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# The liquidity effects of revisions to the S&P 500 index: an empirical analysis ☆

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**Abstract**

We study liquidity effects following S&P 500 index revisions. Using a recent sample of S&P 500 additions, we find a sustained increase in the liquidity of the added stocks. Further, the improvement in the liquidity of added stocks is due primarily to a decrease in the direct cost of transacting and a smaller decline in the asymmetric information component. Finally, the event period cumulative abnormal returns for additions are significantly associated with the decrease in the effective spread, particularly the decline in the direct cost of transacting. In contrast, the liquidity of deleted stocks declines over the three months following deletion.



*JEL classification:* G10

*Keywords:* Liquidity; Asymmetric; Information; Index; Additions; Deletions

## 1. Introduction

We examine the market liquidity effects on a recent sample of stocks added to and deleted from the Standard and Poor's (S&P) 500 index. Our first finding is a significant and sustained improvement in the liquidity of added stocks; specifically, a

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decline in the mean (and median) time-weighted quoted and effective spreads (both absolute and relative). The improvement in liquidity occurs not only during the period covering the announcement and addition of the stocks, but it persists over the subsequent three months of trading. We also find a significant increase in the mean (and median) quoted dollar depth, trading volume, and trade frequency over the three-month post-addition period. Thus, we document a long-term improvement in liquidity (i.e., an improvement that persists over the first three-month period following addition). Further analysis indicates that the improvement in trading costs stems mainly from a significant decrease in the direct cost of trading and, to a lesser extent, from a decrease in the asymmetric information cost of trading. Moreover, we present evidence that the widely documented abnormal returns following the addition of a stock to the S&P 500 index are associated with a decrease in the effective spread, particularly, a decrease in the direct cost of trading. Finally, we find significant deterioration in the liquidity of stocks that are deleted from the index during our study period.

Our findings are important for at least three reasons. First, we are the first to provide a comprehensive analysis of changes in the order flow and trading costs of stocks as they enter (and exit) the S&P 500 index. Shleifer (1986) observes that the addition of a stock to the index may promote closer scrutiny of the firm by analysts and investors, which, in turn, may lead to greater institutional interest, larger trading volume, and lower bid–ask spreads. Consistent with this prediction, Beneish and Whaley (1996), among others, find a permanent (i.e., long-term) increase in trading volume but only a temporary decrease in the quoted spread. Consequently, they rule out improved liquidity as an explanation for the widely documented abnormal returns on newly listed stocks in the S&P 500 index. Their findings seem inconsistent because the literature on liquidity predicts that, all else equal, a permanent increase in trading volume (due to liquidity-motivated trading) should be associated with a permanent, not a temporary, decrease in the bid–ask spread (Stoll, 1978). One possible reason that Beneish and Whaley find only a temporary decrease in the (equally weighted) quoted spread despite a permanent increase in trading volume is perhaps due to the fact that the quoted spread is not quite reliable in measuring actual trading costs. It is well known that the quoted spread overstates the true cost of transacting because many trades are executed inside the quotes (Petersen and Fialkowski, 1994; Madhavan et al., 1997).

To extend this analysis, we use alternative measures of liquidity costs, such as time-weighted quoted spread, effective spread, and time-weighted quoted depth in shares and in dollar terms. Based on these refined and more comprehensive liquidity measures, we find a sustained decrease in trading costs and an increase in quoted depth (in dollars) for stocks added to the S&P 500 index. These findings are consistent with the widely documented persistent increase in trading volume following index addition. The improvements in trading costs and depth persist even after controlling for the effects of changes in trading volume, stock price, and variance of returns. In contrast, stocks that are deleted from the index suffer a decrease in liquidity.

Second, we investigate whether the improvement in liquidity for additions is due to a decline in the direct or the asymmetric information cost of trading. After addition, index funds and investors will own a sizeable portion of shares outstanding in the added stock. According to one estimate, there were over $800 billion invested in mutual funds that tracked the S&P 500 index as of December 1998, which accounted for 8% of the market value of underlying securities.[1] Since most index investors seek to replicate the return performance of the index and minimize their tracking error (rather than "beat the market"), their trading activity will be driven by passive portfolio rebalancing considerations rather than by private information. Our empirical results, derived from the models of Glosten and Harris (1988), the Foster and Viswanathan (1993) adoption of Hasbrouck (1991), as well as Madhavan et al. (1997), show a significant decrease in the direct cost and a slight decline in the asymmetric information cost of trading from the pre- to the post-addition period.

Our third major contribution concerns the liquidity-wealth-effect hypothesis of Amihud and Mendelson (1986), under which if the addition of a stock in the S&P 500 index leads to a decline in expected trading costs, the stock price should permanently rise. Consistent with this prediction, we find that the cumulative abnormal returns associated with the addition of stocks to the index are positively related to the improvement in the liquidity costs of those stocks, especially with respect to the direct cost of trading. In contrast, Beneish and Whaley (1996) uncover only a temporary decrease in the quoted bid–ask spread in their sample and rule out the liquidity hypothesis as an explanation for the abnormal returns surrounding the entry of stocks to the S&P 500 index. While we do find a significant relationship between the improvement in liquidity and abnormal returns, the observed improvement in liquidity can only explain a fraction of the abnormal return and thus is not the only source of abnormal returns due to S&P 500 addition.

