Individual Stock-Option Prices and Credit Spreads*

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In a seminal contribution, Merton (1974) developed the structural firm-value approach to the valuation of corporate bonds. According to this model, corporate debt is simply riskless debt combined with a short position in a credit put option, struck at the face value of the debt. Bankruptcy is triggered, and the credit put exercised, when the firm value drops below the face value of the bond. Subsequently, a large literature has developed and extended this work and a number of papers have studied the empirical implications of the structural firm-value approach to credit risk (see for instance Eom, Helwege and Huang (2003), or Duffie and Singleton (2003) for a textbook treatment). An important finding in this work is that it is challenging to explain variation in credit spreads based solely on credit-risk factors, even when accounting for liquidity proxies (Collin-Dufresne, Goldstein and Martin (2001), henceforth CGM). Furthermore, the magnitude of observed credit spreads is difficult to rationalize in the context of a structural firm-value model (Huang and Huang (2003)).

In this paper, we propose to consider market-based proxies for two fundamental theoretical determinants of credit spreads, volatility and jump risk, that are directly observed in the market for individual options on the equity of the issuing firms. Since corporate bonds embed a short position in out-of-the-money puts, it is natural to consider precisely the market for out-of-the-money puts. Also, whenever the underlying firm value is time-varying and stochastic, this will be reflected in the implied volatility of equity options. For example, the implied volatility will shoot up as the firm approaches the default boundary. Traded individual options directly encode the assessment of market participants of this volatility of equity and of its risk, and would therefore be expected to contain forward-looking information that is highly relevant for credit risk. It is important to point out that traded individual options should only add additional information about credit risk not already captured by equity and riskless debt if the options are indeed non-redundant securities. However, there is ample evidence in the option-pricing literature for violations of the complete-market assumption of the Black-Scholes model, and of priced jump and volatility risk (see Bakshi and Kapadia (2003b) and Bakshi, Kapadia and Madan (2003) for individual options).[1] While the value

[1] The importance of jumps and stochastic volatility in the equity index is studied in Andersen, Benzoni and Lund (2002) and Eraker, Johannes and Polson (2003). Evidence of priced jump and/or volatility risk in

of the equity of an issuer will certainly also be affected by volatility and jump risk, and will therefore contain information about these determinants of credit risk, traded equity-options load more directly on these factors. In particular, both at-the-money options as well as out-of-the-money put options may be needed to fully capture and disentangle their respective effects. Whether individual stock options do indeed add relevant and quantitatively important information for credit risk is ultimately an empirical question and this constitutes the main focus of our paper.

While prices (or equivalently, implied volatilities) of individual options have not previously been suggested as potential determinants of credit spreads on corporate bonds, CGM used information in index options in their analysis of the determinants of credit spread changes. Index options however can by definition not capture firm-specific information and are therefore unable to explain cross-sectional differences in credit spreads across issuers. Additionally, it has been argued in the option-pricing literature that the prices of index options (especially of out-of-the-money puts) may be subject to specific demand and supply shocks (see Bates (2002) and Bollen and Whaley (2003)) and are actually very puzzling from the point of view of rational asset pricing (Bondarenko (2003a and 2003b)). Only if these same 'local' demand and supply shocks equally affect the market for credit risk would for instance the implied-volatility skew of index options be expected to be relevant for understanding credit risk.

Instead, we propose to explain credit spreads over time and across issuing firms based on implied volatilities and implied-volatility skews of the individual options on the issuers' equity. The implied-volatility of at-the-money options is a natural proxy for the volatility of the issuing firm. In recent work, Campbell and Taksler (2003) document a very strong relationship between the historical volatility of equity returns and bond yields. Individual options may provide us with a superior proxy of the volatility of the issuer since the measure is forward-looking rather than historical in nature. Furthermore, to the extent that volatility risk matters and is priced, this would be captured by implied volatilities, but never by a

index options is presented in Ait-Sahalia, Wang and Yared (2001), Bakshi, Cao and Chen (1997), Bakshi and Kapadia (2003a), Buraschi and Jackwerth (2001), Coval and Shumway (2001), Pan (2002) and Rosenberg and Engle (2002), among others. See also Bates (2002) for an excellent survey.

historical measure. As a second important explanatory variable we suggest the implied-volatility skew, which is interpreted as measuring the firm's jump risk, stemming from time-variation in the likelihood and severity of a downward jump in firm value. The results of Bakshi and Kapadia (2003b) and of Bakshi, Kapadia and Madan (2003) suggest that individual-option skews are less subject to changes in the demand for general portfolio insurance.

We use a panel of weekly data on US corporate bond prices and individual option prices of 69 firms, for the 1996-2002 period. In our benchmark analysis, we perform a panel regression of the level of credit spreads on a number of explanatory variables.

As our first contribution, we show that option-implied volatilities are extremely successful in explaining credit spreads, both over time and across firms. In particular, even though we impose a tight panel structure on the coefficients (using a pooled regression), implied volatilities alone can explain close to one third of the total variation in credit spreads. The coefficients on at-the-money implied volatility are highly significant both economically and statistically. Unlike alternative measures of volatility, our proxy is very robust to the inclusion of a large number of control variables. The implied volatility skew also manifests itself as a highly significant explanatory variable, albeit with somewhat less economic impact.

Importantly, option prices are shown to contain substantially more information about credit spreads than do credit ratings. The explanatory power in a pooled regression of credit spreads is 5 to 15 percentage points higher (depending on the maturity of the bonds) when regressing on option-based information than when using credit ratings as an explanatory variable. Interestingly and consistent with the structural firm-value model, we find that the sensitivity of credit spreads to volatility and crash imminence is much larger for poorly rated debt (BBB+ or worse) than for bonds with strong credit ratings (A- or better). We also document that option-implied volatilities anticipate downward credit rating migrations in striking way.

We provide further empirical support for our claim that individual options are relevant for understanding credit risk by showing the impact of option-market liquidity on the credit spreads of short-maturity bonds. Finally, because our measures are based on traded assets, trading strategies can be used to exploit the information that options carry about credit risk. We therefore analyze how individual options can be used to hedge corporate bond returns.

The rest of the paper is structured as follows. Section 1 describes the bond and option data we use and presents summary statistics. The benchmark regressions explaining credit spreads are reported in section 2. Section 3 considers a number of extensions and additional control variables, as well as firm fixed effects and time dummies, and studies the effect of option-market liquidity on credit spreads. Credit ratings are introduced in section 4. Section 5 investigates whether credit risk in corporate bonds can be hedged using traded options on the equity of the issuer. Finally section 6 concludes.

1 Data Description

1.1 Corporate Bond Prices

The data on US-dollar corporate bond prices are taken from the Bloomberg Corporate Bonds Database (BCBD), which contains mid-quotes for corporate bond prices. Besides these mid-quotes, the dataset contains for each bond the maturity date, the coupon size and frequency, the S&P credit rating, the firm's industry sector, and the amount issued. We collect data from January 1996 until September 2002, for a total of 351 weeks.

We restrict ourselves to a set of 69 firms for which both corporate bond data and equity option price data are available. This is a subset of the set of firms analyzed by Duffee (1999). We only use bonds with constant, semiannual coupon payments, and no embedded put or call options or sinking fund provisions. As in Duffee (1999), observations on bond prices with remaining maturity less than one year are dropped. Most bonds are senior unsecured. We only include other bonds, such as subordinated bonds, if a bond has the same rating as the senior unsecured bonds of the particular firm. Most firms are rated investment grade throughout the sample period, but some firms become speculative grade in the last three years of the sample period. Two firms in our sample default, Comdisco and Enron.[2]

[2] The default events however occur after the firms leave the sample and are therefore not driving the results.

In total, we end up with 524 corporate bonds. There are several missing observations in the data, which is typical for corporate bond price datasets. Besides corporate bond price data, we also use Bloomberg data on the 6-month US Treasury bill, and the most recently issued US Treasury bonds with maturities closest to 2, 3, 5, 7, 10, and 30 years.

For our empirical analysis, we use credit spreads of these coupon-paying corporate bonds, defined as the difference between the corporate bond yield and the yield on a government bond with exactly the same maturity and coupon size. Since we do not observe government bond yields for all relevant coupon sizes and maturities, we first estimate the term structure of default-free zero-coupon interest rates. We use the following extended Nelson-Siegel (1987) specification for these zero rates $R(t,T)$:

$$R(t,T) = \delta_{1,t} + \delta_{2,t}\frac{1-e^{-\delta_{3,t}(T-t)}}{\delta_{3,t}(T-t)} + \delta_{4,t}e^{-\delta_{3,t}(T-t)} + \delta_{5,t}\frac{1-e^{-\delta_{6,t}(T-t)}}{\delta_{6,t}(T-t)} \tag{1}$$

Each week, we estimate the parameters $\delta_{1,t}, ..., \delta_{6,t}$ by minimizing the sum of squared bond pricing errors for the Treasury bills and Treasury bonds over these parameters. To account for the fact that long-maturity bond prices are more sensitive to interest rates, we weight each pricing error by the inverse of the duration of the bond. Given this term structure of default-free zero-coupon rates, credit spreads on corporate bonds can readily be calculated. Finally, some bond prices very likely contain data errors. We eliminate observations for which the credit spread is below -50 basis points. Also, we delete the 'middle' observation if the credit spread moves more than 50 basis points in one week, and again more than 50 basis points in the opposite direction in the next week.

1.2 Options

The options data originate from OptionMetrics, LLC. This is a comprehensive dataset, covering all exchange listed call and put options on the entire universe of US equities. The data consist of end-of-day bid and ask quotes, open interest and volume, and cover the period January 1996 to September 2002; there are over 3 million option observations per month in the later part of the sample. In addition, with each option price quote OptionMetrics reports the option's implied volatility (calculated using American or European models, as appropriate). Implied volatility calculations use historical LIBOR/Eurodollar rates for interest rate inputs, and incorporate discrete dividend payments. At any given point in time, exchange traded options on US equities have four distinct expiration dates: there are options expiring over the nearest two months and the next two months of the underlying stock's expiration cycle. So as to keep expiration dates the same across stocks, we only use the prices of options that expire in the month immediately following the current month.

The implied volatility skew is calculated as the (left) slope of the implied volatility smirk: it is the difference between the implied volatility of a put with 0.92 strike-to-spot ratio (or the closest available) and the implied volatility of an at-the-money put, divided by the difference in strike-to-spot ratios.

1.3 Summary Statistics

We now turn to the summary statistics for the dependent variable and for the explanatory variables used in the benchmark analysis. Table 1 reports the following statistics: the global mean of each variable (computed as the average across bonds and firms of the time-series averages), the average time-series standard deviation of the levels of the variables, the average time-series standard deviation of the weekly change of the variables and the cross-sectional standard deviation (across bonds) in the time-series averages. In addition, we also compute the correlations of the explanatory variables with credit spreads, both in the time-series (averaged across bonds and firms) and cross-sectionally (for the time-series averages of firm-specific variables).

