with a continuous variable that measures the sensitivity of default risk to changes in expected inflation. This is a promising area for future research.

Appendix A

Proof of Propositions 1 and 2. The regularity condition of Assumption 5 in the text allows us to focus on the variance of the log of the coverage ratio in proving the propositions. From Assumptions l-5 in the text, the natural logs of the coverage ratios under adjustable and fixed rate debt are, respectively,

$$\begin{aligned}\ln(c) - \ln(p^A) &= \pi_0 + \pi_1 + \ln(x) - r_{f0} - r_{f1} - \Delta^A \\ &= E_0(\pi_0) + u_0 + E_1(\pi_1) + u_1 + \ln(x) \\ &\quad - E_0(i_0) - E_0(\pi_0) - E_1(i_1) - E_1(\pi_1) - \Delta^A \\ &= u_0 + u_1 + \ln(x) - E_0(i_0) - E_1(i_1) - \Delta^A; \quad \text{(A1)}\end{aligned}$$

$$\begin{aligned}\ln(c) - \ln(p^F) &= \pi_0 + \pi_1 + \ln(x) - E_0[r_{f0} + r_{f1}] - \Delta^F \\ &= E_0(\pi_0) + u_0 + E_1(\pi_1) + u_1 + \ln(x) \\ &\quad - E_0(i_0) - E_0(\pi_0) - E_0(i_1) - E_0(\pi_1) - \Delta^F \\ &= u_0 + u_1 + \ln(x) - E_0(i_0) - E_0(i_1) \\ &\quad + E_1(\pi_1) - E_0(\pi_1) - \Delta^F; \quad \text{(A2)}\end{aligned}$$

where u_0 and u_1 are inflation forecast errors. Define ϕ as the difference between the variances of the natural logs of the coverage ratios. Using the stochastic elements in the last lines of Equations (A1) and (A2) and eliminating common variances yields

$$\begin{aligned}\phi &= \mathrm{Var}_0[\ln(c) - \ln(p^A)] - \mathrm{Var}_0[\ln(c) - \ln(p^F)] \\ &= \mathrm{Var}_0[E_1(i_1)] - 2\,\mathrm{Cov}_0[\ln(x), E_1(i_1)] \\ &\quad - \mathrm{Var}_0[E_1(\pi_1)] - 2\,\mathrm{Cov}_0[u_0, E_1(\pi_1) + E_1(i_1)] \\ &= \mathrm{Var}_0[E_1(i_1)] - 2\,\mathrm{Cov}_0[\ln(x), E_1(i_1)] - 3\,\mathrm{Var}_0[E_1(\pi_1)] \\ &\quad - 2\,\mathrm{Cov}_0[E_1(\pi_1), E_1(i_1)] \\ \phi &= \phi\{\mathrm{Var}_0[E_1(i_1)], \mathrm{Var}_0[E_1(\pi_1)], \mathrm{Corr}_0[E_1(i_1), E_1(\pi_1)], \\ &\quad \mathrm{Cov}_0[\ln(x), E_1(i_1)]\} \qquad \text{(A3)}\end{aligned}$$

where use is made of the assumption that inflation follows a random walk and the default premia are known at $t = 0$. The regularity condition of Assumption 5 states that the sign of ϕ equals the sign of the difference in default probabilities, proving Proposition 1.

To prove Proposition 2, notice that $\phi = 0$ is a concave quadratic equation in $\{\mathrm{Var}_0[E_1(\pi_1)]\}^{1/2}$ when the other elements in Equation (A3) are held fixed. While the proposition is stated in terms of a constant correlation between expected inflation and the expected real rate, the function is also concave quadratic if the covariance between expected inflation and the expected real rate is held fixed. The condition assumed in the proposition that fixed-rate debt has lower default risk when there is no price uncertainty implies that $\phi > 0$ when $\{\mathrm{Var}_0[E_1(\pi_1)]\}^{1/2} = 0$. That there exists a single positive real root to the equation $\phi\{\{\mathrm{Var}_0[E_1(\pi_1)]\}^{1/2}\} = 0$ follows from the fact that the equation is concave and quadratic, proving the proposition. ■

Appendix B

Proof of Proposition 3. We examine equilibrium configurations that satisfy the Perfect Bayesian Equilibrium (PBE) concept, which in our model is equivalent to the Sequential Equilibrium criterion of Kreps and Wilson (1982) as described in Fudenberg and Tirole (1991). Denote firm quality by q and refer to the low- and high-default risk-borrowing strategies as L and H. Since managers bear higher expected personal default costs by choosing the high-default risk strategy, any feasible equilibrium has lower-quality types choosing the riskier strat-

egy with probability equal to or lower than that of higher type;[28] that is, $\Pr\{H \mid q'\} \geq \Pr\{H \mid q\}$ for $q' > q$. It follows that average firm quality can never be lower in the high default risk market.