The paper is organized as follows. In Section 2 we discuss the relevant existing literature and motivate our empirical predictions. Section 3 details the data and methodology used in the study, while Section 4 presents the analysis of short-term effects of S&P 500 additions. We analyze the long-term changes in liquidity in Section 5 and focus on long-term changes in the direct and asymmetric information costs of transacting due to addition in Section 6. Section 7 investigates the relationship between abnormal returns and the long-term changes in liquidity. Section 8 presents brief discussions of our findings with respect to deletions. Our conclusions are presented in Section 9.

## 2. Related research and empirical predictions

The liquidity effects due to the addition of a stock to the S&P 500 index as well as the sources of the well-documented wealth effect remain an open empirical research question. This section begins with a review of the literature and is followed by a discussion of alternative explanations for the liquidity effects of S&P 500 index

[1] See Barron's, December 28, 1998, p. MW4.

additions. We follow the existing literature by focusing our study primarily on S&P additions rather than deletions.[2] Although the discussion that follows is in terms of additions to the index, the empirical predictions for deletions should be readily apparent. We return to the issue of deletions in Section 8.

### 2.1. Previous research

It is well established in the literature that when a stock is added to the S&P 500 index, it earns positive abnormal returns on average of approximately 3% and experiences increased trading volume. Previous studies have offered various explanations for the observed increase in price and trading volume. Harris and Gurel (1986) attribute these findings to the *price-pressure hypothesis*, which predicts a temporary increase in the stock's price and trading volume due to index fund purchases of the added security. They find the change in volume is permanent but the change in price is reversed over time. However, Beneish and Whaley (1996), Dhillon and Johnson (1991), Lynch and Mendenhall (1997), and Shleifer (1986) report that the wealth effect due to addition to the S&P 500 index is permanent. This evidence supports the *imperfect substitutes hypothesis* of Scholes (1972) and Shleifer (1986). Under this hypothesis, stocks are characterized by a downward sloping, rather than a perfectly elastic (i.e., horizontal), demand curve. A downward sloping demand curve is inconsistent with general notions of efficient markets in which stock price is based on fundamental asset valuation. Another explanation, the *information hypothesis*, suggests the Standard & Poor's Corporation uses nonpublic information in selecting stocks for addition to the index. Since the addition of a stock is assumed to be good news, the price should register a permanent rise on the announcement of a stock's inclusion in the index. While Jain (1987) and Dhillon and Johnson (1991) find evidence in support of the information hypothesis, both Harris and Gurel (1986) and Shleifer (1986) present evidence inconsistent with this hypothesis.

Erwin and Miller (1998) use a sample of 109 additions over 1984–1988 and find a long-term (up to 30 days following index addition) decrease in daily quoted spreads on added stocks without listed options. Beneish and Whaley (1996) study not only trading volume but also trade size and quoted bid–ask spread in a sample of 103 additions over 1989–1994 and thus provide the most detailed analysis of liquidity effects available in the current literature. They find a long-term increase in trading volume that lasts up to at least 60 days following the effective day of addition and a significant increase in average trade size for the first 10 days after addition. However, the average quoted bid–ask spread in their sample drops significantly only on the day after addition.[3] Since the quoted bid–ask spread returns to the normal (i.e., pre-index addition) level after demand from index funds subsides, they rule out the liquidity

[2] A notable exception is Lynch and Mendenhall (1997) who study S&P 500 index deletions as well as additions in their analysis of the abnormal return patterns following a recent change in announcement policy by Standard and Poor's.

[3] In October 1989, the S&P changed its announcement policy with regard to changes in the composition of the index from just a day to five days prior to the effective day of addition. Beneish and Whaley (1996) find a significantly lower quoted bid–ask spread over the first 10 days following the announcement in the

hypothesis as an explanation for the observed wealth effect. The focus of their research, however, was not liquidity, but rather how changes in S&P's announcement policy introduced a risk arbitrage strategy called the "S&P game". In contrast to the previous research, our aim is to more fully investigate the liquidity effects of S&P 500 additions and liquidity-based explanations for the observed wealth effect.

### 2.2. S&P 500 index revisions and liquidity

The addition of a stock to the S&P 500 index is associated with many significant changes in its ownership and trading activity that affect its market liquidity. By far the most important of these changes is a shift in ownership composition to uninformed index investors. Therefore, we expect an increase (decrease) in liquidity for stocks added to (deleted from) the index. Other potential significant changes that can impact liquidity include a change in the character of order flow, increased analyst following, increased arbitrage trading activity, and order flow migration of liquidity traders from the stock market to the market for index products. The intent of our study is not is to discern between these competing explanations for the liquidity effect as they are not mutually exclusive and are in some cases observationally equivalent. We do, however, briefly discuss the hypothesized effect on liquidity for each aforementioned change.

The most widely recognized change following index addition is the change in ownership composition—the purchase of that stock by index funds and index investors (Lynch and Mendenhall, 1997). This marks a sizeable and long-term shift in the composition of equity ownership to uninformed index investors (Beneish and Whaley, 1996). If the inclusion of a stock in the S&P 500 index results in an increase in the proportion of liquidity-motivated trades and keener competition among informed traders in the stock, the Kyle (1985) type microstructure models suggest an improvement in the liquidity of the stock.

