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Margin regulation and market quality: a microstructure analysis

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Abstract

We find that trading volume increases and market liquidity remains unchanged, while the adverse selection and order-processing cost components of the spread increase and decrease, respectively, after margin levels decline when stocks become margin-eligible. This evidence indicates that the information content of trades has increased, thereby improving market quality. However, no changes were detected after the 1997 regulatory reforms. These results have implications across a broad swath of corporate finance dimensions, including the (1) cost of capital, (2) public vs. private financing decision, (3) form of managerial compensation, (4) type of ownership structure, and (5) degree of shareholder monitoring.

JEL classification: G12; G14; G19
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1. Introduction

The Board of Governors of the Federal Reserve System (hereafter, the "Fed") is empowered to set the percent of a security transaction's value that can be financed by a

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lender such as the initiating broker.[1] For example, if the margin level is set at m%, an investor can borrow up to $(100 - m)$% from the broker when purchasing stock, with the stock serving as collateral for the loan. The effectiveness of margin regulation, in general, and its alleged influence on a stock's volatility, in particular, have been extensively examined with conflicting results.[2]

In this study, we analyze the relation between margin levels and market quality, which we define as the liquidity and the informativeness of stock prices. We investigate links between margin levels and liquidity—namely, spreads, depths, and number of market-makers—and relate these links to a number of corporate finance issues.[3] Seguin (1990) shows that lower margin levels are associated with lower return volatility and larger trading volume. His findings suggest that lower levels of m should lead to some combination of lower spreads, greater depths, and more market-makers, thereby improving market liquidity. According to, for example, Amihud and Mendelson (1986, 2001), greater liquidity should lower the cost of capital for a firm.[4] In contrast, Hardouvelis and Theodossiou (2002) find that lower margin levels are associated with greater return volatility during bull and normal markets. Under their paradigm, lowering the level of m can lead to a deterioration in market liquidity, thereby raising the firm's cost of capital.

We document that lower margin levels lead to a significant increase in daily trading volume, but detect no significant changes in liquidity.[5] Specifically, after considering the impact of certain control variables, we find no significant changes in quoted and effective spreads or in the number of market-makers, but we do find a significant decrease in depths. However, this decrease is economically immaterial. These results indicate that lower margin levels do not change a stock's liquidity. Consequently, the first corporate finance implication of our results is that a lower margin level does not materially affect the firm's cost of capital.

We next investigate possible links between margin levels and the informativeness of stock prices by estimating the components of the spread using the decomposition model of Madhavan et al. (1997). We find that lower margin levels are associated with significant increases in the proportion of the spread attributable to adverse selection. Combining this result with our findings that spreads and depths do not change while trading volume increases, we conclude that the information content of trading and, therefore, the informational environment of the firm and the efficiency of the firm's share price improves when the margin level is lowered.

Because our period of analysis runs from 1993 through 1998 and involves a sample of Nasdaq Small Cap stocks, we investigate the impact of two major market regulatory reforms that were instituted in 1997 on our results. The first reform, the implementation of

[1] Regulation T ("Reg T") covers margin credit granted by broker-dealers; Regulations U and G cover margin credit granted by banks and other lenders, respectively.

[2] Chance (1990) provides a comprehensive review much of the early literature, while Hardouvelis and Theodossiou (2002) provide a review of the more recent literature.

[3] Lee et al. (1993) and Goldstein and Kavajecz (2000) show that changes in both spread and depth are needed to unambiguously infer changes in quoted liquidity.

[4] Also see Brennan and Subrahmanyam (1996), Amihud et al. (1997), Datar et al. (1998), Benveniste et al. (2001), Hasbrouck and Seppi (2001), Pastor and Stambaugh (2002), and Amihud (2002).

[5] Whenever we refer to a change as "significant" (or "insignificant"), we mean statistically significant (or insignificant).

the Order Handling Rules, mandated the public display of limit orders on Nasdaq. The second regulatory reform is the introduction of "teenies" that lowered the minimum tick size from eighths to sixteenths.[6] After examining the entire sample period, we calculate and compare the results from two subsamples—one involving the portion of the sample period before the first of these reforms and the second involving the portion after the second reform was initiated. We find that our aggregate results are driven by the first subsample, because there were no significant changes in liquidity or spread components in the second subsample. Thus, the regulatory reforms appear to have removed any impact of changes in margin levels on market quality.

Finally, while examining the impact of margin regulation on volatility is not our primary objective, we do provide evidence in support of the view that lower margin levels do not significantly increase volatility. Coupled with the liquidity and spread decomposition results, we conclude that lower margin requirements led to improvements in the quality of the market for the affected shares before the 1997 regulatory reforms, but to no change afterwards.

Our conclusion that lower margin levels enhance the informational environment surrounding the firm has numerous and widespread implications across a number of corporate finance paradigms. For example, Subrahmanyam and Titman (1999) argue that the choice between public and private financing depends on the efficiency of the firm's share price. Holmstrom and Tirole (1993, p. 678) argue that the structure of executive compensation depends on "the amount of information contained in the stock price". Finally, a branch of research (e.g., Bolton and von Thadden, 1998; Kahn and Winton, 1998; Maug, 1998; Yu, 2002) argues that there are links between the information content of stock prices, institutional ownership (or ownership concentration), and the degree of monitoring. Although this research suggests that the signs of these relations change depending on various factors, the pertinent studies nonetheless agree that the information content of the publicly traded equity is an important determinant of these links.

Rather than investigating differences in market quality surrounding changes in market-wide alterations in margin levels, our methodological design follows Seguin (1990). We use firm-by-firm changes in the level of m associated with firms becoming margin-eligible.[7] We believe that this design provides a statistically more powerful experiment to conduct our analysis for at least four reasons.

First, margin-eligibility essentially lowers the level of m for these stocks from 100% to 50%, whereas historically, the largest change in the overall margin level was 25%.[8] Thus, our method provides inherently greater statistical power, because the changes in margin

[6] Smith (1998), Barclay et al. (1999), and Chung and Van Ness (2001) describe and analyze the Order Handling Rules; Goldstein and Kavajecz (2000) analyze the change to sixteenths. While the minimum tick is currently a penny, it is unlikely that there would be significant differences between sixteenths and pennies, as the vast majority of stocks in our sample had quoted and effective spreads that were greater than US$3/16.

[7] See Hardouvelis (1988, 1990), the numerous responses that followed these articles including Schwert (1989) and Hsieh and Miller (1990), and Hardouvelis and Theodossiou (2002) for studies of overall changes in margin levels.

[8] See Alexander et al. (2001, p. 23) for a table of the Fed's changes in the margin requirement. The average of the absolute value of the margin changes is approximately 18%.

levels associated with margin-eligibility are of much larger magnitude than the changes in the overall margin level.

Second, the Fed has changed the margin level for stock purchases only 23 times since it was empowered to do so in 1934, and it has remained at 50% since January 3, 1974. In contrast, during our sample period of 1993–1998, the Fed published a list of newly margin-eligible stocks every quarter, yielding a sample of changes that is relatively current and chronologically diverse. Both the number and magnitude of the changes, yielded by our experimental design, econometrically dominate a design based on market-wide changes by offering inherently greater statistical power. Thus, the likelihood of discerning any impact of changes in margin levels on market quality should be greater and more timely in our margin-eligible sample than in a sample of changes in the overall margin requirement.

Third, because changes in margin-eligibility affect only a small subset of equities at any one time, we can control for market-wide factors that impact market quality, including, but not limited to, secular shifts in aggregate volatility, a control conspicuously absent in the work of Hardouvelis and Theodossiou (2002). As a result, we are confident that our approach provides a greater degree of ceteris paribus than studies that examine Fed-mandated changes for all traded stocks.