The average credit spread in our sample is 103.1 basis points for short-maturity bonds and 110.9 basis points for long-term bonds. Credit spreads are highly volatile and exhibit substantial cross-sectional variation. While credit spreads are expressed in basis points, all other variables are expressed in their actual units. Therefore, the average implied volatility for individual options is 34.8% in our sample. Implied volatilities are also highly volatile, both in levels and in first differences, but exhibit somewhat less cross-sectional variation. Interestingly, the individual implied volatility exceeds on average the individual historical volatility, which can be interpreted as evidence of a volatility risk premium. The same is true for the S&P index. Not surprisingly, the volatility of the S&P index (both implied and historical) is substantially smaller than the average individual volatility. Another important finding is that the individual option-implied skew and individual historical skewness are extremely volatile, both in the time-series dimension and cross-sectionally. Unlike for the index, the individual historical skewness is on average positive in our sample, which is quite remarkable.

In terms of correlations with credit spreads, individual volatility stands out. While the historical measure has a somewhat higher time-series correlation with credit spreads (averaged across bonds and firms), the cross-sectional correlation is most pronounced for option-implied volatility. The time-series and cross-sectional relation between credit spreads and the option-based volatility measure is presented graphically in Figures 1 and 3, respectively. The relation between the implied volatility skew and credit spreads is plotted in Figures 2 (average time-series) and 4 (cross-section). While the average time-series correlation is 25.8%, the (univariate) cross-sectional correlation is rather small. Finally, Table 1 shows that historical measures of S&P volatility and skewness are highly correlated with credit spreads during our sample (82.2% and 71.4%, respectively).

A final important comment is in order with respect to the implied volatility skew. The raw time-series for this variable contains a mechanical and periodic pattern that is simply due to the fact that the options do not have constant time to maturity. Because the effect of the shortening in time to maturity is much stronger for the out-of-the-money option than for the at-the-money option, the difference in their implied volatilities (the skew) changes mechanically as the end of the month nears. This pattern would somewhat blur the information in any regression that contains the skew as an explanatory variable to measure jump risk or the likelihood of a downward jump in firm value. We therefore remove the pattern in a model-free fashion, by simply dividing each weekly observation for the skew of an issuer by the ratio of the time-series average of the skews of options with the same time to maturity to the overall time-series average across options. The summary statistics above, as well as all subsequent results, are for this detrended variable.

2 Benchmark Results

We first turn to the benchmark regressions, where credit spreads are explained by alternative measures of volatility and skewness. Based on the insights of a structural firm-value model with jump-diffusions and stochastic volatility, the effect of volatility on credit spreads is expected to be positive. Option-implied skews can be interpreted as measuring jump risk, i.e. the severity and likelihood of a downward jump in firm value, and should therefore have a positive coefficient. Historical measures of skewness on the other hand are expected to affect credit spreads with a negative coefficient, since credit spreads should widen as the return distribution becomes more negatively skewed.

We only retain bonds for which at least 25 observations are available. The regressions are pooled, imposing the same coefficients over time and on different bonds and firms. We distinguish however between short-maturity bonds (between 1 and 5 years) and long-maturity bonds (at least 7 years to maturity), since the results are sufficiently different, as would be expected economically, to warrant separate regressions. All t-statistics are computed using heteroskedasticity-corrected standard errors. The standard errors ignore potential cross-correlations that may exist between observations, as the same issuer may have multiple bonds in the sample at the same time. On average, the number of bond issues per firm is around 5 for short-maturity bonds (334/66) and 3.375 (189/56) for long-maturity debt. An absolute lower bound for the t-statistics reported below would be obtained by assuming that all bonds of the same issuer are perfectly correlated, and therefore by dividing all the reported t-statistics by $\sqrt{334/66} = 2.25$ for short-maturity bonds and by $\sqrt{189/56} = 1.84$ for long-maturity bonds.

Table 2 reports the benchmark regression results for weekly credit spreads of short- and long-maturity bonds. Results are reported for 4 sets of explanatory variables, where each set includes a measure of volatility and of skewness. The first set considers individual implied volatility and the implied-volatility skew of the issuing firm. This constitutes the main focus and contribution of the paper, since these measures are based on traded individual options. As a second set we use the historical volatility and skewness of the issuer. The last 2 sets of regressors are then option-implied (regression 3) and historical (regression

4) measures for both the volatility and skewness of the S&P index. We systematically consider these 4 sets of regressors because it allows a first exploration of the extent to which option-based information is useful in explaining credit spreads, both in absolute terms and in comparison with regressors previously used in the literature. Bakshi, Madan and Zhang (2003) and Campbell and Taksler (2003) investigate historical individual volatilities (but not skewness), CGM include implied volatility and skew for index options, and Campbell and Taksler (2003) and Huang and Kong (2003) use historical index volatility. Finally, we include all 4 sets of regressors simultaneously (in regression 5) as a first step towards assessing the robustness of the different proxies.

2.1 Individual Variables

2.1.1 Option-Implied Measures

When regressing weekly credit spreads on individual implied volatility, the individual implied-volatility skew and a constant (regression 1), both option-based measures are highly statistically significant, for short- as well as long-maturity bonds. The coefficients have the expected sign: an increase in implied volatility and in implied-volatility skew both widen the credit spread, reflecting the rise in the market's assessment of the firm's volatility and jump risk, respectively.

The coefficients are large, suggesting that the measures are also economically significant. To gauge the economic significance more systematically, it is useful to go back to the summary statistics of Table 1. The cross-firm average of the standard deviation of a weekly change in option-implied volatility is 0.046. Thus according to the estimated coefficient, a one-standard-deviation weekly shock in implied volatility leads to a widening of the credit spread by almost 12 basis points for short-maturity bonds and by 20 basis points for long-maturity debt. The implied-volatility skew has a smaller coefficient, but is much more volatile. A typical weekly one-standard-deviation shock in the implied-volatility skew increases the credit spread of that issuer by slightly less than 10 basis points, for both maturities.

Finally, the R^2 of the regression for short-maturity bonds is 14%. Option-implied volatility and skew alone explain more than one seventh of the cross-sectional and time-series variation in credit spreads, even though short-maturity bonds are used (which are typically harder to explain, see for instance the structural-model approach of Huang and Huang (2003)) and even though the pooled regression imposes identical coefficients across all bonds, across all firms and throughout the sample period. For long maturities however, the R^2 of the regression more than doubles: measures of volatility and jump risk based on individual stock-option prices explain roughly one third of the variation across firms and over time in credit spreads, without the inclusion of any other explanatory variables.

The results so far clearly suggest that traded individual options contain information that is relevant for credit risk, in line with a structural firm-value model with stochastic volatility and jumps. The effects are both statistically and economically significant, and consistently so across maturities. The only difference between the maturity groups is quantitative: in line with earlier work, long-maturity credit spreads are easier to explain than short-maturity spreads, making our results particularly strong.

2.1.2 Historical Measures of Volatility and Skewness

Turning now to historical measures of the volatility and skewness of an issuer's stock return in regression 2, we first of all replicate Campbell and Taksler's finding of a very significant coefficient on historical stock return volatility. We extend their work and also include historical stock return skewness. For short-maturity bonds, the coefficient on historical skewness is statistically significant, but has, surprisingly, the wrong sign. Indeed the coefficient on skewness is positive, suggesting that as the return distribution becomes more negatively skewed, credit spreads become smaller, not larger as would be expected economically. The same counter-intuitive sign obtains for long-maturity bonds, but the coefficient is only marginally statistically significant.

Assessing the economic importance is less straightforward than for option-based measures. Since both historical volatility and skewness are calculated using the past 180 return observations, the weekly change in the measure is by construction bound to be small. This is in fact a major disadvantage over option-based measures, which are more forward-looking and which can and do change substantially from week to week. This is clear for instance when comparing the standard deviations of weekly changes for implied and historical volatility in Table 1 (0.046 versus 0.012). In order to therefore interpret the economic significance of historical proxies more directly, we rescale the variables so that they have the same time-series standard deviation (on average across all bonds) as the corresponding implied measures. That way, we can directly compare the estimated coefficients. Doing this reveals that historical volatility has roughly the same economic effect on short- and long-maturity credit spreads as does option-implied volatility. In light of the economically meaningless sign on historical skewness, it is comforting to see that its coefficient is much smaller than the coefficient on implied skew, especially for long-maturity bonds (0.037 versus 0.257).

Although the R^2's are similar to what we found with our option-based measures in regression 1, we want to emphasize that the explanatory power of the regression relies on an estimated coefficient on skewness that has the wrong sign. This raises the question to what extent historical skewness contains information that is relevant for credit risk. Not only is it a third moment and therefore difficult to estimate, it is also not clear that it constitutes an adequate, forward-looking measure of jump risk. This can be interpreted as first evidence of the usefulness of option-based information. Another way to see this is to realize that the time-series correlation (averaged across all firms) between implied skew and historical skew is only –9%. Historical volatility on the other hand tracks implied volatility much more closely: the average correlation between both measures is 60%. The difference between historical skewness and option-implied skew can also be understood in terms of their respective time-series properties. The implied-volatility skew is not only highly volatile, but also mean-reverts much more quickly than the very persistent historical measure. The autocorrelation coefficient for the option-implied measure is 0.396, while it is 0.957 for historical individual skewness. The fact that historical skewness is so highly persistent is quite natural as it is measured over a rolling horizon. But if the 'true' measure of crash imminence or of jump risk that matters for credit spreads, exhibits less persistence and mean-reverts more quickly (like the implied skew), then the historical might simply be too slow to react to changes in jump risk and might in fact by construction be limited in its ability to pick up the dynamics that matter for credit spreads. If this problem is sufficiently severe, and depending on the length of the rolling horizon, it may well produce the wrong sign in a regression, as in Table 2.

2.2 Aggregate Variables

In order to compare our results with CGM, we now regress credit spreads on option-implied measures of volatility and skewness based on S&P index options. For completeness, we finally also consider historical market-wide measures of volatility and skewness.

2.2.1 Option-Implied Measures

The aggregate implied-volatility measure is statistically very significant, but seems to have less economic impact on credit spreads than individual options. A weekly one-standard-deviation shock in the S&P implied volatility changes credit spreads by almost 6 (short-maturity) to 10 (long-maturity) basis points. While this is about half the economic impact of the individual implied-volatility, it is clear that this shock actually affects credit spreads of all bonds simultaneously and is therefore far from negligible. The S&P implied skew comes in with the wrong sign. Its impact is completely insignificant for short-maturity bonds, but seems very strong for long-maturity debt, both statistically and economically. We will show later however that this counter-intuitive effect is not robust. S&P-based variables cannot explain any cross-sectional variation in credit spreads and only pick up time-series variation. This is reflected in the very low R^2 (2% and 5%). CGM obtained similar results, but using a different methodology: they run bond-by-bond regressions and report average results, while we impose constant coefficients in a pooled regression. Also, they analyze the determinants of credit spread changes, and not of credit spread levels as we do here. The results are therefore not fully comparable, although one would certainly expect a much higher average R^2 in bond-by-bond regressions for our sample.

2.2.2 Historical Measures of Volatility and Skewness

Finally, moving to historical aggregate measures of volatility (as used also in Campbell and Taksler (2003) and in Huang and Kong (2003)) and of skewness, we obtain rather surprising results. The coefficient on both historical S&P volatility and skewness are extremely statistically significant. Once again, historical skewness has a counterintuitive (positive) sign, as was the case for individual historical skewness. Furthermore, the R^2 is remarkably large (18% and 32%), in light of the fact that the cross-section remains completely unexplained by market-wide regressors. This raises the suspicion that the historical S&P variables are simply picking up the economic effect of other economy-wide determinants of credit spreads, such as the business cycle and interest rates. We will address this issue in the next section, where we show that these findings are not robust to the inclusion of economic control variables.