The change in ownership composition affects the character of trading activity in the added stock. However, a priori, the impact of index addition on trading activity is not clear. Trading costs (specifically, the asymmetric information cost of trading) may increase if the variance of uninformed trades and the frequency of trades were to drop due to the tendency of index investors to buy and hold. Trading costs may decrease, however, if the trading activity of index investors, which is driven primarily by liquidity shocks and portfolio rebalancing to minimize their tracking error, results in smaller and less frequent trades. Further, the onset of index arbitrage trading for the stock maybe expected to result in smaller and more frequent trades (Sofianos, 1993). Smaller, more frequent, and less variable trades serve to reduce the inventory concerns of the market maker and improve liquidity by reducing the direct cost of trading.

(*footnote continued*)
pre-October 1989 period, but the drop in the quoted spread in their sample is significant only for the day after addition in the post-October 1989 period.

With the inclusion of a stock in the index, its liquidity may improve due to an increase in the number of analysts (Brennan and Subrahmanyam, 1995) following the stock as well as in the degree of market monitoring. Further, the increased visibility of the stock associated with index addition is likely to increase the flow of public information about the stock. Also, as suggested by Easley et al. (1998), these factors may attract more uninformed (i.e., "noise") trade and improve the liquidity of the stock.

Moreover, the added stock becomes effectively "cross-listed" in the market for the stock index as well as in the index derivative markets. This creates alternative trading avenues (i.e., exchange-traded funds and index derivatives) for those investors who trade stock portfolios. As the intuition of Gammill and Perold (1989) suggests, if order flow migrates from the underlying stock market to the associated markets for stock index and/or index derivatives following addition, then the liquidity of the stock will suffer. Models of basket trading (e.g., Subrahmanyam, 1991) and international cross listing (e.g., Domowitz et al., 1998) make it clear that the resulting liquidity effects of "cross-listing" are more complex than intuition suggests and are not easy to predict.

Yet another factor is the change in arbitrage activity in the added stock. Kumar and Seppi (1994) study index arbitrage between the cash and index (S&P 500) futures markets with imperfect information linkages under which new information is locally available in one market first and is transmitted to the other market through trading with a time lag. Their model implies that arbitrage effects on market liquidity from index addition are complex. Holden (1995) suggests index arbitrage between the index shares, futures and options markets and the underlying stocks improves liquidity by eliminating information asymmetries across markets and by providing buying and selling support to correct temporary order imbalances across markets.[4]

### 2.3. *Liquidity-wealth-effect hypothesis*

Amihud and Mendelson (1986) posit that liquidity costs increase the required rate of return on securities. Their model implies that if the inclusion of a stock in the S&P 500 index decreases its transaction costs, then the stock price will rise. More recently, Brennan and Subrahmanyam (1996) present empirical evidence in support of this model and find that equity returns vary directly with the illiquidty costs due to the direct and asymmetric information costs of transacting. Therefore, if the addition of a stock to the S&P 500 index leads to an overall improvement in market quality, then at least part of the previously documented abnormal returns should be positively related to the decline in the effective spread as well as the decline in the direct and/or asymmetric information costs of trading. We call this explanation of the well-documented wealth effect associated with S&P 500 additions the *liquidity-wealth-effect* hypothesis.

[4] Kumar et al. (1998) find the listing of options in individual stocks improves the liquidity and market quality of underlying stocks. Also, see Dhillon and Johnson (1991) and Erwin and Miller (1998).

## 3. Data and methodology

We need data on quotes, prices and order flow to examine changes in the market liquidity of stocks added to the S&P 500 index. Unfortunately, we observe neither the institutional equity ownership structure—the proportional ownership of passive and informed institutions—nor the type of incoming order—whether it is information-motivated or liquidity-motivated. However, the market microstructure models of Glosten and Harris (1988), Glosten and Milgrom (1985), Foster and Viswanathan (1993), Hasbrouck (1991), and Madhavan et al. (1997) allow us to infer the type of trade from the time series properties of transactions prices.

We examine a sample of stocks added to the S&P 500 index during January 1, 1993 through October 31, 1998. We obtained information on the announcement and effective days of listing from the Standard and Poor's Corporation. There were 145 changes to the index during this period. Our final sample consists of firms that meet the following criteria:

(a) The firm is not involved in a merger that immediately preceded its addition to the S&P 500 index.
(b) The firm has available historical data on the New York Stock Exchange (NYSE) Transactions and Quote (TAQ) database for a period commencing 60 trading days before and ending 60 trading days after the effective addition date.[5]
(c) The firm has available historical data on the Center for Research in Securities Prices (CRSP) database for the period commencing 250 trading days before and ending 60 trading days after the effective addition date.
(d) The firm's common stock did not split in the period commencing 60 trading days before and ending 60 trading days after the effective addition date. Specifically, the shares outstanding did not change by greater than 25% in the period.

Criteria (a) and (d) are applied to minimize the effects of confounding events. This selection process yields a sample of 91 additions: 74 are NYSE listed firms and 17 trade on National Association of Securities Dealers Automated Quotations (NASDAQ). We decided to exclude the 17 NASDAQ firms from the sample for two reasons. First, the spread decomposition methodologies employed in this study are more appropriate for an auction market (i.e., NYSE) as opposed to dealer market (i.e., NASDAQ) structure. Second, the application of spread decomposition methodologies to TAQ data for NASDAQ trades is problematic due to the presence of inter-dealer trades.[6] The same selection criteria applied to the corresponding S&P 500 index deletions yields a sample of 27 NYSE listed firms.

[5] We found little or no quote and depth data for October 1994 in the TAQ database. This affected 5 firms in our final sample. Rather than delete the firms from our sample, we omitted the October 1994 observations.