Finally, there are no publicly traded derivative contracts on Nasdaq Small Cap stocks, the set of stocks from which our sample is drawn. In addition, these stocks are, relative to Nasdaq NMS or exchange-listed stocks, thinly traded. Hence, it is reasonable to expect that any impact on market quality associated with changes in margin levels is more likely to be observed for the stocks included in our sample.

Our study is organized as follows. Section 2 contains a review of the literature and identifies the hypotheses, while Section 3 discusses the data. Section 4 presents our aggregate results. In this section, we also demonstrate that any latent selection bias due to the fact that the Fed does not randomly change margin-eligibility status does not affect our results. Section 5 presents the results for the two subsamples straddling the 1997 regulatory reforms, while Section 6 concludes.

2. Background and hypotheses

A major issue among regulators and market participants has been the relationship between the required level of margin and market volatility. The Securities Exchange Act of 1934, enacted in response to the October 1929 stock market crash, gave the Federal Reserve the power to regulate the proportion of a stock purchase that can be financed with debt. In doing so, Congress sought to provide the Fed with an effective policy tool to (i) control the volatility in the stock market and (ii) limit the provision of credit for the purposes of trading to avoid crowding out other types of lending by financial institutions (Schwert, 1989). While calls for raising margin levels to control stock market volatility typically increase during periods of perceived excessive speculation, the effectiveness of using margin as a policy tool has been questioned by numerous economists. For example, early research by Moore (1966) and Officer (1973) found that changes in margin levels introduced by the Fed had no significant effect on stock return volatility.

This debate was re-ignited by Hardouvelis (1988, 1990), who suggested that margin level decreases implemented by the Fed led to greater market volatility. However, subsequent evidence provided by, for example, Ferris and Chance (1988), Kupiec (1989), Schwert (1989), Hsieh and Miller (1990), Seguin (1990), and Seguin and Jarrell (1993) indicated that Hardouvelis' (1988, 1990) studies suffered from methodological flaws. Furthermore, when these flaws were corrected, researchers found that margin levels were either not related to market volatility or were positively related. In rebuttal, Hardouvelis and Theodossiou (2002) argue that the critics themselves committed methodological errors. Instead, they assert that there exists an inverse relationship between margin levels and volatility, but only during bull and normal markets, and that only during bear markets is there no relationship. Finally, Kofman and Moser (2001) investigated the relation between margin levels and volatility indirectly by examining the frequency of price reversals under various margin levels. Because they document a significant negative relationship, their study can be construed as being supportive of Hardouvelis and Theodossiou.

The majority of the above studies examined market-wide Fed-mandated changes in margin. In contrast, Seguin (1990) analyzed the effectiveness of using margin as a policy tool by contrasting the volatility and trading volume of Nasdaq stocks surrounding the date when they were declared margin-eligible by the Fed during the 1977–1987 period. He found that volume increased, while volatility decreased once stocks became margin-eligible. Eligibility appeared to be valued by the market participants, as a roughly 2% abnormal increase in the market value of newly margin-eligible stocks was observed around the time of the announcement by the Fed. Furthermore, Grube et al. (1987) and Wolfe et al. (1992) also found a similar-size abnormal increase in market value.

Mayhew et al. (1995) paralleled this analysis by examining the impact of changes in option margin levels on the market quality in the underlying stocks. They found that bid-ask spreads increased while quoted depth decreased following the January 1986 decrease in margin levels. However, they found no changes in these variables associated with the June 1988 increase in margin levels. Furthermore, the 1986 decrease (1988 increase) in margin levels led to decreases (increases) in the bid-ask spreads of the options. The authors interpreted their findings to suggest that a decrease in option margin levels is associated with a move by uninformed investors from trading the underlying stocks to the options themselves, resulting in proportionally more uninformed investors in options but fewer of them in the underlying stocks.

A related line of research has examined the introduction of publicly traded options on the market quality of the underlying stock. Lowering option margin levels and introducing options are two avenues that provide investors with an increased ability to take positions with greater leverage in a firm's stock. Evidence provided by, for example, Conrad (1989), Skinner (1989), DeTemple and Jorion (1990), Damodaran and Lim (1991), and Fedenia and Grammatikos (1992) indicated that bid-ask spreads and return volatility of the underlying stocks decline while trading volume rises after options are introduced.[9] In addition, Kumar et al. (1998) found that quoted depth increases while the adverse-

[9] Bollen (1998) found that option introduction has no impact on return volatility. In terms of the stock price reaction to option introduction, Sorescu (2000) found that stock prices reacted favorably to option introduction before 1980, but after 1980, the reaction was puzzlingly unfavorable.

selection component of the spread decreases, following the introduction of options. They conclude that the introduction of options causes informed investors to migrate from trading the underlying stocks to trading options. This, in turn, results in an increase in the proportion of traders in the underlying stocks that are uninformed.

Our primary objective is to use our sample of margin-eligibility events to analyze the effects of margin regulation on stock market quality using richer, intraday data. Given Seguin's (1990) findings that margin-eligibility was associated with a decrease in volatility and an increase in trading volume, we posit our first four hypotheses:

H1. A stock's quoted spread decreases when it becomes margin-eligible.

H2. A stock's effective spread decreases when it becomes margin-eligible.

H3. A stock's quoted depth increases when a stock becomes margin-eligible.

H4. The number of market-makers for a stock increases when the stock becomes margin-eligible.

We also extend Seguin's analyses by estimating spread components and their changes surrounding margin-eligibility. Assuming that increased volume is indeed associated with margin-eligibility and that, as traditional microstructure theory predicts, the fixed component of order-processing costs is apportioned across all trades, we posit the fifth hypothesis:

H5. The proportion of the spread that is attributable to order-processing costs decreases when a stock becomes margin-eligible.

Finally, because exchange traded options are unavailable for the stocks in our sample, we argue that margin-eligibility increases the leverage opportunities for an informed trader with finite wealth, resulting in the sixth hypothesis:

H6. The proportion of the spread that is attributable to adverse selection increases when a stock becomes margin-eligible.[10]

These six hypotheses underpin the econometric analyses that follow.

3. Data

3.1. Sample selection

We use the list of newly margin-eligible domestic stocks from the Federal Reserve Bulletin for the years 1993–1998 to form our sample.[11] The sample begins in 1993, corresponding with the initiation of the Trade and Quote (TAQ) database. During our sample

[10] It should be noted that the fifth and sixth hypotheses are in reality one hypothesis, as a change in one component will automatically lead to a change in the opposite direction in the other component.

[11] The Bulletin also lists stocks that are deleted from the margin-eligible list. Relative to additions, these stocks are few in number, trade infrequently (roughly 60% as often as the additions), and typically have prices of less than US$2 per share. Hence, we did not examine them. See Grube et al. (1987, p. 29) and Wolfe et al. (1992, pp. 95–96) for a discussion of the difficulties encountered in analyzing such stocks.

period, the Fed employed a "vetting" process where stocks were considered for addition to the margin-eligible list only after a set of criteria were met. These requirements included, among others: (i) four or more dealers making a market in the stock; (ii) a share price of US$5 or more; (iii) issuer's capital is at least US$4 million; (iv) 400,000 or more shares are held by the public; (v) 1200 shareholders of record that are not directly associated with the firm; and (vi) issuer or a predecessor has been in existence for at least 3 years. Seguin (1990, p. 118) notes that once a stock meets the eligibility criteria, it "would almost certainly appear on one of the next two eligibility lists". Our sample period ends in 1998, because starting with January 1999, the Fed ceased publishing its quarterly lists of newly margin-eligible domestic securities and simultaneously abandoned their vetting process in favor of an automatic rule. From that point, all domestic securities would become margin-eligible once listed on Nasdaq. The Fed still screens foreign stocks quarterly, but such stocks are excluded from our sample because of possible complications associated with multiple listing.