2.3 Individual and Aggregate Variables

Combining all 4 sets of regressors to explain credit spreads gives a first indication about the robustness of the results. As can be seen from regressions 5 in Table 2, the individual option-based measures remain very significant. The point estimates are naturally somewhat smaller, but the economic impact of these variables on credit spreads continues to be nontrivial: a weekly one-standard-deviation shock to either variable moves short-maturity credit spreads by roughly 5 basis points. For long-maturity bonds, the results are particularly striking: individual option-implied volatility emerges as the most important firm-specific determinant of credit spreads. In spite of the inclusion of other proxies for volatility among the set of regressors, individual implied-volatility has now the largest (2.84) and most significant coefficient (t-statistic of 30.05). Its economic impact is substantial: a one-standard-deviation weekly increase in the implied volatility of an issuer widens its credit spread by 13 basis points. An equivalent increase in the implied-volatility skew induces a 6 basis point increase.

The coefficients on the historical individual measures also shrink somewhat, but remain overall significant. This also means that the skew continues to have the wrong sign. Noteworthy is that individual historical volatility becomes remarkably less important for credit spreads in long-term bonds, with a coefficient of 0.40 and a t-statistic of 4.49. Interestingly, the signs of both S&P implied variables flip, due to the correlation structure of the variables included in the regression. In contrast, the individual measures are much more robust to these correlation effects. Historical S&P volatility and skew are also affected by the inclusion of other regressors, but continue to have large and significant coefficients (albeit with the wrong sign for skew). However, the next section will show that the impact of these variables is not robust to the inclusion of additional control variables.

Even though we impose a tight structure on the coefficients through the use of a panel and even though any other explanatory variables are lacking, the pooled panel regression explains 25% of the cross-sectional and time-series variation in credit spreads of short-maturity bonds. For long-maturity debt, the R^2 of the regression without firm-specific measures is 32.5%. Adding the firm-specific option-based measures brings the R^2 to 41.9%. The R^2 of the regression with all 8 explanatory variables is 42%.

At this point, it seems that measures implied from index options behave quite differently in relation to credit spreads than individual option-implied variables, especially the implied-volatility skew. This may be consistent with the conclusions of for instance Bates (2003) and Bollen and Whaley (2003) that specific demand and supply shocks affect the market for portfolio insurance (and for out-of-the-money index puts), such that S&P implied skew is not always a good measure of crash imminence. In fact, Bates even shows that the index implied-volatility skew tends to be highest after a crash, not before. In this case, the S&P implied-volatility skew may not be expected to contain information that is relevant for other markets, like the market for credit risk. At the same time, this may be less of an issue for individual options, making them a more relevant determinant of credit spreads. Indeed Bakshi, Kapadia and Madan (2003) have shown that individual skews have quite different dynamics than S&P skews.

In summary, the results so far suggest that individual options contain relevant information for credit spreads. Especially implied volatility emerges as an important determinant. The implied-volatility skew matters as well, but to a lesser extent. The evidence is strongest for long-maturity bonds. In the next section, we explore the robustness of these findings.

3 Robustness and Sensitivity Analysis

To analyze the robustness of the benchmark results, we now consider a variety of extensions. First, we add a number of regressors that have been shown in the literature to have explanatory power for credit spreads. Second, we introduce year dummies and firm fixed effects to explore to what extent our option-based measures explain cross-sectional versus time-series variation. Finally, we present some evidence of liquidity spillovers between the options market and the corporate bond market, by showing that simple proxies for liquidity in the options market affect credit spreads.

3.1 Control variables

A number of papers have examined the determinants of credit spreads. It is therefore important to investigate whether our option-based variables are just proxying for these determinants or whether they provide indeed additional explanatory power.

A first natural control variable is the firm's past stock return. Kwan (1996), CGM and Campbell and Taksler (2003) document a negative relationship between the firm's past stock return and credit spreads. The equity return can be interpreted as reflecting the firm's health, or alternatively, as being a crude but high-frequency proxy for leverage. In addition, the overall state of the economy may matter and can be captured by the market (S&P) return, as in Longstaff and Schwarz (1995), CGM, Campbell and Taksler (2003) and Huang and Kong (2003). Both the firm and market return are calculated over the past 180 days and obtained from CRSP.

To control for the level and slope of the term structure of interest rates, we include the yield on 2- and 10-year Treasury bonds from Datastream, following Duffee (1998 and 1999), CGM, Campbell and Taksler (2003), Driessen (2003), Elton, Gruber, Agrawal and Mann (2002) and Huang and Kong (2003), among others. The general empirical finding is a negative relationship between default-free rates and credit spreads. One explanation for this effect is given by Longstaff and Schwartz (1995). In their model, a rise in the level of interest rates increases the drift of the risk-neutral process for the value of the firm, thus reducing the risk-neutral probability of default as well as credit spreads. CGM interpret the slope of the term structure as a proxy for the overall state of the economy, as well as a measure of expected future short rates. A negative sign is therefore expected.

The general trend in the level of credit spreads over time is controlled for by including the BAA rate. CGM show it has explanatory power on top of many other variables. Finally, liquidity may be an important factor driving credit spreads. We use the 10-year swap rate as a first proxy for liquidity (as in CGM) and the difference between the 30-day Eurodollar and the Treasury yield as a second control (following Campbell and Taksler (2003)). All these data are obtained from Datastream.

The introduction of these additional economically motivated control variables has a number of interesting implications in Table 3. First, the individual option-based measures turn out to be very robust. Although the coefficients on the implied volatility and implied-volatility skew of the issuing firm shrink, they remain statistically and economically significant. Individual implied volatility comes out as a very important determinant of credit spreads on long-maturity bonds: the coefficient of 2.26 means that a weekly one-standard-deviation shock to implied volatility moves credit spreads for long-maturity bonds by more than 10 basis points. Statistically, it is the single most significant regressor with a t-statistic of 29.86. Interestingly, the skew variable now matters most for short-maturity credit spreads. It should not come as a surprise though that the coefficients become smaller. The controls include for instance the BAA rate, which captures the overall trend in credit spreads. To the extent that the option-implied variables measure this trend as well, it is to be expected that their coefficients become smaller. The fact that they remain important highlights that the measures capture more than just the overall trend, i.e. they carry relevant information beyond the BAA rate.

Individual historical estimates of volatility continue to play a role, but mainly for short-maturity credit spreads. The introduction of other macroeconomic and market-wide variables clearly matters for the aggregate S&P-based measures. These are not robust to the inclusion of additional controls and have in fact the wrong sign in most cases. For instance the coefficient on S&P historical volatility drops from 7.4 in Table 2 (long-maturity bonds, no controls) to essentially zero in the last column of Table 3. This suggests that the earlier results simply reflected to a large extent the correlation of the S&P-based measures with other market-wide variables: for instance both the S&P implied volatility and the S&P historical volatility are highly correlated with the 10-year yield (correlation of -56% and -70% respectively) and with the market return (-44% and -64% respectively).

Consistent with earlier work, we find significant and robust coefficients for the interest rate variables considered: the 2-year yield, 10-year yield, BAA rate and to a lesser extent also the swap rate emerge as relevant explanatory variables for credit spreads. The firm and market return are somewhat important, but have rather small effects economically speaking. This may of course also be due to their correlations with other regressors, so that their effect on credit spreads becomes hard to disentangle. Liquidity matters mainly for short-term bonds, which is sensible and in line with for instance Janosi, Jarrow and Yildirim (2002) and Driessen (2003). Finally, the R^2 increases by 6 to 14% for long-maturity bonds and by 5 to 12% for short-term corporate debt.

3.2 Year dummies

As a next robustness check, we introduce year dummies in the regression with all other controls of Table 3. Two conclusions can be drawn from this exercise.

First, individual implied volatilities and implied skews pick up more than just time-variation in credit spreads, both for short- and long-maturity bonds. This can be seen in Table 4, since year dummies have very little impact on the effect of these variables. If anything, the coefficients become actually slightly larger and more significant. The coefficients on the S&P-based measures however, as well as on some of the control variables (e.g. the market return), shrink substantially with the inclusion of year dummies, since they only pick up time-series variation. Second, the fact that the R^2 is essentially unchanged means that our explanatory variables already account for all low-frequency time-variation.

3.3 Firm Fixed Effects

To understand how much cross-sectional variation in credit spreads is left unaccounted for, we augment the regression to include firm dummies or firm fixed effects. Unlike year dummies, issuer fixed effects do change the results somewhat. The coefficient on individual implied-volatility increases from 0.78 to 1.27 for short-maturity bonds and drops from 2.32 to 1.68 for long-term bonds. In both cases the coefficients remain very statistically significant. The biggest change can be observed for the coefficient on the individual option-implied skew for long-term bonds, which increases from 0.03 to 0.09 and now has a t-statistic of 10. This suggests that, for long-term bonds, the implied-volatility skew variable is more closely related to individual time-series variation in credit spreads, and less related to cross-sectional variation (or at least with a smaller coefficient): introducing firm fixed effects allows the variables to 'focus' on individual time-series variation, since the firm dummies can take care of the cross-sectional variation. When going back to the simple descriptive statistics studied in the previous section (Table 1 and Figure 4), this is not too surprising, since the cross-sectional relation between option-implied skews and credit spreads was considerably less strong.

Another important finding in Table 4 is that the R^2 goes up substantially when issuer fixed effects are introduced. The explanatory variables explain almost half the variation in credit spreads on short-maturity bonds and more than two thirds of the variation for long-term bonds. Keeping in mind that we impose panel regressions, these numbers are quite remarkable. At the same time, the fact that the R^2 was 29% (short-maturity) and 49% (long-maturity) without fixed effects suggests that individual options do not exhaustively explain the cross-section of credit spreads, even though they are a very important determinants. Other issuer-specific factors seem to be reflected in credit spreads.

3.4 Option-Market Liquidity

As a final extension, we now consider the liquidity of the market for individual options. Although several articles have analyzed the liquidity of corporate bond markets, a study of the impact of option-market liquidity on credit spreads is lacking. A first motivation for considering option-market liquidity is that it may have an effect on credit spreads beyond and in addition to the influence of implied volatilities and option-implied smirks. This may happen because of a liquidity-spillover effect: some corporate bonds may be rather illiquid and investors may require an additional premium as compensation for this illiquidity. If the liquidity of a corporate bond of an issuer is correlated with the liquidity of its traded options, then an issuer-specific measure of option-market liquidity should matter for the credit spreads of that issuer. This relation between option-market illiquidity and credit spreads may in fact reflect hedging activities along the line of what we investigate in the final section of the paper. Issuer-specific credit risk as present in corporate bonds can to some extent be hedged by trading in individual options of that issuer. If these options are particularly illiquid, then hedging becomes more difficult and more costly. This cost may manifest itself in the discount at which the corporate bonds are trading, i.e. in the credit spread.

Secondly, this extension can be viewed as introducing an additional control variable. It is possible that part of the impact of the individual implied volatility and implied-volatility skew on credit spreads is driven by the fact that these two variables are correlated with the (il)liquidity of the options market (to the extent that illiquid options carry a liquidity premium). Then, if the liquidity of options markets matters for credit spreads, the coefficients on implied volatility and implied-volatility skew should shrink towards zero when a direct proxy for option-market liquidity is added to the regression.