[6] When the abnormal return and volume event study is conducted with the entire sample of 91 firms, the event study results remain qualitatively similar to those reported in this study.

The stock market data primarily comes from two sources. For the event study, the daily return data for individual stocks and the market proxy (i.e., CRSP value-weighted index) are taken from the CRSP daily master file for 1998. The data on shares outstanding and daily total trading volume are also obtained from CRSP. The trade and quote data are taken from the NYSE TAQ database. We discard the first trade of each day (and thus overnight price changes) in order to avoid the effects of market opening procedures, overnight news arrivals, and dividends. The daily NYSE trading volume was obtained from the NYSE. Only primary exchange (i.e., NYSE) data are included in the analysis. We do not have access to data on the proportional equity ownership of passive and informed institutions. Instead, we obtain quarterly data on aggregate institutional ownership from the Value Line Investment Survey. As is customary in the microstructure literature (e.g., Brennan and Subrahmanyam, 1995), analyst following is proxied by the number of quarterly earnings estimates from the I/B/E/S summary history.

### *3.1. Descriptive statistics*

Descriptive statistics for the final sample of 74 additions and 27 deletions to the S&P 500 index are presented in Table 1. From Panel A, while the number of index changes has increased over time, the percentage of firms retained in the additions sample is fairly constant across years at about 50%. Of the 74 additions retained in the sample, 64 were listed in the index on independent dates, thus minimizing the problem of clustering of event dates. Panel B presents summary statistics on the characteristics of the additions sample estimated over 121 trading days surrounding the effective day of addition (i.e., event day 0). The firms added to the S&P 500 index are large (median capitalization of $5.35 billion) and are traded actively (median average daily volume of 525,213 shares). The intraday analysis of additions uses 2.5 million quotes and 1.8 million trades. The last three columns of Table 1 present common measures of trading costs and depth derived from intraday data. The reported quoted spread is the daily time-weighted average of the difference between ask and bid quotes for the added stock over the extended (i.e., six month) event period. The weights are based on the length of time over which each quote is valid as a fraction of time between the daily opening trade and the close of trading. Quoted relative spread is defined as the dollar quoted spread divided by the quote midpoint. We examine both the absolute (i.e., dollar) and relative (i.e., percent) spreads to ensure that the change in relative spread is not simply due to a change in the price level of a stock following its listing in the market index. Depth denotes the total number of shares offered by the market maker at the bid–ask quotes. Like the dollar spread, both the quoted relative spreads and depth are daily time-weighted averages. The estimates in Panel B indicate the newly listed firms have a median quoted spread of $0.15 with a median quoted depth of 127 round lots. From Panel C, only 27 firms remain following our screening procedures for index deletions. Compared with the additions sample, the deleted firms are on average considerably smaller ($929 million vs. $6.0 billion), trade less frequently (8,156 vs. 24,876 trades over the sample period), have higher relative quoted spreads (1.60% vs. 0.41%), and greater depth

Table 1

Descriptive statistics

The final sample consists of 74 (27) NYSE listed firms that were added to (deleted from) the S&P 500 index from 1993 to 1998. Firms retained in the final sample did not have confounding events (e.g., splits, mergers, etc.) and have complete data available in the TAQ as well as CRSP databases. The data on additions and deletions are obtained from Standard and Poor's Corporation.

*Panel A: Distribution of S&P 500 index revisions by calendar year*

| Year | Total No. of S&P 500 changes | No. of additions retained in sample | No. of independent dates for additions | No. of deletions retained in sample | No. of independent dates for deletions |
|---|---|---|---|---|---|
| 1993 | 8 | 3 | 3 | 2 | 2 |
| 1994 | 15 | 7 | 5 | 4 | 4 |
| 1995 | 30 | 15 | 14 | 8 | 7 |
| 1996 | 21 | 15 | 12 | 6 | 5 |
| 1997 | 23 | 13 | 11 | 3 | 3 |
| 1998 | 48 | 21 | 19 | 4 | 4 |
| Total | 145 | 74 | 64 | 27 | 25 |

*Panel B: Descriptive statistics for index additions*

Daily average closing price, market capitalization, NYSE trading volume in shares, daily number of NYSE quotes and trades, daily time-weighted average of quoted spread (relative and absolute) and depth in shares for the index additions are drawn from the 121 trading day interval [−60, +60], centered on the effective day of addition. Price, capitalization, and volume data are taken from the CRSP daily master file. Data on quotes, trades, spread, and depth are obtained from the NYSE TAQ database

| | Average closing price ($) | Market value ($000) | Daily NYSE volume (shares) | Total No. of NYSE quotes | Total No. of NYSE trades | Quoted relative spread (%) | Quoted spread ($) | Quoted depth (00 shares) |
|---|---|---|---|---|---|---|---|---|
| ($n = 74$) | | | | | | | | |
| Mean | 43.75 | 6,034,637 | 639,086 | 33,898 | 24,876 | 0.41 | 0.1571 | 186 |
| Median | 39.17 | 5,350,185 | 525,213 | 29,782 | 20,918 | 0.38 | 0.1510 | 127 |
| Maximum | 109.22 | 18,093,665 | 2,175,781 | 109,586 | 93,148 | 1.50 | 0.2455 | 1,522 |
| Minimum | 8.32 | 720,460 | 116,462 | 5,839 | 7,315 | 0.17 | 0.0954 | 25 |
| Total No. of trades/quotes | | | | 2,474,542 | 1,840,842 | | | |