In Panel (A) of Table 1, we report that 3778 common stocks became margin-eligible over our sample period. Our final sample, however, contains only 543 stocks. This decline is attributable to numerous data requirements. We excluded stocks without continuous data beginning 120 trading days before through 100 trading days after the effective date of margin-eligibility. This requirement is imposed in order to permit comparisons of the pre-period, defined as days -100 to -12 relative to the eligibility effective date, with the post-period, defined as days $+5$ to $+100$ relative to the effective date. The additional 20 days before day -100 are included to avoid any initial listing effects such as price stabilization, during which spreads have been documented to be artificially low (e.g., Hanley et al., 1993, 1996). We cut off the pre-period at day -12 to avoid any confounding announcement effects, as the Fed announced its list between 11 and 2 trading days before each of the 24 effective dates. The -120 to $+100$ trading day requirement reduced our sample size from 3778 stocks to 1011, primarily because the median stock has only 60 days of trading data prior to margin-eligibility.

We further eliminated stocks that split within 100 trading days of the effective date to avoid any confounding affects attributable to stock splits. The stock split requirement further reduced the sample by another 126 stocks to 885. Finally, we required the stocks to be on the Nasdaq Small Cap list for the entire pre-period (some stocks migrate during this period between Small Cap and NMS status), reducing the sample by another 342 stocks, leaving a final sample of 543 stocks. We examine Nasdaq Small Cap stocks exclusive of NMS stocks because following the Fed's 1984 amendments to margin-eligibility rules, NMS stocks become automatically margin-eligible when they attain NMS status on Nasdaq.

3.2. Subsamples

We duplicate our analyses on our data dichotomized into subsamples based on the introduction of the Order Handling Rules and teenies in 1997. Such an analysis is of interest for at least two reasons. First, econometrically, the introduction of these regulatory changes may represent a regime shift. If so, then our aggregate results may inadvertently reflect a weighted average of two distinct, regime-specific sets of results. Second, economically, the introduction of these measures was, in part, intended to improve the quality of the market. Panel (B) of Table 1 categorizes the 543 firms by time period, i.e.,

Table 1
Test sample of stocks added to the margin-eligibility list

(A) By effective date	Number of additions	Number in sample	Cumulative number in sample
19930208	93	9	9
19930510	141	14	23
19930809	167	18	41
19931108	174	21	62
19940214	222	23	85
19940509	185	17	102
19940808	187	34	136
19941114	131	13	149
19950213	111	22	171
19950508	105	12	183
19950814	157	22	205
19951113	162	18	223
19960212	196	26	249
19960513	209	24	273
19960812	236	36	309
19961112	224	37	346
19970210	187	40	386
19970512	126	20	406
19970811	141	27	433
19971110	131	15	448
19980209	147	26	474
19980511	131	32	506
19980810	144	29	535
19981109	71	8	543
Total	3778	543	

(B) By subsample	Start date	End date	Number of additions
Subsample I	19930208	19970210	386
Subsample II	19980209	19981109	95
Between I and II	19970512	19971110	62
Total			543

In this table, we present the number of securities that are added to the list of margin-eligible stocks by effective date. Securities in the test sample include only those common stocks with CRSP and TAQ coverage that are an addition to the List of Marginable OTC Stocks from 1993 to 1998. To remain in the sample, a stock must have trading data on CRSP from 120 trading days before through 100 trading after the effective date and must not split during the ± 100 trading days surrounding the effective date. Stocks becoming margin-eligible by virtue of becoming a Nasdaq NMS stock or that were NMS stocks for part of the -100 to 0 period are excluded. Subsample I refers to additions during the pre-reform period; Subsample II refers to additions during the post-reform period.

whether the margin-eligibility date occurred prior to (“Subsample I”) or subsequent to (“Subsample II”) both the implementation of the Order Handling Rules and the introduction of teenies. The Order Handling Rules were implemented in 22 waves beginning January 1, 1997 and ending October 13, 1997; Small Cap Nasdaq stocks were not included in this process before August 4, 1997. Teenies were implemented on Nasdaq on June 24, 1997. Hence, the 17 quarterly additions to the Fed's margin-eligible list from February 8, 1993 through February 10, 1997 form the first subsample of 386 additions that took place before either of these reforms. The four quarterly additions from February 9, 1998 through November 9, 1998 form a second subsample of 95 additions that took place after the two reforms.

The set of 40 stocks that became margin-eligible on February 10, 1997 have approximately 80 days of trading with a minimum tick size of US$1/8 and 20 days of trading with a minimum tick size of US$1/16. Because our control sample covers the exact same time period and the number of days of overlap is small, we allocated these stocks in the first subsample. The set of 26 stocks that became margin-eligible on February 9, 1998 had approximately 92 days of trading after the final wave of the Order Handling Rules (on October 13, 1997) and 97 days after the penultimate wave (on October 6, 1997). Given that (1) our control sample covers the exact same time period; (2) none of the 26 stocks were in the final wave and 13 were in the penultimate wave; and (3) the number of days of overlap is very small, we allocated these stocks to the second subsample.

3.3. Control sample

To accommodate changes in general market liquidity (Chordia et al., 2000, 2001), we contrast the observed changes for our sample to changes observed from a control sample. We selected a control stock for each stock in our test sample on the following basis. First, control sample stocks must, like the test sample stocks, have been Nasdaq Small Cap stocks during the entire pre-period.[12] Second, the control sample stocks must have met the same data requirements as the test sample stocks. Specifically, we required trading data from 120 days before the margin-eligibility date to 100 days after, and control sample stocks must not have split in the period of 100 days before to 100 days after the margin-eligibility date. Third, we required the control sample stocks to have the same minimum tick size as the test sample stocks in the pre-period. The minimum tick size for stocks selling below US$10 was US$1/32, while the minimum tick was US$1/8 if the price was above US$10 (US$1/16 after the change to teenies in 1997). Hence, we required the control sample stocks to have an average price above (below) US$10 in the pre-period if the test sample stocks have an average price above (below) US$10. Fourth, the control sample stocks could not be in the test sample for that margin-eligibility date, nor could they be foreign stocks. For each test sample stock, one stock was randomly selected from the set that passes these four screens to serve as the control sample stock. Sampling was done without replacement for each effective date. Our control sample selection process provides close but inexact matches for our analysis. To accommodate any remaining differences in variables known to affect market quality, we also calculated the values of four control variables—average trade size, volume, volatility, and price. These variables are then included in cross-sectional regressions below.

[12] Thus, the control stocks are either margin-ineligible or margin-eligible. Restricting the control sample to either type would bias the results. Relative to the test sample, using just margin-eligible stocks would result in a control sample of larger, more prosperous stocks, while using just margin-ineligible stocks would result in a control sample of smaller, less prosperous stocks.

4. Overall results

Our analysis begins with an examination of the impact of margin-eligibility on equity trading. We first calculate the average number of trades per day, trade size, and daily volume in both the pre- and post-eligibility period for each firm. Panel (A) of Table 2 reports the cross-sectional averages and medians in the pre- and post-periods for the full sample as well as the cross-sectional averages of the corresponding ln(post/pre) values. We used natural log transformations to mitigate the effects of the inherent skewness in the sample distributions. As an additional benefit, the use of this transformation allows the resulting value to be interpreted as the approximate percentage change in the variable in question. The table also reports the percentage of the ln(post/pre) values that are negative. We report both a *t*-test and a Wilcoxon signed rank test for the cross-sectional average comparisons, and a non-parametric sign test for the percent negative.

As shown in Table 2, following margin-eligibility, there were insignificant increases of roughly 2.8% in the average number of trades per day, 0.9% in average trade size, and 3.7% in average daily volume.[13] Relative to the control sample, only the average trade size increased by a significant amount.