As a firm-specific measure of the liquidity of its traded options we use the bid-ask spreads on both out-of-the-money and at-the-money options. Table 5 shows the pooled regression results when adding these measures to all other controls. Two findings are noteworthy. First, the coefficients on implied volatility and on the implied-volatility skew change slightly for short-maturity bonds, but remain identical for long-maturity bonds. This highlights once again the strong robustness of our results for long-maturity bonds. Second, the coefficients on the option-liquidity proxies are rather large and significant for short-maturity bonds and essentially zero in the last column (long-maturity bonds). This is quite sensible, since previous articles have shown that the liquidity spread is largest for short-term bonds (e.g., Janosi, Jarrow, and Yildirim (2002) and Driessen (2003)). Also, for short-maturity bonds, it is the effect of the bid-ask spread on out-of-the-money options that is most precisely estimated (t-statistic of 11.6), which is also reasonable since out-of-the-money options tend to be more illiquid than at-the-money ones. The fact that there is some evidence of a liquidity-spillover effect for short-maturity bonds, but not for long-maturity bonds, can also be seen from the R^2 of the regressions: firm-specific option-liquidity proxies increase the R^2 by 1 percentage point for short-term bonds, while there is no gain in explanatory power for long-maturity bonds.

4 Incorporating Credit Ratings

Credit ratings have been shown to have explanatory power for credit spreads, even when controling for economic determinants of spreads (e.g. Campbell and Taksler (2003)). We therefore include these ratings along with the other variables considered so far. We then proceed to study the interaction between credit ratings and our measures of the option-market assessment of the volatility and jump risk of a firm, in order to analyze whether these measures matter more for bonds of issuers that are closer to default, as would be predicted by a structural firm-value model. Finally, we investigate to what extent option markets incorporate credit information more quickly or differently than do rating agencies, by documenting the dynamics of implied volatility and the implied-volatility skew around rating migrations.

4.1 Credit Ratings as a Control Variable

Standard and Poor's classify issuing firms into 26 different categories based on their risk of default. Even with our reasonably large sample, it is not meaningful to distinguish all 26 distinct groups. We therefore aggregate up the different ratings into 5 groups: AAA, AA, A, BBB, and finally BB and lower.

We first regress credit spreads for short- and long-maturity bonds on a constant and four rating dummies (Table 6). The rating dummies are highly statistically significant and have the expected sign: poorly rated bonds on average have higher credit spreads.

The explanatory power of these regressions is quite limited however, with an R^2 of 9% for short-maturity bonds and 17% for long-maturity bonds. Interestingly, this is substantially less than when regressing credit spreads on the measures of volatility and of jump risk implied by individual option markets. In Table 2 we found that implied volatility and the option skew alone explain 14% of the variation in short-maturity credit spreads and 32% for long-maturity bonds, which is respectively 5 and 15 percentage points higher.

The arguments made by Campbell and Taksler (2003) to explain the relatively poor results with ratings apply here a fortiori. First, ratings are updated slowly and gradually, while our measures exhibit much more high-frequency variation. In addition, our option-based measures are forward-looking in nature. Second, credit ratings have less cross-sectional explanatory power by construction, due to the discreteness of the rating categories (which we exacerbated by further aggregation into just 5 categories): they cannot explain differences in credit spreads for bonds with the same rating, unlike our issuer-specific variables.

In spite of their rather limited explanatory power, the rating dummies remain statistically singificant when all other control variables are added, especially the ones for the categories closest to default. The individual implied-volatility measure continues to be highly statistically and economically significant. In fact, the point estimate in the regression for short-maturity bonds increases slightly. The implied-volatility skew variable on the other hand shrinks when controling for ratings. Even though the coefficient remains statistically significant, its economic significance becomes quite small: a 1-standard-deviation weekly shock to the option-implied skew changes credit spreads by around 1 basis point. It seems that the skew variable, at least to some extent, picks up information that is also conveyed by credit ratings. Since we interpret the skew as a measure of the option market's assessment of the likelihood and severity of a downward jump in firm value, this is perhaps not too surprising, since ratings are designed to carry essentially this same information. Also, the smaller coefficients on the skew variable may well be due to the large number of regressors and control variables included, so that disentangling the different effects becomes difficult. Nonetheless, the finding does suggest that the skew variable is somewhat less powerful and less robust than the level of the option-implied volatility.

Finally, it can be seen that credit ratings are quite complementary to the fundamental economic variables: the R^2 increases from 29.2% to 32.7% for credit spreads on short-maturity bonds and from 49% to 56% for long-term bonds, suggesting that ratings do convey additional information not already present in the economic variables we consider.

4.2 Interaction Terms

So far, we have found empirical support for the prediction of an extended structural firm-value model (augmented to allow for stochastic volatility and jumps) that credit spreads on corporate bonds are positively related to measures of the volatility of the firm value of the issuer and of its jump risk. A further and more precise prediction is that the sensitivity of the credit spreads to volatility and jump risk increases as the firm gets closer to the default boundary. To test whether the impact of volatility and of jump risk on credit spreads is indeed higher for junk-bond issuers than for investment-grade firms, we now interact the credit rating with our option-implied measures of the volatility of the firm value and of the jump risk of the issuer. Since the historical volatility and skewness of the return distribution are potential alternative measures of volatility and of jump risk, we also interact the credit ratings with these historical proxies. Because some of the rating categories contain too few bonds, we pool the data for this purpose into 3 categories: AAA to A-, BBB+ to BBB-, and BB+ and lower.

The first regressions reported in Table 7 replicate the basic regressions of Table 2, where spreads are explained by individual implied volatility and the implied-volatility skew only (with a constant), but now allow for interaction with the credit rating, for short- and long-maturity bonds respectively. In general the sensitivity of credit spreads to implied volatilities increases substantially and significantly as the credit rating deteriorates, especially for long-maturity bonds. The reported coefficients are additive, so that (for instance) a BB long-maturity issuer faces a 6.3 total coefficient on its implied volatility, which is substantially higher than the impact for investment-grade issuers in category 1 (a coefficient of 3.52), in line with the prediction of the model. This means that a one-standard-deviation weekly shock in implied volatility changes the credit spread by 16 basis points for an investment-grade issuer and by 29 basis points for a BB firm (versus 20 basis points in the benchmark regression without credit-rating interaction terms). The results of the interaction between the implied-volatility skew variable and credit ratings are also intuitive: highly-rated bonds have a small coefficient on this option-implied measure of jump risk, but category 2 issuers face a substantially larger impact. The incremental effect for short-maturity issuers in category 3 tends to be insignificant because of the small number of observations in that cell. Allowing for interaction increases the fit of the regression substantially: from 14% to 22% for short-mat and from 32% to 42% for long-term bonds.

When adding the firm-specific historical and economy-wide measures of volatility and jump risk, along with the other controls, the same results obtain. For the option-implied skew measure we now even find that highly-rated bonds have no (for short-maturity) or even a slightly negative (for long-maturity) but economically small coefficient on this proxy for jump risk, while category 2 and 3 have remarkably large sensitivities. This pattern seems also present for the historical measures, but it is less pronounced. However, it is noteworthy that the effect of historical skewness finally becomes negative, as it should be, for categories 2 and 3 of the long-maturity bonds. With the credit rating interaction, the R^2 grows to 34% for short-maturity bonds (from 29%) and to 58% for long-term bonds (from 49%).

4.3 Implied Volatility Patterns around Rating Migrations

In order to shed more light on the relationship between credit markets and individual option markets, and in particular to understand to what extent individual stock-option prices carry information that is relevant for credit risk, we now study the dynamics of implied volatilities of individual options during periods of rating migrations. A number of papers have analyzed the impact of changes in credit ratings on stock and bond prices. The effect of rating migrations on option prices however has not been documented before. If prices of options already incorporate the information that triggers the change in credit rating, we would expect to observe different dynamics of option prices for migrating firms than for non-migrating firms before the announcement. If however the change in credit rating represents new information, not previously reflected in option prices, the main effect on the dynamics of option prices would be expected after the announcement. A detailed study could try to test whether option prices can be used to predict rating changes with a probit analysis or could use option information to model the dynamics of the migration probability matrix (extending for instance Kavvathas (2000)). Since our sample only contains a limited number of rating changes[3], we instead conduct an informal and purely exploratory analysis. We report the patterns in the implied volatilities of firms that migrate and compare with a control group of all non-migrating firms during the same time-period in order to correct for general trends. We distinguish between upgrades and downgrades, because the work on stock and bond price reactions to rating changes has found that the results depend on whether the change concerns an upward or a downward migration (see for instance Hand, Holthausen and Leftwich (1992), Hull, Predescu and White (2003) and Katz (1974)).

Figure 5 shows the average implied volatility for 4 groups of firms, computed as follows. For each downgraded firm, we first find the date of the rating change and the implied volatility of its traded at-the-money options during the 32 weeks around the event date. We then calculate the average implied volatility across all non-downgraded firms (the control group) over the same 32 weeks. Having done this for each migration event, we then average across all downgrades, to obtain the average implied volatility for downgraded firms and for the control group, as plotted in Figure 5. The same is done for upgrades, where the control group consists of all non-upgraded issuers.

A number of interesting results emerge in Figure 5. First, it is clear that downgrades tend to occur during volatile times, while upgrades happen in periods with lower implied volatilities, since the controls groups exhibit on average higher implied volatilities around downgrade events than around upward migrations. Second, the issuers that are downgraded have substantially higher implied volatilities. Focusing now on the time pattern of implied volatilities, we find for downgraded issuers that the option market's assessment of the volatility of firm value rises during the months preceding the migration and peaks 2 to 3 weeks before the rating change announcement. This suggests that option markets anticipate to a large extent the downgrade, or at least, that most of the information triggering the

[3]The number of observations declines further if we insist on having data on implied-volatility skews of the migrating firms. For this reason, we are forced to focus on patterns in implied volatilities only.

migration is already reflected in the option prices. Subsequent to the downgrade, the implied volatility of the affected firms tends to decrease, but only very gradually. The implied volatility remains well above the level for non-downgraded firms, even 16 weeks after the event. Upgraded firms on the other hand are less dramatically different from other firms. Although the implied volatility for issuers that are upgraded is typically lower than for all other firms, the difference is quite small. Furthermore, firms that experience upward migrations see essentially no change in their implied volatilities during the months preceding the rating change. Interestingly, their implied volatility fully reverts back to the mean for non-upgraded firms during the month subsequent to the rating change.

These findings are consistent with the results in the literature that downgrades have a larger effect on credit spreads and equity prices than do upgrades. As a caveat however, it should be pointed out that the lack of striking results for upgrades may also simply be due to the small number of upward migrations in the sample (34 upgrades versus 73 downgrades).

As a next step, we classify firms into the 3 broad rating categories of the previous subsection (AAA to A-, BBB+ to BBB-, and BB+ and lower), based on their pre-migration rating. However, we continue to identify all rating changes: for instance, a firm that is downgraded from AA to AA- is included in the group of downgrades (of rating category 1), even though it stays within the same category. The various control groups are also constructed using firms belonging to the respective rating category. A first finding is that within each pre-migration rating category, downgraded firms have substantially higher implied-volatilities. This lends further support to our claim that firm-specific option-implied volatility matters for credit risk, since within a given rating category the implied-volatility essentially provides information (on average) about subsequent downgrades, i.e. about increases in credit risk as perceived by the rating agency. Comparing Figures 6, 7 and 8, a striking result is that the strongest quantitative effect occurs for downgrades in categories 2 and 3. Firms that are already poorly rated and that experience a further downgrade, have extremely high implied volatilities that (on average) reach twice (category 2) or even three times (category 3) the implied volatility of firms in the corresponding control group. It is also noteworthy that the implied volatility of downgraded issuers in the weakest group (category 3) continues to increase after the downgrade and peaks roughly half a month later. As in Figure 5, the difference between upgraded firms and other issuers is very small.