Table 1 (*continued*)

*Panel C: Descriptive statistics for index deletions*

Daily average closing price, market capitalization, NYSE trading volume in shares, daily number of NYSE quotes and trades, daily time-weighted average of quoted spread (relative and absolute) and depth in shares for the final sample of 27 S&P 500 index deletions are from the 121 trading day interval [−60, +60], centered on the effective day of deletion. Capitalization data are taken from the CRSP daily master file. The remaining data are obtained from the NYSE TAQ database

| | Average daily closing price ($) | Market value ($000) | Daily NYSE volume (shares) | Total No. of NYSE quotes | Total No. of NYSE trades | Quoted relative spread (%) | Quoted spread ($) | Quoted depth (00 shares) |
|---|---|---|---|---|---|---|---|---|
| ($n = 27$) | | | | | | | | |
| Mean | 17.47 | 928,547 | 2,195,848 | 14,030 | 8,156 | 1.60 | 0.1450 | 268 |
| Median | 13.86 | 372,231 | 1,678,100 | 8,101 | 5,586 | 1.00 | 0.1331 | 203 |
| Maximum | 76.15 | 7,108,663 | 8,931,400 | 49,145 | 46,039 | 5.18 | 0.3212 | 894 |
| Minimum | 2.54 | 82,500 | 352,700 | 3,711 | 2,000 | 0.21 | 0.0807 | 38 |
| Total No. of trades/quotes | | | | 378,823 | 220,200 | | | |

(268 vs. 186 round lots). The intraday data used for the deletions sample numbers 220,200 trades and 379,823 quotes.

### 3.2. Event study methodology

We employ the standard event study methodology as outlined in Campbell et al. (1997) to estimate abnormal returns and to verify that our sample replicates the S&P 500 addition results presented in previous studies. We define the event (i.e., day $t=0$) as the effective day of adding a stock to the S&P 500 index. The announcement procedure for adding firms to the index that was in effect during the sample period was first implemented in October 1989. Under this policy, the S&P Corporation announces that a firm will be added (deleted) to (from) the S&P 500 index after the close of trading on, generally, the 5th trading day prior to the effective day. In other words, there will generally be four trading days between the announcement and the open of trading on the effective day of revision. The mean time period between the announcement date and the effective day of addition is 3.35 trading days and the median is 4 trading days. There are 11 occurrences of no lag; that is, the first day of trading following the announcement is also the effective day of addition. There are three occurrences of lags greater than 10 trading days.

We estimate the market model parameters over the interval $[-250, -30]$. The abnormal returns for each security will be aggregated through event time to obtain cumulative abnormal returns (CAR). We then aggregate CARs across securities to compute the average cumulative abnormal return (ACAR). The standard parametric test assumes the abnormal returns are jointly normally distributed, their mean and variance are not affected by the event, and they are independent across securities. In addition, we also use the large sample approximation of the standard sign test to check the robustness of our parametric tests of abnormal returns.

## 4. Empirical results over the event window

This section focuses on the short-term effects of revisions to the S&P 500 index. Abnormal returns and changes in liquidity around index revision, $[-5, +5]$, are investigated using the standard event study methodology. Previous work in this area has found positive (negative) abnormal returns and increased trading volume for the added (deleted) stock in the period surrounding the announcement and effective days of revision. We complement this evidence by analyzing the effects of addition on quoted time-weighted spread, depth, and trade size.

### 4.1. Abnormal returns

The results for the event window $[-5, +5]$ are presented in Panel A of Table 2. For additions, the first evidence of statistically significant average abnormal returns (AAR) arises on event day $[-4]$, which is generally the first day of trading after the announcement of the addition of a stock to the index. Significant positive AARs

Table 2
Event window abnormal returns and trading volume for S&P 500 additions

*Panel A*
The sample consists of 74 NYSE listed firms that were added to the S&P 500 index during the period 1993–1998. Average abnormal returns (AAR) and average cumulative abnormal returns (ACAR) are calculated using the market model and the standard event study methodology. The estimation window for calculating the market model parameters is the event time interval [−250, −30]. AAR and ACAR are both tested for significance using a *t*-test and a standard sign test.

| Event day/interval | AAR (%) or ACAR (%) | Proportion with AAR or ACAR > 0 | *t*-Test $H_0$: AAR = 0 or ACAR = 0 | Sign test for AAR and ACAR |
|---|---|---|---|---|
| −5 | 0.10 | 0.46 | 0.36 | −0.70 |
| −4 | 1.61 | 0.72 | 5.75*** | 3.72*** |
| −3 | 0.78 | 0.57 | 2.78*** | 1.16 |
| −2 | 1.33 | 0.59 | 4.76*** | 1.63 |
| −1 | 1.69 | 0.69 | 6.05*** | 3.25*** |
| 0 | 2.30 | 0.68 | 8.24*** | 3.02*** |
| +1 | −0.74 | 0.27 | −2.63*** | −3.95*** |
| +2 | −0.64 | 0.32 | −2.30** | −3.02*** |
| +3 | −0.47 | 0.39 | −1.67* | −1.86* |
| +4 | −0.23 | 0.39 | −0.84 | −1.86* |
| +5 | −0.08 | 0.39 | −0.29 | −1.86* |
| [−1, +1] | 3.26 | 0.68 | 3.90*** | 3.02*** |
| [−2, +2] | 3.95 | 0.73 | 2.83*** | 3.95*** |
| [−3, +3] | 4.26 | 0.73 | 2.18** | 4.18*** |
| [−4, +4] | 5.63 | 0.84 | 2.25** | 5.81*** |
| [−4, 0] | 7.72 | 0.95 | 5.53*** | 7.67*** |
| [+1, +4] | −2.08 | 0.31 | −1.87* | −3.25*** |