One of the econometric shortfalls associated with these univariate tests is that the assumption of cross-sectional independence is violated because the additions are clustered around 24 margin-eligibility dates. Following Seguin (1990), we mitigate this problem by deflating the data for each test and control sample stock by dividing its value for each day from $t = -100$ through $t = -12$ and from $t = +5$ through $t = +100$ (Y_{it}) by its pre-period mean ($\overline{Y}_i$) as follows:

$$Y'_{it} = Y_{it}/\overline{Y}_i \tag{1}$$

where:

$$\overline{Y}_i = \sum_{t=-100}^{-12} Y_{it}/89 \tag{2}$$

These deflated values (Y'_{it}) are then aggregated cross-sectionally for each day to get a daily average value (Y_t):

$$Y_t = \sum_{i=1}^{n} Y'_{it}/n \tag{3}$$

After calculating the time-series of these daily cross-sectional averages, we estimate the following time-series regression:

$$Y_{\mathrm{TEST}t} - Y_{\mathrm{CONTROL}t} = a + b\mathrm{MARDUM}_t + e_t \tag{4}$$

Note that this equation involves the daily difference between the test and control sample averages; MARDUM_t is a dummy variable that equals 1 if observation Y_t is from the post-period, and 0 if otherwise.

[13] Chordia et al. (2001, pp. 511–512) found an increase in NYSE trading volume and number of trades during the 1988–1998 period.

Table 2
Cross-sectional analysis of trading activity and volatility of stocks that have been added to the margin-eligibility list

	Pre-mean (median)	Post-mean (median)	% < 0	Test sample difference (standard error)	Control sample difference (standard error)	Difference in differences (standard error)
(A) Trading activity						
Number of trades/day	22.7 (8.9)	27.6 (8.6)	53.7	0.028 (0.656)	0.037 (0.658)	− 0.008 (0.905)
Trade size	1610 (1410)	1660 (1440)	50.2	0.009 (0.355)	− 0.038**† (0.418)	0.047** (0.541)
Daily volume	33,500 (13,100)	40,300 (15,800)	51.5	0.037 (0.773)	− 0.002 (0.772)	0.039 (1.060)
(B) Volatility	0.034 (0.031)	0.039 (0.034)	36.3###	0.161***††† (0.582)	0.093***††† (0.611)	0.068*†† (0.842)

In this table, we report the cross-sectional analysis of trading activity and volatility for our entire sample (see Table 1). A control sample was formed by matching each test sample stock with a randomly selected Nasdaq Small Cap stock. Potential control sample stocks were excluded from consideration if they split within $\pm$ 100 trading days relative to the effective date. Control sample stocks were matched on the average price (less than US\$10 or greater than US\$10) during the pre-period. Columns with the heading ‘Pre’ refer to pre-period trading days $-$ 100 to $-$ 12 relative to the effective date of margin-eligibility, while columns with the heading ‘Post’ refers to post-period trading days + 5 to + 100 relative to the effective date. The column with the heading ‘Test sample difference’ is the cross-sectional average of the ln(post/pre) values for the test sample firms. The column with the heading ‘Control sample difference’ is the difference for the control sample firms. The column with the heading ‘Difference in differences’ is the average difference between the test and control samples. Volatility refers to the standard deviation of the continuously compounded close-to-close returns. Trades and quotes are filtered following Bessembinder (1997). Significance differences from zero are indicated with ‘*’/‘†’ for a *t*-test/Wilcoxon signed rank test. ‘#’ indicates that percentage negative is significantly different from 50% using a sign test. Three/two/one significance symbol(s) indicate(s) significance at the 1%/5%/10% level.

Table 3
Time-series analysis of measures of trading activity and volatility of stocks that have been added to the margin-eligibility list

Standardized dependent variable	Intercept	MARDUM	Regression R^2
(A) Trading activity			
Number of trades	− 0.289*** (0.100)	− 0.027 (0.135)	0.001
Trade size	0.016 (0.010)	0.025* (0.013)	0.023
Daily volume	− 0.295*** (0.056)	0.150* (0.078)	0.036
(B) Volatility	− 0.177 (0.275)	− 0.509 (0.381)	0.010

In this table, we report the results of time-series regressions for standardized variables. We standardize trading activity variables by deflating each day's value by the pre-eligibility mean; volatility is standardized by deflating the squared deviation of each day's return by the pre-eligibility variance. Next, we form portfolios for the test and control sample (see Tables 1 and 2). The time series used in the regression is calculated as the test-control variable. Significance at the 1%/5%/10% level is indicated with ***/**/*. MARDUM is a binary variable set to 1 if the day is after the effective date, 0 if otherwise. Time-series regressions are estimated after explicitly accommodating autocorrelation. Standard errors are in parentheses. The reported R^2 is a measure of fit for the structural part of the model after transforming for the autocorrelation and is the R^2 for the transformed regression.

The results from the time-series regressions that explicitly accommodate autocorrelation are reported in Table 3. The key coefficient in these regressions is b, which we interpret as the percent change in the relevant dependent variable upon eligibility. These values are reported under the MARDUM column in the table. The results indicate that the average trade size and daily volume increased significantly after margin-eligibility by 2.5% and 15%, respectively, relative to the control sample.[14]

Panel (B) of Table 2 shows that, in contrast to Seguin (1990), there is a significant 16.1% increase in volatility following margin-eligibility. This increase in volatility exceeded the 9.3% increase for the control sample. To further explore this issue, we first examined the monthly volatility of continuously compounded closing returns of the Nasdaq Composite, Russell 2000, and Wilshire 5000 indexes during our 6-year sample period of 1993–1998. The slope coefficient of the 72 monthly measures of volatility regressed on a trend variable indicate that market-wide volatility was increasing by a significant amount over our sample period. Thus, it appears that there was a trend of increasing volatility during the sample period that was reflected in the observed increases for both the test and control samples.

Having noted this secular increase in volatility during the sample period, we next addressed the lack of cross-sectional independence in the Table 2 tests by utilizing a time-series test of volatility using Eq. (4). However, a modified or standardized value of Y'_{it} was used:

$$Y'_{it} = (Y_{it} - \overline{Y}_i)^2/[(88/89)s_i^2] \tag{5}$$

where s_i^2 is the unbiased estimator of the variance of stock i during the pre-period. With these values of Y'_{it}, values of Y_t were calculated for each day from − 100 to − 12 and + 5

[14] Although Table 3 indicates significance at only the 10% level, the p-values corresponding to the MARDUM coefficient for the Trade Size and Daily Volume regressions were 0.059 and 0.056.

to +100 using Eq. (3). Panel (B) of Table 3 indicates that the value of b, the coefficient to MARDUM in Eq. (4), was not significantly different from 0 at even the 10% level.[15] Thus, the increase in volatility in the test sample that was observed in Table 2 was not significantly different from the increase in the control sample; instead, it was simply a manifestation of a secular increase in volatility in the market.

4.1. Market liquidity

The first major objective of our study is to investigate the change in liquidity following a margin-eligibility. Our measures of liquidity include time-weighted quoted spreads, trade-weighted effective spreads, time-weighted quoted depth (measured as the average of the inside bid and ask depths), and the number of market-makers. Panel (A) of Table 4 reports the cross-sectional averages and medians in the pre- and post-periods for the full sample as well as the cross-sectional averages of the corresponding ln(post/pre) values.

The results under the "Test Sample Difference" column detail an economically and statistically significant increase in liquidity following margin-eligibility. Cross-sectional average quoted and effective spreads decreased by roughly 6.6% and 7.5% (H1 and H2), respectively, while depth and the number of market-makers increased by 2.5% (H3) and 0.4% (H4), respectively.[16] With the exception of the market-maker tests, all findings are significant using both parametric and non-parametric tests. However, similar results occurred for the control sample, as shown under the "Control sample difference" column. Given the control sample results, it is not surprising that the "Difference in differences" column indicates that none of the changes were significantly different between the test and control samples.