5 Hedging Credit Risk

To further extend the analysis we now study whether individual options can be used to hedge the credit risk in corporate bond returns. This highlights a potentially important advantage of using options rather than historical proxies for measuring the volatility of the firm value and its jump risk, as we have proposed in this paper: unlike the historical proxies, options are traded assets and one can therefore think about implementing trading strategies that are designed to exploit the information these assets contain about credit risk for hedging purposes. Furthermore, it has been argued that hedging credit risk in corporate bonds is extremely difficult (Clare, Ioannides and Skinner (2000)).

In line with the rest of our analysis, the hedging strategies will be regression-based. We regress four-weekly returns on individual corporate bonds on the returns of a number of natural hedging instruments: the equity of the issuer, a short- and long-maturity government bond and finally two individual options, an at-the-money straddle and an out-of-the-money put. The straddle can be thought of as an instrument that loads heavily on the volatility risk of the firm value and the put can be interpreted as mainly capturing the jump risk. We run the regressions for each bond separately using a time-series of returns and retaining only the bonds with at least 25 observations. The results are reported in the form of cross-sectional averages of the coefficients and $R^{2\prime}s$, both for short-maturity and for long-maturity bonds. All standard errors and t-statistics are computed using the methodology of Fama and MacBeth (1973), acknowledging that the coefficients are obtained as cross-sectional averages of time-series regressions.

Table 8 reports the results for short- and long-maturity bonds, starting with a regression of corporate bond returns on equity and government bond returns. We then progressively add the ATM straddle, OTM put and finally both instruments simultaneously. For the first regression, the coefficient on the firm's equity is negative on average, but insignificant. The hedge ratios for short-maturity and long-maturity government bonds are positive and highly significant. Not surprisingly, the largest coefficient obtains for short-maturity Treasury bonds. The average R^2 is 22%. When adding either the straddle (regression 2) or the OTM put (regression 3), the R^2 increases by 4 to 5 percentage points. The coefficients on the derivatives are small in absolute value, which reflects the leverage implicit in the derivatives. As expected, both coefficients are negative on average, since corporate bond returns should load negatively on volatility and jump risk. When adding both derivatives simultaneously, the hedge ratios on government bonds and derivatives remain significantly positive and negative, respectively. Equity is insignificant. Most importantly, the availability of both derivatives increases the R^2 from 22% to 31%, suggesting that individual options can indeed be used successfully to hedge credit risk. However, a large part of the variation in corporate bond returns remains unhedged. Although individual options do help to hedge credit risk, the hedge is still quite imperfect.

For long-maturity bonds, equity and bonds alone hedge 36% of the variation in corporate bond returns. The coefficient on equity is now positive on average, but still insignificant. The hedge ratio for short-maturity bonds becomes negative, while the coefficient on long-maturity Treasuries is very large and positive. Although the straddle coefficient is significantly negative, it only helps the effectiveness of the hedge to a very limited extent. The OTM put on the other hand is insignificant, but raises the R^2 by 6 percentage points. As for short-maturity bonds, adding both derivatives to the menu of securities available for hedging purposes improves the hedge quite a bit more and the R^2 becomes 42.6%. However, the hedge ratios for both derivatives are imprecisely estimated and insignificant, suggesting that hedge ratios vary substantially across bonds.

6 Conclusion

We have demonstrated empirically that prices of traded individual equity-options contain important information about credit risk. In particular, volatilities implied by at-the-money and out-of-the-money options are very useful proxies for two fundamental theoretical determinants of credit spreads, namely volatility and jump risk. The effect of at-the-money implied volatility on credit spreads is economically and statistically very significant, and robust to a number of extensions. Together with option-implied skews, but without any other explanatory variables, 32% of the variation in long-term credit spreads can be explained, even though the pooled regression imposes constant coefficients over time and across issuers. This is almost twice the explanatory power that can be achieved by credit ratings. Importantly, interacting our measures of volatility and jump risk with credit ratings, we find the economically meaningful result that the credit spreads of poorly rated bonds are substantially more sensitive to these determinants.

Further empirical evidence for the interaction between the individual equity-option market and the market for corporate bonds is provided by the following findings. Short-maturity credit spreads are significantly affected by measures of firm-specific option-market liquidity, suggesting the existence of a liquidity-spillover effect. Furthermore, option-implied volatilities anticipate downward credit rating migrations in striking way, especially for issuers that already have a low credit rating. Finally, we present evidence that traded individual options can help to hedge the credit risk in corporate bonds.

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Table 1: Summary Statistics

Variable	Mean	TS Std. Dev. Levels	TS Std. Dev. Change	CS Std. Dev.	Correlation TS	Correlation CS
Short-Mat. Credit Spread	1.031	0.352	0.097	0.663		
Long-Mat. Credit Spread	1.109	0.402	0.083	0.583		
Ind. Implied Volatility	0.348	0.098	0.046	0.062	0.752	0.400
Ind. Skew	0.314	0.330	0.361	0.129	0.258	-0.122
Ind. Historical Volatility	0.325	0.119	0.012	0.090	0.862	0.289
Ind. Historical Skewness	0.121	0.662	0.184	0.267	-0.566	0.059
S&P Implied Volatility	0.205	0.050	0.029		0.419	
S&P Skew	0.773	0.179	0.148		-0.220	
S&P Historical Volatility	0.183	0.053	0.004		0.822	
S&P Historical Skewness	-0.302	0.505	0.136		0.714	

Table 2: Benchmark Regressions for Short- and Long-Maturity Bonds

	Short-Maturity					Long-Maturity				
Regression	1	2	3	4	5	1	2	3	4	5
Ind. IV	2.52693				1.02896	4.4642				2.84149
	(26.75)				(11.14)	(67.02)				(30.05)
Ind. Skew	0.24294				0.14862	0.25738				0.17201
	(23.46)				(13.85)	(20.55)				(14.30)
Ind. Hist. Vol.		2.77448			1.14919		4.32151			0.39753
		(30.66)			(10.12)		(72.91)			(4.49)
Ind. Hist. Skewness		0.12017			0.08648		0.03717			0.03822
		(11.11)			(8.26)		(2.50)			(3.15)
S&P IV			1.93699		-0.2461			3.44365		-1.7307
			(27.56)		(-3.10)			(35.60)		(-17.8)
S&P Skew			-0.0328		0.37712			-0.4868		0.0581
			(-1.84)		(22.23)			(-18.96)		(2.58)
S&P Hist. Vol.				4.30727	2.05109				7.37849	4.565
				(72.16)	(14.57)				(87.70)	(35.78)
S&P Hist. Skewness				0.54223	0.61076				0.84107	0.63302
				(30.11)	(36.78)				(37.99)	(25.33)
Constant	0.00137	-0.0471	0.54379	-0.4162	-1.0514	-0.3627	-0.2525	0.90762	-0.9381	-1.0534
	(0.043)	(-1.584)	(31.827)	(-21.854)	(-32.995)	(-16.29)	(-12.70)	(33.36)	(-42.58)	(-31.58)
R^2	0.143	0.172	0.022	0.178	0.245	0.322	0.311	0.056	0.323	0.420
Adjusted R^2	0.143	0.172	0.022	0.178	0.245	0.322	0.311	0.056	0.323	0.420
Number of bonds	334	334	334	334	334	189	189	189	189	189
Number of firms	66	66	66	66	66	56	56	56	56	56
Total number of bond-weeks	34419	34419	34419	34419	34419	24261	24261	24261	24261	24261

Table 3: Regressions with Additional Control Variables

Regression	Short-Maturity		Long-Maturity	
	1	5	1	5
Ind. IV	1.31551	0.78199	2.25859	2.31634
	(11.82)	(8.76)	(29.86)	(26.72)
Ind. Skew	0.08253	0.06976	0.03375	0.0335
	(8.99)	(7.84)	(3.22)	(3.17)
Ind. Hist. Vol.		1.605		0.55922
		(10.97)		(5.92)
Ind. Hist. Skewness		0.13801		0.17172
		(10.87)		(13.45)
S&P IV		-0.8215		-2.0516
		(-10.73)		(-21.19)
S&P Skew		0.14399		-0.2002
		(8.09)		(-8.77)
S&P Hist. Vol.		-1.5835		-0.0753
		(-7.83)		(-0.37)
S&P Hist. Skewness		0.06455		-0.1578
		(2.73)		(-5.02)
Market Return	0.0324	-0.0871	-0.3033	-0.4146
	(0.97)	(-2.53)	(-7.13)	(-9.96)
Firm Stock Return	-0.0253	-0.0522	-0.118	-0.1782
	(-1.56)	(-3.03)	(-8.60)	(-11.92)
Two-year Yield	-0.159	-0.177	-0.1781	-0.1967
	(-12.47)	(-12.22)	(-11.43)	(-10.00)
Ten-year Yield	-0.3117	-0.2677	-0.4512	-0.4371
	(-9.69)	(-7.78)	(-14.56)	(-11.41)
BAA Rate	0.32899	0.27103	0.37952	0.33461
	(14.51)	(12.29)	(13.75)	(12.28)

Table 3 (continued): Regressions with Additional Control Variables

Regression	Short-Maturity		Long-Maturity	
	1	5	1	5
Swap Rate	0.17201	0.11835	0.32223	0.28939
	(3.28)	(2.17)	(5.63)	(4.89)
Liquidity	0.12883	0.15166	0.00188	0.09612
	(8.76)	(10.17)	(0.10)	(4.76)
Constant	-0.4768	0.08559	-0.7874	0.21774
	(-7.58)	(0.98)	(-11.26)	(2.01)
R^2	0.265	0.292	0.468	0.487
Adjusted R^2	0.265	0.292	0.468	0.486
Number of bonds	334	334	189	189
Number of firms	66	66	56	56
Total number of bond-weeks	34419	34419	24261	24261