*Panel B: Abnormal volume around S&P 500 addition*

A pooled time series cross-sectional regression is estimated with the White (1980) covariance matrix to investigate the volume patterns surrounding the dates of announcement and addition of stocks to the S&P 500 index as follows:

$$\log Volume_{jt} = \mathbf{i}\alpha_j + t\gamma + \sum_{i=-5}^{+5} D_i\beta_i + \varepsilon_j \quad \text{for } j = 1, 74 \text{ and } t = -250, +5,$$

where log $Volume_{jt}$ is the log of trading volume in shares for stock $j$ and day $t$, $\mathbf{i}$ is a $74 \times 74$ identity matrix, $D_i$ are dummy variables for each trading day in the event time interval $[-5, +5]$, $\varepsilon_j$ is a zero mean heteroscedastic disturbance term, and $\alpha_j$, $\gamma$, and $\beta_i$ are parameters to be estimated. The estimates for the linear drift ($\gamma$) and event period dummy variables ($\beta_i$) for the interval $[-5, +5]$ are reported with $t$-statistics in parentheses below

Parameter estimates

| $\gamma$ | $\beta_{-5}$ | $\beta_{-4}$ | $\beta_{-3}$ | $\beta_{-2}$ | $\beta_{-1}$ | $\beta_0$ | $\beta_{+1}$ | $\beta_{+2}$ | $\beta_{+3}$ | $\beta_{+4}$ | $\beta_{+5}$ |
|---|---|---|---|---|---|---|---|---|---|---|---|
| 0.0019 | 0.1886 | 0.4819 | 0.5848 | 0.7182 | 0.9969 | 2.7412 | 1.3608 | 0.7685 | 0.6037 | 0.5230 | 0.3868 |
| (25.99) | (1.90) | (4.02) | (6.27) | (9.17) | (11.54) | (31.33) | (18.44) | (10.36) | (7.47) | (5.89) | (5.40) |

Adjusted $R^2 = 0.61$

*Significant at 10% level.
**Significant at 5% level.
***Significant at 1% level.

persist over the next four days, with the largest average return of 2.30% occurring on the effective date of addition (i.e., event day 0). Of the 74 firms in the sample, 50 firms have positive abnormal returns on the effective day of addition. There is some evidence of the reversal of positive excess returns in the post-addition period, but not enough to eliminate the gains accrued during the event interval [−4, 0]. The average cumulative abnormal return over the nine trading day interval [−4, +4] is 5.63%, with 62 of the 74 firms in the sample experiencing positive abnormal returns.[7] These results are consistent with the pattern of returns noted by previous studies, notably Beneish and Whaley (1996) and Lynch and Mendenhall (1997).

### 4.2. Trading volume

Another frequent finding of previous work on the effect of S&P 500 addition is a substantial increase in trading volume in the event period presumably due to the portfolio rebalancing trades of index investors as well as the activities of risk arbitrageurs. In testing for abnormal volume in the event period, we control for the likely increase in trading volume over time.[8] We employ the following least squares dummy variable model for pooled times series regression.[9]

$$\log \mathit{Volume}_{jt} = \mathbf{i}\alpha_j + t\gamma + \sum_{i=-5}^{+5} D_i\beta_i + \varepsilon_j \quad \text{for } j = 1, 74 \quad \text{and} \quad t = -250, +5, \tag{1}$$

where log $\mathit{Volume}_{jt}$ is the log of trading volume in shares for stock $j$ and day $t$, $\mathbf{i}$ is a $74 \times 74$ identity matrix, $D_i$ are dummy variables for each trading day in the event time interval [−10, +10], $\varepsilon_j$ is a zero mean heteroscedastic disturbance term, and $\alpha_j$, $\gamma$, and $\beta_i$ are parameters to be estimated. The intercept parameters, $\alpha_j$, differ across securities and capture the variation in average trading volume across securities. The linear drift term, $\gamma$, estimates the average change in trading volume per trading day. The coefficients of the series of eleven dummy variables, $\beta_i$, capture abnormal volume for each trading day in the event period [−5, +5]. Eq. (1) is estimated for our sample of additions using the White (1980) heteroscedastic consistent covariance matrix; the results are presented in Panel B of Table 2. The results confirm the dramatic increase in trading volume reported by earlier studies. For example, on the typical announcement day (i.e., event day −4), the coefficient on the dummy ($\beta_{-4}$) is 0.4819 and is highly significant ($t = 4.02$), indicating trading volume is approximately 1.6 times higher than normal after controlling for the linear drift in volume over time. The abnormal volume continues to rise and reaches its maximum on the effective day of addition (i.e, event day 0). On this day, the estimate of the dummy

[7] The median cumulative abnormal return is 5.69% (not reported in Table 2). Also, for this nine-day event window [−4, +4], the cumulative abnormal returns for individual firms range from a min. of −15.64% to a max. of +24.56%.

[8] We thank the referee for this suggestion.

[9] We also estimate Eq. (1) separately for each stock in the additions sample thus allowing for different drift terms and slope coefficients across stocks. The results of these regressions are consistent with the pooled approach as 73 of 74 stocks in the sample exhibit statistically significant abnormal volume on the day of addition.