The results from the Eq. (4) time-series regressions that explicitly accommodate autocorrelation are reported for the full sample in Panel (A) of Table 5. As noted with respect to Table 3, the key coefficient in these regressions is b, reported in the MARDUM row. The estimates indicate no significant changes in quoted spreads, effective spreads, and depth. However, the number of market-makers decreases by 0.6% (H4), an amount that is marginally significant, but not economically material.

In Table 6, we test the robustness of our results by controlling for variables known to affect spreads given earlier in Table 2. Following Mayhew et al. (1995) and Kumar et al. (1998), we estimate cross-sectional regressions of the change in the four measures of market quality given in Table 5—quoted spreads, effective spreads, depth, and the number of market-makers—on changes in four control variables. Our explanatory variables are chosen based on Benston and Hagerman (1974) and Stoll (1978), who show that spreads are negatively related to price and volume and positively related to risk. Because Easley and O'Hara (1987) show that trade size is related to the adverse-selection

[15] Insignificant changes in volatility were also observed when Eq. (5) was replaced with:

$$Y'_{it} = (Y_{it} - \overline{Y}_i)^2 - (88/89)s_i^2.$$

[16] Chordia et al. (2001, p. 508) found a decreasing trend in quoted and effective spreads for NYSE stocks during the 1988–1998 period.

Table 4
Cross-sectional analysis of market liquidity measures of stocks that have been added to the margin-eligibility list

	Pre-mean (median)	Post-mean (median)	% < 0	Test sample difference (standard error)	Control sample difference (standard error)	Difference in differences (standard error)
(A) Full sample						
Quoted spread (¢)	48.7 (39.2)	48.3 (36.2)	56.7###	– 0.066***††† (0.339)	– 0.061***††† (0.301)	– 0.005 (0.438)
Effective spread (¢)	40.0 (30.1)	36.1 (27.3)	58.0###	– 0.075***††† (0.396)	– 0.057***††† (0.319)	– 0.018 (0.485)
Depth (shares)	736 (727)	783 (732)	40.7##	0.025**†† (0.261)	0.044***††† (0.283)	– 0.019 (0.347)
No. of market-makers	9.9 (8.8)	10.1 (9.1)	48.7	0.004 (0.221)	0.012 (0.201)	– 0.008 (0.282)
(B) Subsamples						
Subsample I:						
Quoted spread (¢)	51.1 (43.8)	51.6 (39.9)	53.9	– 0.045***†† (0.326)	– 0.051***††† (0.268)	0.007 (0.410)
Effective spread (¢)	43.5 (33.5)	38.1 (29.5)	55.6###	– 0.063***††† (0.413)	– 0.045***††† (0.285)	– 0.018 (0.479)
Depth (shares)	699 (687)	699 (673)	40.8	0.001 (0.163)	0.010 (0.161)	– 0.008 (0.222)
No. of market-makers	10.4 (9.6)	10.3 (9.5)	53.2	– 0.012 (0.206)	0.001 (0.186)	– 0.012 (0.267)
Subsample II:						
Quoted spread (¢)	37.3 (30.5)	37.8 (26.0)	57.4	– 0.063*†† (0.375)	– 0.059 (0.373)	– 0.004 (0.533)
Effective spread (¢)	27.4 (22.8)	27.7 (20.5)	60.6##	– 0.056*† (0.333)	– 0.051 (0.385)	– 0.005 (0.521)
Depth (shares)	937 (859)	906 (844)	56.4	– 0.042 (0.293)	0.007 (0.457)	– 0.049 (0.544)
No. of market-makers	9.1 (7.7)	9.9 (8.5)	34.0###	0.069***††† (0.256)	0.051**† (0.220)	0.018 (0.327)

In this table, we report the cross-sectional analysis of various measures of market liquidity for our entire sample and subsamples defined in Tables 1 and 2. Columns with the heading 'Pre' refer to pre-period trading days – 100 to – 12 relative to the effective date of margin-eligibility, while columns with the heading 'Post' refers to post-period trading days + 5 to + 100 relative to the effective date. The column with the heading 'Test sample difference' is the cross-sectional average of the ln(post/pre) values for the test sample firms. The column with the heading 'Control sample difference' is the difference for the control sample firms. The column with the heading 'Difference in differences' is the average difference between the test and control samples. Quoted depths are measured as the average of the bid and ask depths. Quoted spreads and depths are time-weighted. Effective spreads are calculated as 2 × | Price – Quote Midpoint | ; reported numbers are trade-weighted. Trades and quotes are filtered following Bessembinder (1997). The number of market-makers is time-weighted. Significance differences from zero are indicated with '*'/'†' for a *t*-test/Wilcoxon signed rank test. '#' indicates the percentage negative is significantly different from 50% using a sign test. Three/two/one significance symbol(s) indicate(s) significance at the 1%/5%/10% level.

Table 5
Time-series analysis of market liquidity measures of stocks that have been added to the margin-eligibility list

	Quoted spread	Effective spread	Depth	No. of market-makers
(A) Full sample				
Intercept	– 0.006 (0.009)	– 0.008 (0.009)	0.004 (0.015)	0.004 (0.003)
MARDUM	0.011 (0.012)	0.009 (0.013)	– 0.001 (0.020)	– 0.006* (0.004)
R^2	0.005	0.002	0.000	0.016
(B) Subsamples				
Subsample I				
Intercept	0.001 (0.011)	– 0.015 (0.013)	– 0.000 (0.003)	0.003 (0.002)
MARDUM	0.013 (0.013)	0.026 (0.018)	– 0.009** (0.005)	– 0.009*** (0.003)
R^2	0.006	0.011	0.021	0.049
Subsample II				
Intercept	– 0.002 (0.011)	0.015 (0.011)	0.011 (0.046)	0.006 (0.009)
MARDUM	– 0.028* (0.016)	– 0.015 (0.015)	– 0.153*** (0.061)	0.017 (0.011)
R^2	0.018	0.006	0.033	0.013
Subsample I – Subsample II				
Intercept	0.002 (0.011)	– 0.030* (0.016)	– 0.012 (0.046)	– 0.003 (0.009)
MARDUM	0.050*** (0.015)	0.041* (0.022)	0.144*** (0.062)	– 0.026*** (0.011)
R^2	0.055	0.018	0.029	0.032

In this table, we report the results of time-series regressions for standardized variables. We standardize variables by deflating each day's value by the pre-eligibility mean. Next, we form portfolios for the test and control sample (defined in Tables 1 and 2). The time series used in the regression is calculated as the test-control variable. Significance at the 1%/5%/10% level is indicated with ***/**/*. The quoted spread and depth are time-weighted and the effective spread is trade-weighted. MARDUM is a binary variable set to 1 if the day is after the effective date, 0 if otherwise. Time-series regressions are estimated after explicitly accommodating autocorrelation. Standard errors are in parentheses. The reported R^2 is a measure of fit for the structural part of the model.

component of the spread, we include average trade size as well. We estimate regressions of the form:

$$\Delta Y_i = a + b_1 \Delta \text{PRICE}_i + b_2 \Delta \text{TRADESIZE}_i + b_3 \Delta \text{VOLUME}_i + b_4 \Delta \text{VOLATILITY}_i + e_i \tag{6}$$

where for test sample stock i, ΔY_i is the change in the market-quality measure, ΔPRICE_i is the change in the average price, $\Delta\text{TRADESIZE}_i$ is the change in the average trade size, ΔVOLUME_i is the change in the average daily volume, and $\Delta\text{VOLATILITY}_i$ is the change in volatility. In each case, the change is measured relative to the control sample stock.