Table 4: Regressions with Year Dummies and Firm Fixed Effects

Regression	Short-Maturity		Long-Maturity	
	Year Dummies	Fixed Effects	Year Dummies	Fixed Effects
Ind. IV	0.80141	1.2744	2.39065	1.67692
	(8.75)	(10.69)	(27.39)	(20.52)
Ind. Skew	0.07192	0.06298	0.03732	0.09067
	(8.22)	(6.86)	(3.52)	(10.18)
Ind. Hist. Vol.	1.59329	1.80913	0.4641	1.0069
	(11.13)	(8.02)	(4.93)	(9.73)
Ind. Hist. Skewness	0.13849	0.11699	0.17756	0.07405
	(10.58)	(8.29)	(13.55)	(6.96)
S&P IV	-0.8525	-0.9524	-1.9425	-1.5182
	(-9.45)	(-12.21)	(-16.641)	(-19.445)
S&P Skew	0.10023	0.17348	-0.1866	-0.1518
	(5.93)	(11.31)	(-7.95)	(-8.50)
S&P Hist. Vol.	-1.3306	-2.0465	0.2485	-0.134
	(-4.90)	(-8.27)	(0.86)	(-0.79)
S&P Hist. Skewness	0.03706	0.0199	-0.2121	-0.162
	(1.13)	(1.04)	(-4.83)	(-6.48)
Market Return	-0.0416	-0.1117	-0.1281	-0.3789
	(-0.95)	(-3.59)	(-2.35)	(-11.80)
Firm Stock Return	-0.0578	0.06986	-0.1981	-0.0715
	(-3.04)	(4.56)	(-12.77)	(-5.48)
Two-year Yield	-0.2359	-0.1786	-0.1485	-0.1893
	(-9.16)	(-15.71)	(-5.35)	(-13.34)
Ten-year Yield	-0.3111	-0.2591	-0.6993	-0.4963
	(-8.21)	(-6.50)	(-13.88)	(-15.51)
BAA Rate	0.36063	0.2088	0.4034	0.37347
	(14.62)	(11.33)	(12.51)	(18.19)

Table 4 (continued): Regressions with Year Dummies and Firm Fixed Effects

Regression	Short-Maturity		Long-Maturity	
	Year Dummies	Fixed Effects	Year Dummies	Fixed Effects
Swap Rate	0.16722	0.15033	0.42813	0.31678
	(3.16)	(2.96)	(6.98)	(6.60)
Liquidity	0.0906	0.14618	0.05894	0.10765
	(4.73)	(11.03)	(2.44)	(6.53)
Constant	-0.4787	0.22081	0.03537	-0.0644
	(-3.50)	(2.38)	(0.21)	(-0.76)
R^2	0.294	0.485	0.491	0.673
Adjusted R^2	0.294	0.484	0.490	0.672
Number of bonds	334	334	189	189
Number of firms	66	66	56	56
Total number of bond-weeks	34419	34419	24261	24261

Table 5: The Effect of Option-Market Liquidity

Regression	Short-Maturity	Long-Maturity
Ind. IV	0.94654	2.32476
	(9.70)	(25.57)
Ind. Skew	0.05502	0.03779
	(5.77)	(3.28)
Ind. Hist. Vol.	1.57	0.53768
	(10.56)	(5.59)
Ind. Hist. Skewness	0.12859	0.17327
	(10.11)	(13.43)
S&P IV	-0.8044	-2.0711
	(-10.59)	(-21.13)
S&P Skew	0.15001	-0.1835
	(8.46)	(-8.04)
S&P Hist. Vol.	-1.5222	-0.031
	(-7.48)	(-0.15)
S&P Hist. Skewness	0.07407	-0.1374
	(3.12)	(-4.37)
Market Return	-0.0859	-0.4234
	(-2.50)	(-10.02)
Firm Stock Return	-0.0334	-0.1801
	(-1.96)	(-12.00)
Two-year Yield	-0.1741	-0.196
	(-11.75)	(-10.74)
Ten-year Yield	-0.2862	-0.4446
	(-8.20)	(-11.42)
BAA Rate	0.28422	0.32628
	(12.60)	(11.68)

Table 5 (continued): The Effect of Option-Market Liquidity

Regression	Short-Maturity	Long-Maturity
Swap Rate	0.12264	0.3074
	(2.19)	(5.06)
Bond Liquidity	0.13714	0.09323
	(9.08)	(4.57)
OTM bid-ask spread	0.15433	-0.0082
	(11.61)	(-0.58)
ATM bid-ask spread	0.23011	-0.0735
	(2.67)	(-0.72)
Constant	-0.0677	0.18554
	(-0.77)	(1.70)
R^2	0.303	0.489
Adjusted R^2	0.302	0.489
Number of bonds	333	189
Number of firms	66	56
Total number of bond-weeks	33692	23885

Table 6: Controlling for Credit Ratings

Regression	Short-Maturity 1	Short-Maturity 5	Long-Maturity 1	Long-Maturity 5
Ind. IV		0.86383		2.02008
		(8.61)		(24.18)
Ind. Skew		0.0434		0.02399
		(4.44)		(2.24)
Ind. Hist. Vol.		1.23846		0.26613
		(7.23)		(2.77)
Ind. Hist. Skewness		0.1441		0.1168
		(10.71)		(9.17)
S&P IV		-0.897		-1.9746
		(-10.89)		(-21.20)
S&P Skew		0.15121		-0.1734
		(7.82)		(-7.89)
S&P Hist. Vol.		-1.2267		0.23646
		(-5.45)		(1.22)
S&P Hist. Skewness		0.10107		-0.0685
		(4.13)		(-2.28)
Constant	0.64059	-0.0997	0.70095	0.05835
	(74.39)	(-1.03)	(54.08)	(0.56)
Rating Group 2	0.27037	0.17017	0.11351	-0.0719
	(23.26)	(11.78)	(7.25)	(-4.68)
Rating Group 3	0.21025	0.13542	0.35623	0.11908
	(23.20)	(10.19)	(25.82)	(9.96)
Rating Group 4	0.70488	0.46315	0.99108	0.53731
	(39.99)	(34.21)	(57.12)	(42.54)
Rating Group 5	0.16142	0.28955	0.73387	0.8056
	(13.56)	(12.55)	(40.37)	(42.41)

Table 6 (continued): Controlling for Credit Ratings

Regression	Short-Maturity		Long-Maturity	
	1	5	1	5
Market Return		-0.1249		-0.4976
		(-3.49)		(-12.53)
Firm Stock Return		-0.0616		-0.144
		(-3.23)		(-9.97)
Two-year Yield		-0.1523		-0.1347
		(-10.43)		(-8.07)
Ten-year Yield		-25.61		-0.5028
		(-6.86)		(-13.73)
BAA Rate		0.24508		0.30458
		(11.24)		(12.14)
Swap Rate		0.11752		0.32493
		(2.03)		(5.74)
Bond Liquidity		0.17088		0.10104
		10.92		(5.15)
R^2	0.093	0.327	0.171	0.560
Adjusted R^2	0.093	0.326	0.170	0.560
Number of bonds	308	308	168	168
Number of firms	61	61	48	48
Total number of bond-weeks	30507	30507	22792	22792

Table 7: Interaction with Credit Ratings

Regression	Short-Maturity 1	Short-Maturity 5	Long-Maturity 1	Long-Maturity 5
Ind. IV	2.17726	0.63667	3.51589	1.4813
	(28.59)	(9.20)	(59.90)	(20.95)
Ind. IV × Rating2	1.29332	0.9425	1.49576	0.77289
	(23.52)	(3.36)	(47.98)	(5.19)
Ind. IV × Rating3	0.08642	-0.1182	1.28488	2.42888
	(2.15)	(-0.42)	(11.59)	(3.42)
Ind. Skew	0.01123	0.00061	0.00796	-0.0022
	(14.75)	(0.87)	(9.11)	(-2.86)
Ind. Skew × Rating2	0.03275	0.02323	0.02641	0.00706
	(14.50)	(10.94)	(12.76)	(3.73)
Ind. Skew × Rating3	-0.0042	0.00127	0.01295	0.01575
	(-1.68)	(0.65)	(2.99)	(4.68)
Ind. Hist. Vol.		1.057		0.062
		(9.58)		(0.78)
Ind. Hist. Vol.×Rating2		0.1626		0.53268
		(0.61)		(3.66)
Ind. Hist. Vol.×Rating3		0.77461		-0.294
		(2.82)		(-0.43)
Ind. Hist. Skewness		0.00564		0.01479
		(7.54)		(20.83)
Ind. Hist. Skewness		0.01038		-0.0172
× Rating2		(4.29)		(-7.72)
Ind. Hist. Skewness		-0.0087		-0.1571
× Rating3		(-0.77)		(-7.22)
S&P IV		-0.6483		-1.9166
		(-8.36)		(-20.64)
S&P Skew		0.01291		-0.0104
		(9.57)		(-6.15)
S&P Hist. Vol.		-1.2054		0.44675
		(-6.39)		(2.42)
S&P Hist. Skewness		0.00989		-0.0017
		(5.44)		(-0.77)

Table 7 (continued): Interaction with Credit Ratings

Regression	Short-Maturity		Long-Maturity	
	1	5	1	5
Market Return		-0.1107		-0.55
		(-3.34)		(-14.40)
Firm Stock Return		-0.0258		-0.146
		(-1.55)		(-10.63)
Two-year Yield		-0.1213		-0.1036
		(-8.30)		(-6.30)
Ten-year Yield		-0.3407		-0.508
		(-10.83)		(-14.79)
BAA Rate		0.25661		0.31375
		(13.08)		(12.94)
Swap Rate		0.16084		0.30043
		(3.18)		(5.55)
Bond Liquidity		0.16364		0.09896
		(11.35)		(5.23)
Constant	0.12672	0.15345	-0.1439	0.24733
	(4.81)	(1.78)	(-7.50)	(2.55)
R^2	0.218	0.344	0.421	0.577
Adjusted R^2	0.218	0.343	0.421	0.577
Number of bonds/firms	334/66	334/66	189/56	189/56
Total number of bond-weeks	34419	34419	24261	24261

Table 8: Hedging Corporate Bonds with Individual Options

	Short-Maturity				Long-Maturity			
Regression	1	2	3	4	1	2	3	4
Firm's stock return	-0.00092	0.00274	-0.00224	-0.00227	0.00171	0.00397	0.00073	0.00785
	(-0.50)	(1.61)	(-0.98)	(-0.83)	(0.59)	(1.22)	(0.16)	(1.55)
Short-Mat. Treasury	0.84285	0.77251	0.95257	0.70875	-0.9429	-0.94577	-1.16313	-1.17198
	(8.73)	(7.44)	(6.94)	(4.86)	(-4.51)	(-5.14)	(-4.98)	(-4.97)
Long-Mat. Treasury	0.14011	0.16475	0.10803	0.18719	1.15978	1.19114	1.29512	1.32367
	(4.86)	(5.35)	(2.67)	(4.55)	(16.43)	(17.56)	(16.82)	(16.56)
Firm's straddle return		-0.00096		-0.00116		-0.00176		-0.00067
		(-2.55)		(-2.10)		(-2.44)		(-0.75)
Firm's OTM put return			-0.00025	-0.00042			-0.0003	-0.00011
			(-1.38)	(-1.94)			(-1.06)	(-0.28)
Constant	-0.00449	-0.0043	-0.00499	-0.0042	-0.00212	-0.00253	-0.00218	-0.00242
	(-14.10)	(-12.23)	(-10.85)	(-8.23)	(-2.96)	(-4.33)	(-2.84)	(-3.05)
R^2	0.219	0.261	0.267	0.309	0.358	0.375	0.414	0.426
Adjusted R^2	0.149	0.158	0.158	0.171	0.311	0.306	0.349	0.338
Number of Bonds	262	243	192	158	148	137	117	107