($\beta_0$) is 2.7412 and highly significant ($t = 31.33$) indicating trading volume is approximately 15 times greater than normal. Following addition, the abnormal volume subsides from its peak but remains significant throughout the event period. However, the analysis of aggregate trading volume alone cannot tell us whether the observed increase is due to liquidity—or information—motivated trading. Therefore, we turn to the analysis of trading costs and market depth.

### 4.3. *Quoted liquidity*

We construct ratios of the daily average quoted absolute and relative time-weighted spreads over various event intervals to their counterparts in the pre-addition period over trading days [−60, −10]. Consistent with Branch and Freed (1977), we focus on the relative quoted spread since it captures the economic significance of the spread to the dealer. We realize, however, that the relative spread of a stock will change, in spite of a constant dollar spread, solely due to an increase in its price after the addition. Therefore, we also analyze the absolute quoted spread defined as the difference between the quoted ask and bid prices in dollar terms. In addition, consistent with Lee et al. (1993) we examine quoted depth to complement the spread analysis.

From Table 3, the time-weighted quoted absolute and relative spreads show a clear decrease over the event window. For example, in the [−4, +4] interval the daily mean and median absolute quoted spread ratios are both 0.95 and highly significant, indicating that spreads are lower over the nine trading day period centered on the effective day of addition. For the average quoted relative spread over the same interval, the decrease is more pronounced due to higher stock prices in the post-addition period and is highly significant under the standard *t*-test and sign test. Further, the decrease in spreads, both absolute and relative, is widespread. For instance, in the post-addition interval [+1, +4], only 12% (18%) of the sample stocks experience elevated quoted absolute (relative) spreads. The greatest observed decrease in spreads occurs in the period following addition [+1, +4].

In addition, we analyze the ratios of time-weighted quoted depth (number of shares offered by the market maker at the bid and ask quotes) and trade size, which are constructed along the same lines as the spread ratios. From column 6 of Table 3, the average depth is significantly greater over the event period as compared to the pre-addition time interval [−60, −10]. For example, in the [−4, +4] interval, the mean depth is 2.49 times greater than normal, and all but two stocks in the sample experience an increase in depth. The last two columns show that, along with the increased trading volume and depth, the average trade size also increases in the event period. The significant increase in average trade size during the event window is consistent with the discussion of the temporary "S&P game" examined in Beneish and Whaley (1996). The "S&P game" can be described as buying large blocks of the added stock between the announcement and addition dates and then selling them back at a potentially higher price to index funds in the post-addition period. Thus, the results on quoted spreads and depth are consistent with the conjecture that the

Table 3
Short-term (event window) effects of S&P 500 index additions on stock market liquidity
Liquidity is measured by the absolute and relative quoted spreads, quoted depth, and average trade size on the NYSE for the added stocks. The sample consists of 74 NYSE listed firms that were added to the S&P 500 Index during the period 1993–1998. The spread (depth) ratio for each security in the sample is defined as the ratio of time-weighted average spread (depth) measured over the indicated event time interval to time-weighted average spread (depth) measured over the pre-S&P 500 addition period [−60, −10]. The trade size ratio for a security is defined as the ratio of average trade size over the indicated event time interval to average trade size in the pre-S&P 500 addition period. Quoted spread is defined as ask price less bid price. Quoted relative spread is defined as quoted spread divided by quote midpoint. Quoted depth the sum of depth (number of shares) at the bid and ask prices. Data are taken from the NYSE TAQ database. The null hypothesis that the mean of the reported ratio is equal to unity is tested using a standard $t$-test.

| Event time interval | Ratios of event interval to pre-addition period liquidity measures | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Quoted spread ($) | | Quoted spread (%) | | Quoted depth | | Average trade size | |
| | Mean (median) | Proportion with ratio >1.0 | Mean (median) | Proportion with ratio >1.0 | Mean (median) | Proportion with ratio >1.0 | Mean (median) | Proportion with ratio >1.0 |
| [0, 0] | 1.07 (0.92)** | 0.35 | 0.96 (0.85)*** | 0.28 | 2.68*** (2.20)*** | 0.86 | 3.02*** (2.78)*** | 0.99 |
| [−1, +1] | 0.97 (0.91)*** | 0.28 | 0.95 (0.90)*** | 0.31 | 2.66*** (2.25)*** | 0.97 | 2.32*** (2.17)*** | 0.99 |
| [−2, +2] | 0.95*** (0.91)*** | 0.28 | 0.91*** (0.86)*** | 0.26 | 2.75*** (2.36)*** | 0.97 | 2.03*** (1.87)*** | 0.99 |
| [−3, +3] | 0.94*** (0.92)*** | 0.28 | 0.91*** (0.88)*** | 0.26 | 2.62*** (2.32)*** | 0.97 | 1.88*** (1.80)*** | 0.97 |
| [−4, +4] | 0.95*** (0.95)*** | 0.31 | 0.91*** (0.88)*** | 0.26 | 2.49*** (2.23)*** | 0.97 | 1.78*** (1.71)*** | 0.95 |
| [−4, −1] | 1.02 (0.98) | 0.46 | 0.98 (0.95)** | 0.36 | 1.96*** (1.53)*** | 0.80 | 1.47*** (1.40)*** | 0.80 |
| [+1, +4] | 0.86*** (0.85)*** | 0.12 | 0.84*** (0.80)*** | 0.18 | 2.96*** (2.44)*** | 1.00 | 1.49*** (1.37)*** | 0.88 |

**Significant at 5% level.
***Significant at 1% level.

inclusion of a stock in the S&P 500 is followed by an increase in the proportion of liquidity-motivated trading.