Consistent with extant literature, the estimated coefficients indicate that higher share price, trade size, and volatility, along with lower volume, were associated with wider quoted and effective spreads. Of primary importance, however, are the intercepts, which measure the change in the spread measure, ceteris paribus. The intercepts for both the quoted and effective spread regressions again indicate that there was an insignificant change associated with becoming margin-eligible.

Table 6
Cross-sectional regression analysis involving market liquidity measures of stocks that have been added to the margin-eligibility list

	(A) Full sample				(B) Subsamples			
	ΔQuoted spread	ΔEffective spread	ΔQuoted depth	ΔMarket-makers	ΔQuoted spread	ΔEffective spread	ΔQuoted depth	ΔMarket-makers
Intercept	0.005	0.002	−0.030**	0.001				
	(0.016)	(0.018)	(0.015)	(0.012)				
Intercept I					0.026	0.014	−0.020	−0.004
					(0.020)	(0.022)	(0.016)	(0.015)
Intercept II					−0.015	−0.015	−0.050	0.024
					(0.038)	(0.042)	(0.036)	(0.027)
SUB					−0.041	−0.029	−0.030	0.029
					(0.043)	(0.048)	(0.036)	(0.031)
ΔPrice	0.448***	0.512***	−0.183***	−0.049*	0.451***	0.507***	−0.117***	−0.066**
	(0.037)	(0.040)	(0.033)	(0.027)	(0.041)	(0.046)	(0.034)	(0.030)
ΔTrade size	0.181***	0.132***	0.018	−0.116***	0.183***	0.128***	0.045	−0.136***
	(0.042)	(0.046)	(0.038)	(0.032)	(0.044)	(0.049)	(0.037)	(0.033)
ΔVolume	−0.240***	−0.242***	0.061***	0.126***	−0.244***	−0.243***	0.039**	0.130***
	(0.022)	(0.024)	(0.020)	(0.016)	(0.024)	(0.027)	(0.020)	(0.017)
ΔVolatility	0.153***	0.176***	−0.029	−0.098***	0.153***	0.183***	−0.031*	−0.094***
	(0.021)	(0.023)	(0.019)	(0.016)	(0.022)	(0.024)	(0.018)	(0.016)
Adjusted R^2	0.297	0.322	0.085	0.137	0.288	0.304	0.057	0.138

In this table, we report the results of cross-sectional regressions that examine the determinants of the abnormal changes in market liquidity for our entire sample and subsamples defined in Tables 1 and 2. The dependent variables in the regressions are the abnormal changes in market liquidity variables, i.e., test-control. The explanatory variables are calculated similarly. The 'Δ' operator indicates ln(post/pre) values for the variables, where the Pre-period and Post-period are defined in Table 4. Significance at the 1%/5%/10% level is indicated with ***/**/*. In Panel (B), the value for Intercept I corresponds to Intercept with the binary variable SUB defined as 0 if the margin-eligibility date occurred in Subsample I, 1 if otherwise. Thus, Intercept I is the value of the intercept for the first subsample. The value for Intercept II corresponds to Intercept with SUB defined as 0 if the margin-eligibility date occurred in Subsample II, 1 if otherwise. Thus, Intercept II is the value of the intercept for the second subsample. Full and subsamples are defined in Tables 1 and 2. Standard errors are in parentheses.

Furthermore, lower share price and higher volume were associated with greater depths. More importantly, the intercept indicates a significant decrease in depth associated with eligibility. However, given an average depth of slightly more than 700 shares as shown in Table 2, the decrease of three shares (depths are measured in round lots, so 0.03 is equivalent to three shares) is not economically material.

Lastly, the table shows that there was no significant change in the number of market-makers, and that higher share price, trade size, and volatility, along with lower volume, were associated with fewer market-makers. In sum, we interpret the results of Tables 4–6 as providing little evidence supporting any of the first four hypotheses. We conclude that margin-eligibility has no substantive impact on liquidity.

4.2. Spread decomposition surrounding eligibility

The second major objective of our study is to evaluate changes in the adverse-selection and order-processing cost spread components associated with margin-eligibility with the spread decomposition model of Madhavan et al. (1997).

4.2.1. The Madhavan–Richardson–Roomans model

Using the Madhavan–Richardson–Roomans model, Panel (A) of Table 7 reports that for the full sample, the proportion of the spread attributable to adverse selection decreased by, on average, an insignificant 4.4%, while the proportion associated with order-processing costs increased by a significant 6.4% following margin-eligibility. Both of these changes appear to contradict H5 and H6. However, the changes in the control sample firms were even greater. As a result, we focus on the final column, which reports that, compared to the control sample, the order-processing portion of the spread declined, while the adverse-selection component increased. Thus, Table 7 provides mild support for H5 and H6.

To control for market-wide changes and to generate measures of statistical significance, we estimate the cross-sectional regressions of Eq. (6) using the four control variables from Table 2 with changes in the components of the spread used as dependent variables. As in Table 6, in each case the change is measured relative to the change in the control sample firm. The results are reported in Panel (A) of Table 8. Examination of the intercept reveals that upon margin-eligibility, the adverse-selection component increased significantly, while the order-processing cost component decreased significantly. In sum, we interpret the results of Tables 7 and 8 as evidence in support of the fifth and sixth hypotheses: the adverse-selection spread component increased and the order-processing costs component decreased following margin-eligibility. In conjunction with our earlier evidence of no change in liquidity and an increase in volume, we interpret this evidence as supporting our view that the proportion of informed investors increases when a stock becomes margin-eligible, thereby improving the information content of share prices.

4.2.2. The Glosten–Harris model

As a robustness test, we repeated the above analyses using the Glosten–Harris (1988) model as formulated by Brennan and Subrahmanyam (1996, p. 444):

$$\Delta p_t = \lambda q_t + \Psi(I_t - I_{t-1}) + e_t \tag{7}$$

Table 7
Cross-sectional analysis of spread components for stocks that have been added to the margin-eligibility list

	Pre-mean (median)	Post-mean (median)	% < 0	Test sample difference (standard error)	Control sample difference (standard error)	Difference in differences (standard error)
(A) Full sample						
Adverse selection	0.110 (0.065)	0.082 (0.068)	54.2	− 0.044 (1.225)	− 0.272***††† (1.234)	0.255**† (1.716)
Order processing	0.890 (0.935)	0.918 (0.932)	45.8	0.064***†† (0.397)	0.131***††† (0.400)	− 0.054†† (0.541)
(B) Subsamples						
Subsample I						
Adverse selection	0.091 (0.049)	0.065 (0.055)	52.9	− 0.031 (1.325)	− 0.456***††† (1.329)	0.478***††† (1.859)
Order processing	0.909 (0.951)	0.935 (0.945)	47.1	0.049**† (0.338)	0.151***††† (0.377)	− 0.093***††† (0.341)
Subsample II						
Adverse selection	0.164 (0.113)	0.123 (0.111)	60.8*	− 0.175 (0.989)	0.041 (0.940)	− 0.192 (1.360)
Order processing	0.836 (0.886)	0.877 (0.889)	39.2*	0.101*† (0.498)	0.073** (0.292)	0.048 (0.651)

In this table, we report the cross-sectional analysis of spread decomposition measures for our entire sample and subsamples defined in Tables 1 and 2. Column with the heading 'Pre' refers to pre-period trading days − 100 to − 12 relative to the effective date of margin-eligibility, while column with the heading 'Post' refers to post-period trading days + 5 to + 100 relative to the effective date. The column with the heading 'Test sample difference' is the cross-sectional average of the ln(post/pre) values for the test sample firms. The column with the heading 'Control Sample Difference' is the difference for the control sample firms (see Table 2). The column with the heading 'Difference in differences' is the average difference between the test and control samples. The spread components are estimated using the Madhavan et al. (1997) method. Any parameter estimate implying a negative cost is excluded. This procedure may result in difference in differences not being equal to the difference of the sample difference and control difference means. Significance differences from zero are indicated with '*'/'†' for a *t*-test/Wilcoxon signed rank test. '#' indicates the percentage negative is significantly different from 50% using a sign test. Three/two/one significance symbol(s) indicate(s) significance at the 1%/5%/10% level.