Figure 1

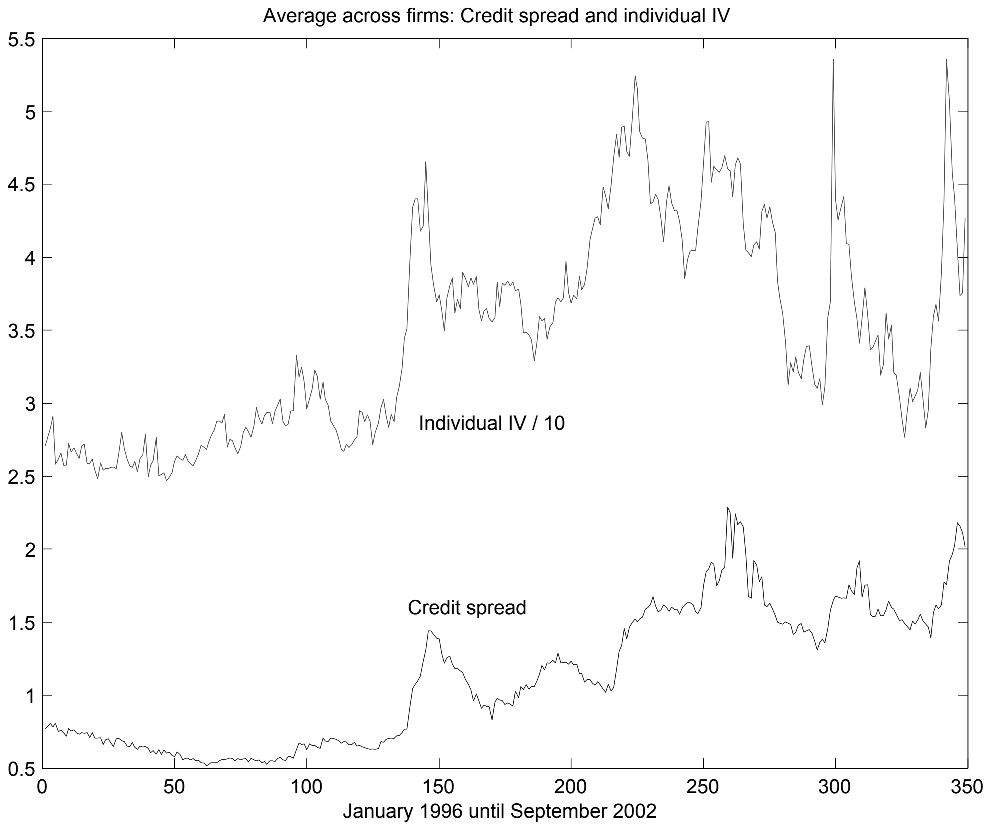

Average across firms: Credit spread and individual IV
5.5
5
4.5
4
3.5
3
2.5
2
1.5
1
0.5
0
50
100
150
200
250
300
350
January 1996 until September 2002
Individual IV / 10
Credit spread

Figure 2

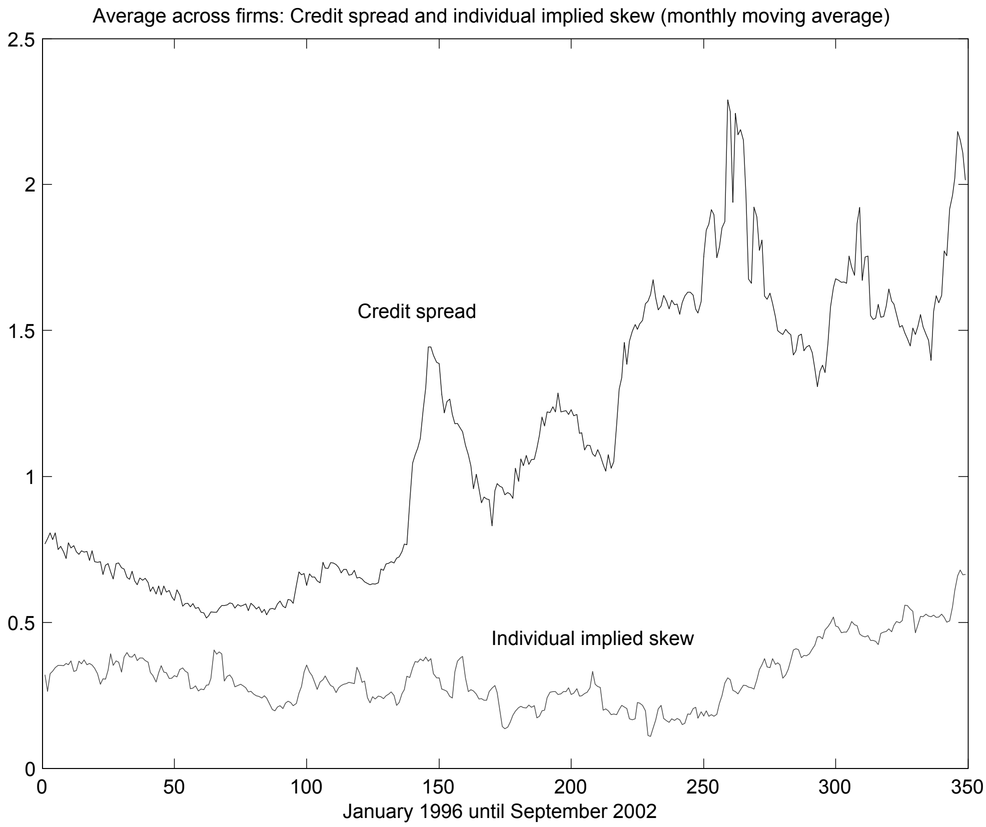

Average across firms: Credit spread and individual implied skew (monthly moving average)
2.5
2
1.5
1
0.5
0
0
50
100
150
200
250
300
350
Credit spread
Individual implied skew
January 1996 until September 2002

Figure 3

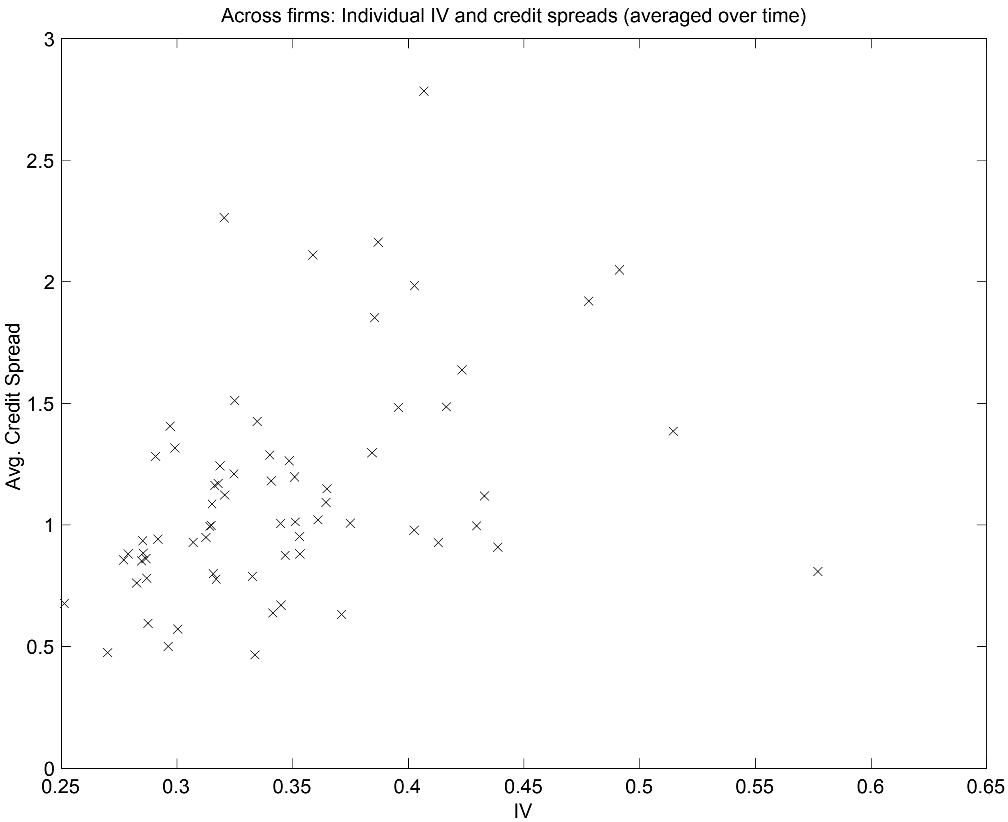
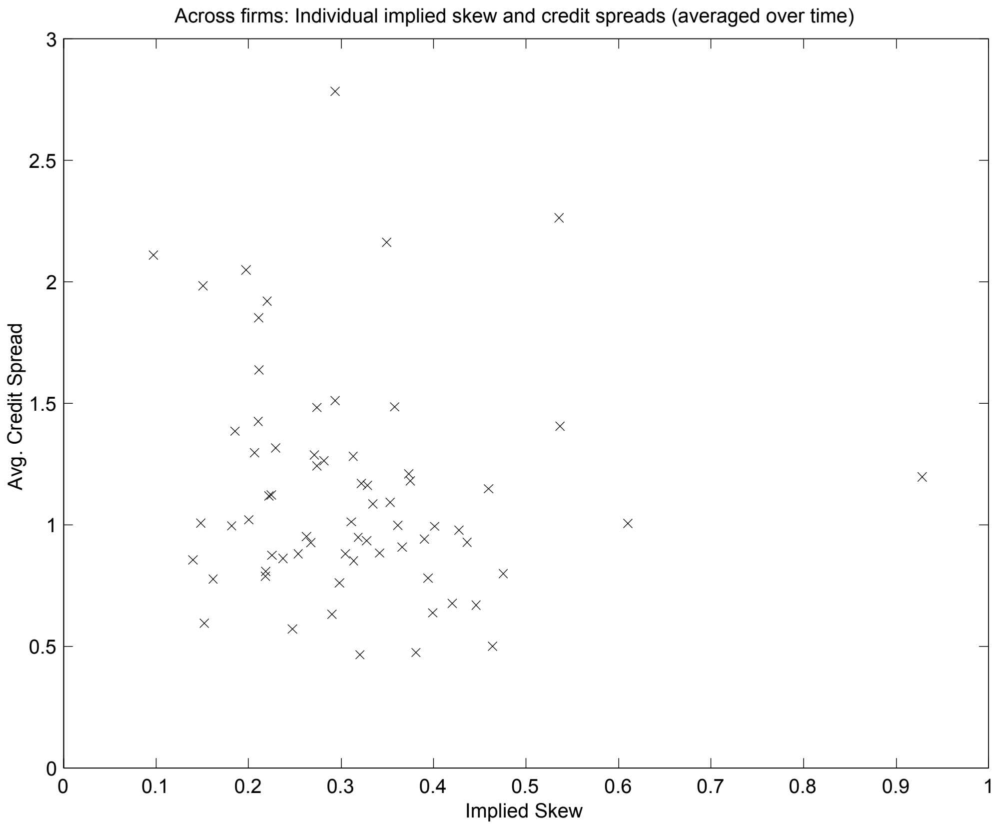

Across firms: Individual IV and credit spreads (averaged over time)
Avg. Credit Spread
IV
0
0.5
1
1.5
2
2.5
3
0.25
0.3
0.35
0.4
0.45
0.5
0.55
0.6
0.65

Figure 4



Across firms: Individual implied skew and credit spreads (averaged over time)
3
2.5
2
1.5
1
0.5
0
Avg. Credit Spread
0
0.1
0.2
0.3
0.4
0.5
0.6
0.7
0.8
0.9
1
Implied Skew