Consider first equilibrium outcomes in pure strategies. A separating outcome has all types $q < q^*$ selecting the lowdefault risk strategy and all types $q \geq q^*$ choosing the high one. Separation exists if such a conjecture is self-fulfilling. Define $EVo(q \mid q \geq q^*)$ and $EVo(q \mid q < q^*)$ as the expected values of the firm at $t = 0$ conditional on, respectively, $q \geq q^*$ and $q < q^*$. Under the assumed conjecture, types $q \geq q^*$ are better off with the high-risk strategy H if, for all $q \geq q^*$,

$$EV_0(q \mid q \geq q^*) - EV_0(q \mid q < q^*) > \Omega\{\Pr[c - p^H < 0 \mid q] - \Pr[c - p^L < 0 \mid q]\}, \quad \text{(B1)}$$

while types $q < q^*$ are better off with the low-risk strategy L if the inequality in Equation (Bl) is reversed for all $q < q^*$. Since the right-hand side of Equation (Bl) is decreasing in firm quality, the solution, q^*, to the equation

$$EV_0(q \mid q \geq q^*) - EV_0(q \mid q < q^*) = \Omega\{\Pr[c - p^H < 0 \mid q^*] - \Pr[c - p^L < 0 \mid q^*]\} \quad \text{(B2)}$$

satisfies both inequalities. There must exist specifications for which Equation (B2) can be solved because both sides of Equation (B2) are positive.[29] ■

Similar logic shows the existence of pooling equilibria, where all firms choose the same debt type. Pooling in the low-default-risk market is more efficient because default costs are minimized. Moreover, pooling in low-default-risk debt is robust to a wider range of out-of-equilibrium beliefs held by investors about which firms might defect from the pooling equilibrium. For example, if out-of-equilibrium moves result from random mistakes and the probability of a mistake is independent of firm type (i.e., investors hold *passive* conjectures about types making out-of-equilibrium moves), only the pooling outcome in lowdefault-risk debt is sustained. While it is possible to con-

[28] Any separating or semiseparating equlibrium whereby lower types choose the riskier strategy more frequently than higher types violates the Spence-Mirrlees sorting condition assumed in Assumption 6.

[29] In discussing separating equilibria similar to ours, Riley (1979) shows that the potential for alternative offers by price searching buyers (Le.. yield searching investors in our model) can create instability at the lower endpoint of the signal, Causing the equilibrium to unravel. In our environment, however, the separating equilibrium is stable in a "reactive" sense. Details are available upon request.

struct beliefs under which firms may pool in high-risk debt,[30] we hypothesize that pooling in low-default-risk debt is the dominant alternative to separating equilibria when pure strategies are employed by firms?'

Finally, consider mixed strategies. The following demonstrates that at most one quality level will randomize, ruling out a measurable set of firms employing mixed strategies in equilibrium. If any type q' randomizes, the manager must be indifferent between the two borrowing strategies. Type q' is indifferent if

$$EV_0(q \mid H) - EV_0(q \mid L) = \Omega\{\Pr[c - p^H < 0 \mid q'] - \Pr[c - p^L < 0 \mid q']\}, \quad \text{(B3)}$$

where $EVo(q \mid H)$ and $EVo(q \mid L)$ are, respectively, the values at $t = 0$ of the firms choosing H and L determined by integrating over all types weighted by their probabilities of choosing strategies H and L. Since the left-hand side of Equation (B3) is independent of q', while the right-hand side decreases with q', there exists at most one quality that is indifferent between the two borrowing strategies. With a continuum of firm types, any one quality is of measure zero.

In summary, only equilibria in pure strategies are meaningful. Among these are separating equilibria whereby higherquality types choose the high-default risk strategy and pooling equilibria in both strategies. We emphasize the pooling equilibrium in the low-risk debt type as the most robust and efficient.

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[30] For example, the belief that a defector of equilibrium is always of the worst type may sustain a high-risk pooling out come.

[31] Usually, one can prune multiple informational equilibria by using stronger equilibrium refinements. For example, a well-known result is that in a signalling game with a continuum of types and a continuous signal, the universal divinity criterion of Banks and Sobel (1987) eliminates all equilibrium but the most efficient signaling equilibrium. Interestingly, the universal divinity criterion does not eliminate either of the two pooling equilibria in our model. The reason is that the signal **is** a binary variable. The pooling equilibrium in high-default-risk debt, however, **is** eliminated if we give investors the opportunity to make counteroffers to companies [as in Rothschild and Stiglitz (1976)], When firms pool in high default risk debt, an offer to buy low-risk bonds at a price marginally lower than that consistent with average firm quality will be attractive to every firm and cam money for the investor.

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