where Δp_t is the change in a stock's price from trade $t-1$ to trade t, q_t is the size of trade t, I_t is $+1$ for buyer-initiated trades and -1 for seller-initiated trades, and e_t is an error term. In Eq. (7) the parameter λ is the adverse-selection component of the spread, while Ψ is the order-processing component. Because our concern is the change in λ between the pre-period and post-period, a dummy variable D_t was added to this equation, taking on a value of 0 if trade t was in the pre-period and 1 if the trade was in the post-period. Thus, we estimated:

$$\Delta p_t = \lambda_1 q_t + \lambda_2 q_t D_t + \Psi_1 (I_t - I_{t-1}) + \Psi_2 (I_t - I_{t-1}) D_t + e_t \tag{8}$$

with ordinary least squares allowing for an intercept term. If either a test firm or its control firm had a negative estimate for λ_1, then the pair of firms was omitted from further analysis (Brennan and Subrahmanyam, 1996, p. 447, also discard firms that have negative estimates). In order to account for the average transaction size, λ_2/P and $\lambda_2\ q/P$ were

Table 8
Cross-sectional regression analysis of spread components of stocks that have been added to margin-eligibility list

	(A) Full sample		(B) Subsamples	
	ΔAdverse selection	ΔOrder processing	ΔAdverse selection	ΔOrder processing
Intercept	0.2718** (0.1139)	− 0.0598* (0.0367)		
Intercept I			0.4648*** (0.1458)	− 0.0857** (0.0250)
Intercept II			− 0.2039 (0.2290)	0.0220 (0.0616)
SUB			− 0.6687** (0.2723)	0.1077 (0.0733)
ΔPrice	0.0686 (0.2685)	0.1040 (0.0866)	0.0569 (0.3094)	0.0563 (0.0833)
ΔTrade size	0.8257*** (0.3346)	− 0.1384 (0.1079)	0.8820*** (0.3581)	− 0.1388 (0.0964)
ΔVolume	− 0.6341*** (0.1615)	− 0.0034 (0.0521)	− 0.6412*** (0.1768)	0.0273 (0.0476)
ΔVolatility	0.0579 (0.1782)	0.1218** (0.0575)	0.1864 (0.1925)	0.0914* (0.0518)
Adjusted R^2	0.077	0.036	0.096	0.041

In this table, we report the results of cross-sectional regressions that examine the determinants of abnormal changes in percentage spread components for our entire sample and subsamples defined in Tables 1 and 2. The dependent variables in the regressions are the abnormal changes in spread components, i.e., test-control. The explanatory variables are calculated similarly. We employ the Madhavan et al. (1997) method for estimating these components, excluding observations that imply negative costs. The 'Δ' operator indicates ln(post/pre) values for the variables. Significance at the 1%/5%/10% level is indicated with ***/**/*. In Panel (B), the value for Intercept I corresponds to Intercept with the binary variable SUB defined as 0 if the margin-eligibility date occurred in Subsample I, 1 if otherwise. Thus, Intercept I is the value of the intercept for the first subsample. The value for Intercept II corresponds to Intercept with SUB defined as 0 if the margin-eligibility date occurred in Subsample II, 1 if otherwise. Thus, Intercept II is the value of the intercept for the second subsample. Standard errors are in parentheses.

analyzed, where q and P are the average trade size and stock price in the post-period, respectively.

While the average value of the change in the adverse-selection component, whether measured by λ_2/P or $\lambda_2 q/P$, was negative for the test sample, it was even more negative than the control sample, indicating the average value of the adverse-selection component rose in the test sample relative to the control sample. Lastly, the estimated value of Ψ_2 for the test sample, relative to the control sample, was negative, indicating the order-processing costs declined. While the difference between the test and control samples was not significant in any of these comparisons when using either a t-test or Wilcoxon signed rank test (except for when the Wilcoxon test was applied to λ_2/P), it was of the same sign as the Madhavan, Richardson, and Roomans results reported in Tables 7 and 8.

4.3. Selection bias

It could be argued that because the Fed did not randomly select OTC stocks to become margin-eligible, these results are driven by selection bias. Despite the fact that the Fed used a publicly known set of criteria (described earlier) to decide whether a particular OTC firm should be included in the list of margin-eligible stocks, one could argue that the Fed's decision reflects an improvement in a firm's viability since it has now met the eligibility criteria, even though the Fed actively denies this (Grube et al., 1987, p. 19). If this were true, one could criticize our tests as merely detecting changes in the market quality that follow the evolution of the company as a successful enterprise, and not necessarily changes that are attributable to the introduction of margin trading.

However, under this unlikely scenario the effective date when margin trading is introduced would be a non-event, because many of the changes in a company's prospects would be spread across time prior to the introduction of margin trading, perhaps as far as six months in the past according to Seguin (1990, p. 118). Consequently, our tests would not pick up significant changes in market quality around the effective date. The fact that we find significant improvements in market quality in the post- vs. pre-period suggests that the introduction of margin trading, not selection bias, is causative.

Nevertheless, we further examined three variables during the pre-period in order to see if their evolution in the test sample differed from the evolution in the control sample. The first variable investigated was the quoted spread. First, we determined the time-weighted quoted spread for each day from -100 to -12 for each test sample stock. Second, first differences were determined for each stock. Thus, for each stock, we have a vector of 88 (-100 to -12, less 1 day due to differencing) event days that correspond to daily changes in its time-weighted quoted spread. Third, for each of the 24 margin-eligibility dates, the changes were averaged across all the stocks for each particular event day, leaving us with 24 sets of 88-element vectors. Each of these 24 sets contains a time-series of averaged first-differences in time-weighted quoted spreads, with each set pertaining to a different margin-eligibility date. Fourth, for each of the 88 event days, we took an equally weighted average across the 24 sets. A similar time series was constructed from the control sample.

The resulting test and control sample time-series columns of average differences were compared by a *t*-test and a Wilcoxon signed rank test. Note that by reducing the data into two vectors and using the empirical distribution and standard errors of the aggregated averages themselves, we have circumvented any spurious correlation due to clustering.[17] The test was repeated for changes in depth (where the time-weighted average of the bid and ask depths was used) and price (where the time-weighted midpoint in the quoted spread was used).

The mean difference for daily changes in quoted spreads, with *p*-values in parentheses, was -0.00004 (0.98). Similarly, the mean difference for daily changes in depth was -0.01740 (0.63), and for daily changes in prices was -0.00593 (0.06).[18] A Pearson product–moment correlation coefficient was also computed to compare the test and control sample time-series for changes in spreads, resulting in a coefficient and *p*-value of -0.04 (0.71). Similarly, the coefficient and *p*-value for changes in depth were -0.07 (0.52), and for changes in prices were 0.13 (0.22).[19]

In sum, we found no evidence that the test sample securities evolved differently than their control sample counterparts in the pre-period. Overall, these results suggest that selection bias is either not present or has no impact on our variables of interest in the pre-period for our test sample.

[17] If we had, for example, calculated standard deviations from individual observations, s, and inferred that the standard error for a particular event day was $s/\sqrt{543}$, then our inferences would be in error (our test "sizes" would be inaccurate) since the 543 observations are not independent.

[18] Similar results were obtained when examining median differences.