Figure 5

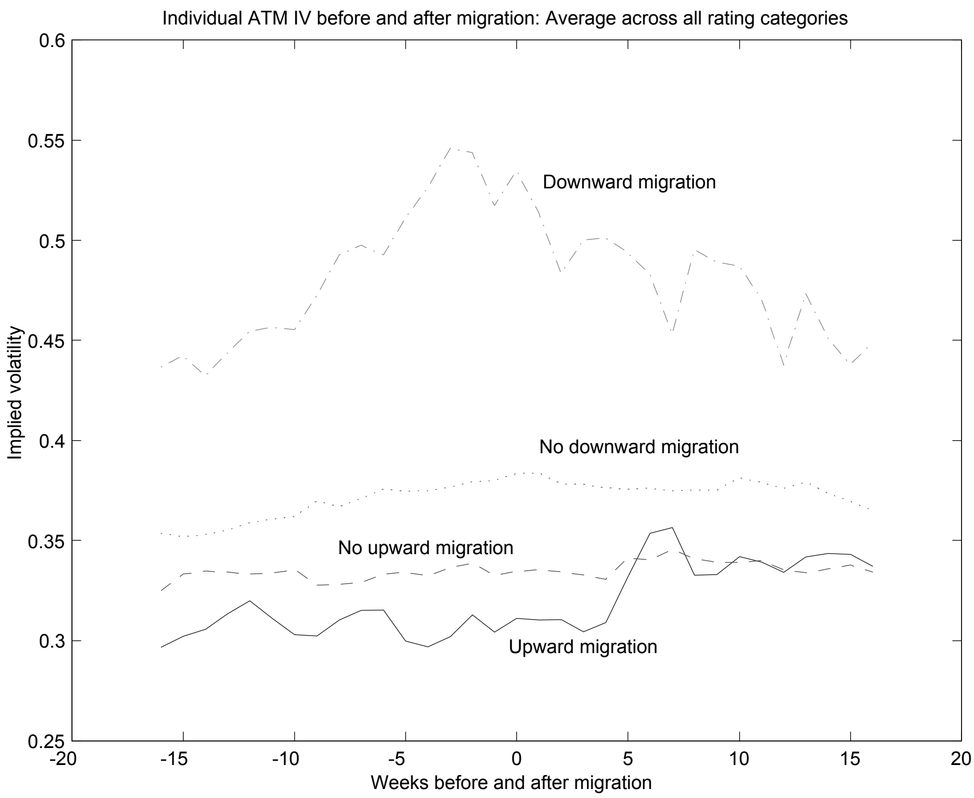
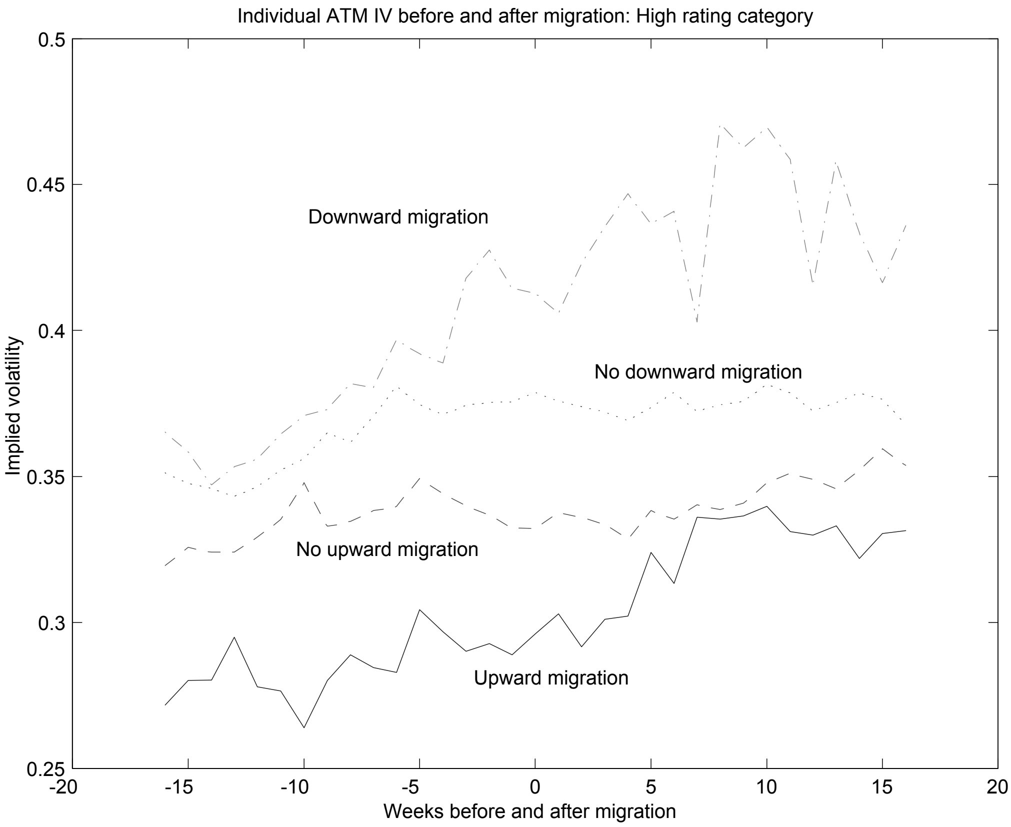
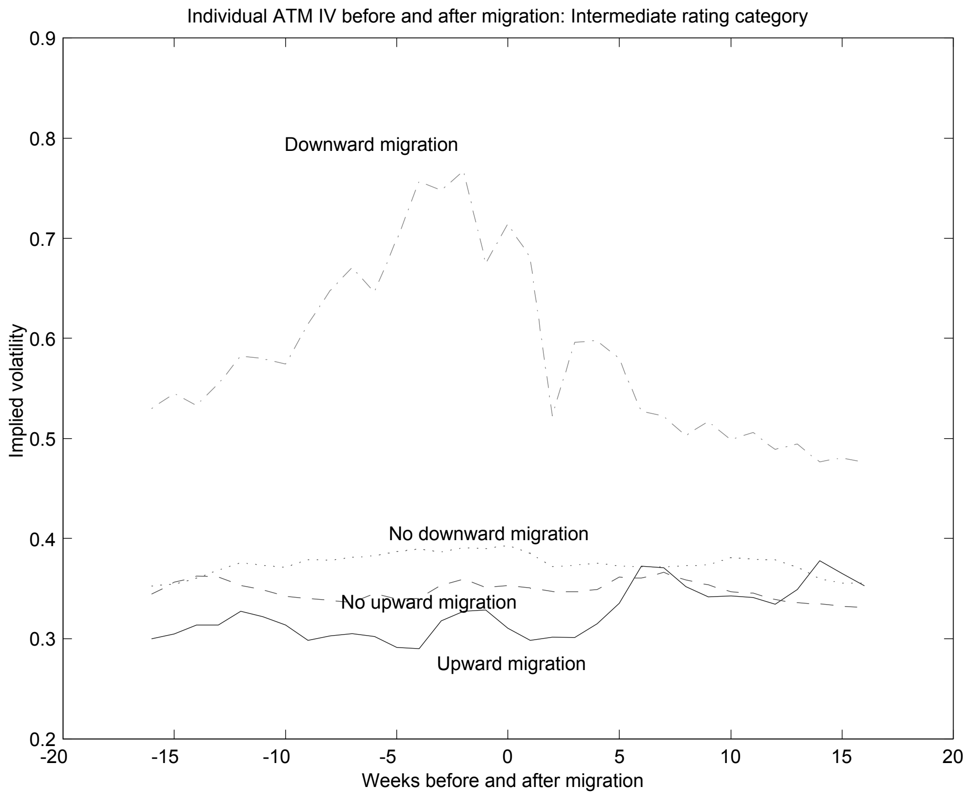

Individual ATM IV before and after migration: Average across all rating categories
Implied volatility
Weeks before and after migration
Downward migration
No downward migration
No upward migration
Upward migration
0.6
0.55
0.5
0.45
0.4
0.35
0.3
0.25
-20
-15
-10
-5
0
5
10
15
20

Figure 6



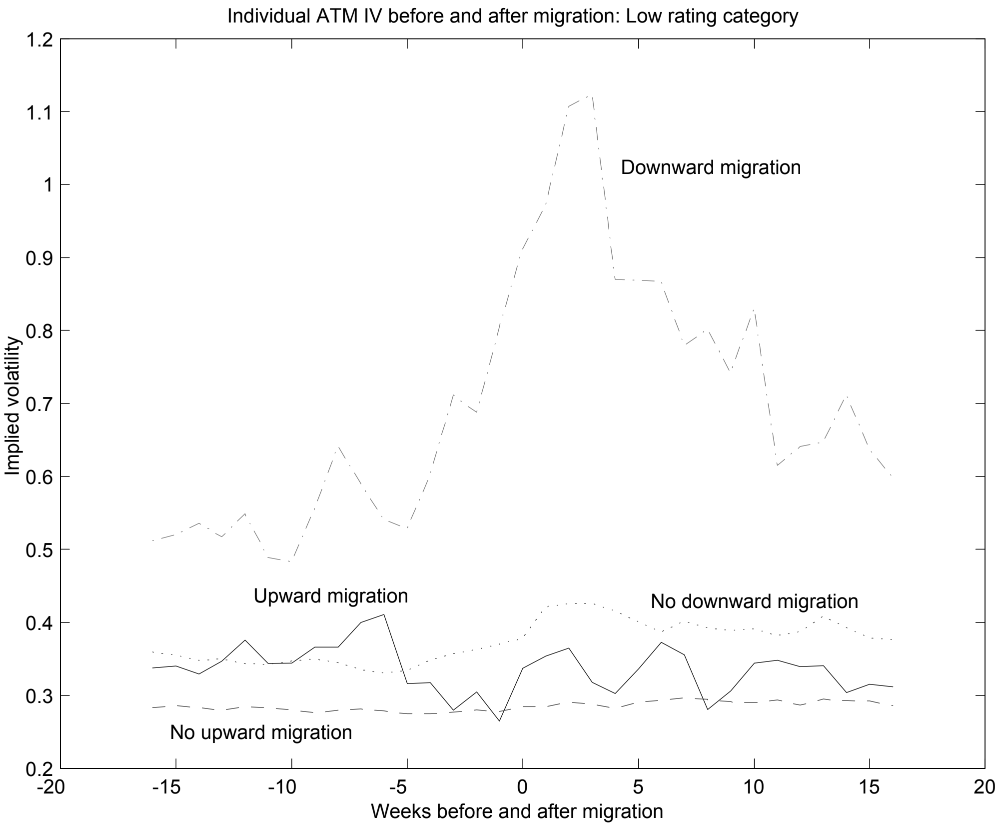
Individual ATM IV before and after migration: High rating category
Downward migration
No downward migration
No upward migration
Upward migration
Implied volatility
0.5
0.45
0.4
0.35
0.3
0.25
-20
-15
-10
-5
0
5
10
15
20
Weeks before and after migration

Figure 7



Individual ATM IV before and after migration: Intermediate rating category
Downward migration
No downward migration
No upward migration
Upward migration
Implied volatility
Weeks before and after migration
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
-20
-15
-10
-5
0
5
10
15
20

Figure 8


Individual ATM IV before and after migration: Low rating category
Implied volatility
Weeks before and after migration
Downward migration
Upward migration
No downward migration
No upward migration
1.2
1.1
1
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
-20
-15
-10
-5
0
5
10
15
20