[19] Similar results were obtained with Spearman rank-order correlation coefficients. In addition, all of these tests were repeated using globally equally weighted time series, again with similar results.

5. The impact of "teenies" and Order Handling Rules

In this section, we investigate the impact of the two major regulatory reforms that occurred during our sample period of 1993–1998. Specifically, "teenies" were introduced and the Order Handling Rules were implemented in 1997. The objective of this section is to determine whether these reforms affect our overall results. It is important to perform these subsample analyses for a number of reasons. First, with the introduction of teenies, minimum quoted spreads were essentially halved.[20] In turn, this could conceivably lead to a reduction in the number of observations at the minimum spread. If so, the dispersions in our variables of interest increase, and the variables are measured with less error. Both should enhance the statistical power of our tests. Second, with the notably increased ability to use limit orders, greater transparency, and improved access to ECNs like Instinet and SelectNet resulting from the Order Handling Rules, we argue that the ability of informed traders to exploit information is significantly different after the reforms.

To determine the impact of these reforms, we partition our sample into two subsamples: the 386 additions before and the 95 additions after the implementation dates for both regulatory reforms. As mentioned earlier, Panel (B) of Table 1 presents these two subsamples. In Panel (B) of Tables 4–8, we report the subsample results.

5.1. Market liquidity statistics

Panel (B) of Table 4 indicates that there are no significant differences in quoted spreads, effective spreads, depths, or the number of market-makers in either subsample. However, Panel (B) of Table 5, adjusting for cross-sectional dependencies by applying Eq. (4) to each subsample, indicates that there is a marginally significant decline in the quoted spread in Subsample II. Furthermore, there is a significant decline in depths in both subsamples, particularly Subsample II. While this may seem odd in that the decline for the full sample, shown in Panel (A), is insignificant, the difference can be attributed to the three 1997 effective dates indicated in Table 1 that were in the full sample, but not in either subsample. Lastly, note that the decline in the number of market-makers observed for the full sample appears to have occurred only during the first subsample, as there is an increase, albeit insignificant, in the second subsample.

In order to see if the differences between the two subsamples are significant, the second subsample values, $Y_{\mathrm{TESTII}t} - Y_{\mathrm{CONTROLII}t}$, were subtracted from the first subsample values, $Y_{\mathrm{TESTI}t} - Y_{\mathrm{CONTROLI}t}$, with the resulting variables applied to Eq. (4) as follows:

$$(Y_{\mathrm{TESTI}t} - Y_{\mathrm{CONTROLI}t}) - (Y_{\mathrm{TESTII}t} - Y_{\mathrm{CONTROLII}t}) = a + b\mathrm{MARDUM}_t + e_t \qquad (9)$$

Thus, the coefficient b indicates the magnitude of change in the variable around the margin-eligibility date for the first subsample relative to the second subsample.

[20] This was true only for stocks whose price is above US\$10; the minimum spread for those priced below US\$10 remained at US\$1/32 during our full sample period.

The bottom panel of Table 5 shows that the difference in the MARDUM coefficients between the two subsamples is significant for all four variables. The Subsample II decrease in both quoted and effective spread is significantly greater than in Subsample I. Furthermore, the decrease in the number of market-makers is significantly greater in the first subsample. While these observations suggest that becoming margin-eligible in the post-1997 period led to greater liquidity, the greater decrease in depth serves as an offset.

The dummy variable SUB was added to cross-sectional Eq. (6) to test for differences in spreads between the two subsamples as follows:

$$\Delta SPD_i = a + b_1\Delta\mathrm{PRICE}_i + b_2\Delta\mathrm{TRADESIZE}_i + b_3\Delta\mathrm{VOLUME}_i + b_4\Delta\mathrm{VOLATILITY}_i + b_5\mathrm{SUB}_i + e_i \tag{10}$$

Here SUB equals 0 if margin-eligibility date i occurred in the first subsample and 1 if it was in the second subsample. The first key coefficient in Eq. (10) is the intercept a, which indicates whether the change in spread is of the hypothesized sign in the first subsample; it is thus referred to as Intercept I. The second key coefficient is b_5, which indicates whether the impact of margin-eligibility significantly differs between the two subsamples. Because the sum of the coefficients for the intercept and SUB equals the value of the intercept for the second subsample, we refer to the sum as Intercept II.

The regression results for Eq. (10) are shown in Panel (B) of Table 6. Examination of the estimated coefficients associated with SUB reveals no significant difference in spreads between subsamples after adjusting for changes in the control variables. Furthermore, the coefficients for the intercept indicate that the change in spread is not significant in any of the four cases. Furthermore, there is an insignificant change in depth and the number of market-makers in both subsamples. In sum, Panel (B) of Tables 5 and 6 indicate that the 1997 reforms have had little impact. That is, margin-eligibility does not have an impact on a stock's liquidity either before or after the reforms of 1997. It is tempting to make the case that these reforms have improved liquidity, as afterwards there is a greater relative decrease in spreads and an increase in the number of market-makers. However, given the magnitude of the coefficients and their general lack of significance, these changes are not economically material.

5.2. Spread components

Table 7 indicates that, relative to the control sample, the increase in the adverse-selection component and the decrease in the order-processing component of the spread associated with margin-eligibility are notable and significant in the first subsample. However, neither spread component in the second subsample has a significant change. Similar to the subsample analysis in Table 7, the cross-sectional regressions reported in Table 8 indicate that the first subsample results are similar to the full sample results, but that the second subsample results were not significant after adjusting for changes in the control variables. Thus, it appears that the 1997 reforms have largely removed the impact of margin-eligibility on the components of the spread.

5.3. Market volatility

While not reported, we also used Eqs. (3)–(5) to analyze changes in volatility for each subsample. Of most interest is the first subsample, because there are two key features of this sample that eases comparisons with the time period studied by Hardouvelis and Theodossiou (2002). First, both the first subsample and the period studied by Hardouvelis and Theodossiou (2002) occur before the 1997 reforms (although their time period started and ended much earlier than our time period). Second, Hardouvelis and Theodossiou (2002) argue that the inverse relation between margin levels and volatility occurs only during bull (or normal) markets. During our first subsample (1993–1997), annual arithmetic returns to the Russell 2000, Wilshire 5000, and S&P 500 were 15.1%, 17.5%, and 18.2%, respectively. However, our estimate of the impact of a reduction in margin levels is negative, but insignificant at even the 10% level, as the estimated coefficient associated with MARDUM was − 0.8149 with a standard error of 0.5204. We interpret this last finding as "out-of-sample" refutation of the predictions of Hardouvelis and Theodossiou (2002).

6. Conclusions

In this study, we investigate the impact of margin-eligibility on the market quality of the affected equities. First, we find that the market liquidity of stocks is unchanged when they become margin-eligible, as there are insignificant changes in spreads, depths, and the number of market-makers. Second, we find that the portion of the spread that is attributable to adverse selection increases after margin-eligibility, suggesting there are more informed investors, while the portion attributable to order-processing costs decreases. While this increase should typically lead to a widening of spreads, the observed increase in trading volume appears sufficient to offset the increase in adverse selection so that, on balance, spreads are unaffected. In sum, once firms become margin-eligible, the quality of the market for their shares improves.

These findings suggest that margin-eligibility has no impact on a firm's cost of capital, as liquidity is unaffected. However, the observed increase in the information content of market prices has potential repercussions across a broad swath of other corporate finance dimensions. Examples include the public vs. private financing decision, the form of managerial compensation, the type of ownership structure, and the degree of shareholder monitoring. However, the recent introduction of two regulatory reforms—the introduction of teenies and the implementation of Order Handling Rules—appear to have eliminated the impact of margin-eligibility on all variables under consideration. Hence, on balance, a reduction in margin is a non-event for corporations in the current regulatory environment.

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