## 5. Long-term effects on returns and liquidity

What is of greater interest is not the short-term improvement in quoted liquidity but whether it will persist over time. In other words, we wish to check if the influx of index investors, the added visibility of the newly listed stock, increased intermarket arbitrage, and potential order flow migration lead to a long-term (i.e., permanent) improvement or deterioration in its liquidity. To investigate this issue, we compare several measures of spread, depth, and trading activity computed over the post-addition time interval $[+10, +60]$ with their counterparts over the pre-addition period $[-60, -10]$.[10] These time intervals are chosen to reduce the bias in liquidity measures due to the initial portfolio rebalancing of index fund managers in the days following the effective date of addition and any other possible announcement effects.

### *5.1. Long-term abnormal returns*

If sustained improvement in market liquidity is an important source of abnormal returns associated with the index addition of a stock, then we are less likely to observe price reversals to the pre-announcement level. Table 4 presents the evidence on long-horizon returns, ACARs (i.e., between trading days $-4$ and $+60$). Two findings are noteworthy. First, the results indicate significant average abnormal returns at least 10–20 trading days following the index inclusion of a stock, but not beyond. Since the added stocks are generally large in firm size and are actively traded (see panel B of Table 1), price increases due to excess demand for the added stocks and other transient trading frictions are unlikely to endure up to a full calendar month. Therefore, we interpret the observed persistence of abnormal returns as a sign of a lasting improvement in the value of added stocks. Second, the last four rows of Table 4 suggest strong price reversals relative to the market over the post-addition period. These reversals are consistent with the price pressure hypothesis. However, it is important to note that such long horizon tests have low power because of their greater sampling variability.[11] Overall, we find evidence of strong price reversals as well as price persistence in the post-addition period.

### *5.2. Long term abnormal returns and momentum*

Jegadeesh and Titman (1993) and Chan et al. (1996) find momentum in stock returns over long-term (i.e., six months to one year) horizons. It is likely that stocks added to the S&P 500 index are positive momentum stocks. If additions are indeed

[10] Throughout the study, we will use the terms ''pre-addition'' and ''post-addition'' periods to refer to the 51 trading day event time intervals $[-60, -10]$ and $[+10, +60]$, respectively.

[11] See Brown and Warner (1985) and also Chapter 4, Section 6 of Campbell et al. (1997).

Table 4
Long-term abnormal returns and momentum effects

*Panel A: Long-term abnormal returns*

The sample consists of 74 NYSE listed firms that were added to the S&P 500 index during the period 1993–1998. Average cumulative abnormal returns (ACAR) are calculated using the market model and the standard event study methodology. The estimation window for calculating the market model parameters is the event time interval [−250, −30]. The significance of ACARs over varying time intervals is examined using a *t*-test and a standard sign test. A two-tailed *t*-test is used to test whether abnormal returns accrued as of the addition date [−4, 0] are fully reversed following index addition ([+1, *T*], where $T = +10$, +20, +30, and +60)

| Event interval | ACAR (%) | Proportion with ACAR >0 | *t*-Test $H_0$: ACAR = 0 | Sign test for ACAR = 0 | *t*-Test: $H_0$: ACAR[−4,0] = −ACAR[+1,*T*]. |
|---|---|---|---|---|---|
| [−4, +10] | 4.41 | 0.73 | 1.05 | 3.95*** | |
| [−4, +20] | 1.91 | 0.64 | 0.27 | 2.32** | |
| [−4, +30] | 2.01 | 0.53 | 0.21 | 0.46 | |
| [−4, +60] | −1.09 | 0.38 | −0.06 | −1.86* | |
| [+1, +10] | −3.31 | 0.24 | −1.19 | −4.42*** | 4.61*** |
| [+1, +20] | −5.81 | 0.19 | −1.04 | −5.35*** | 1.57 |
| [+1, +30] | −5.70 | 0.27 | −0.68 | −3.95*** | 1.46 |
| [+1, +60] | −8.81 | 0.26 | −0.53 | −4.18*** | −0.61 |

*Panel B: Momentum and long-term abnormal returns*

A measure of momentum in the pre-addition period is formed as the buy and hold return for each stock in the additions sample in the period [−250, +30] less the buy and hold return from the CRSP value-weighted index over the same period. The sample is then divided into quartiles based on this measure of excess return. The mean of the cumulative abnormal returns (CAR) are reported for each pre-event excess return quartile and for the long-term event intervals

| Quartile based on pre-event period buy and hold excess returns | Mean of pre-event buy and hold excess returns (%) | Mean of CAR for indicated pre-event buy and hold excess return quartile and long-term event interval (%) | | | |
|---|---|---|---|---|---|
| | | [−4, +10] | [−4, +20] | [−4, +30] | [−4, +60] |
| 1st | 72.53 | −0.51 | −4.51 | −6.27% | −11.98% |
| 2nd | 23.01 | 3.86 | 0.47 | 0.40% | −2.54% |
| 3rd | 2.49 | 7.02 | 5.38 | 5.24% | 3.39% |
| 4th | −21.41 | 7.14 | 6.18 | 8.60% | 6.60% |
| Entire sample | 18.98 | 4.41 | 1.91 | 2.01% | −1.09% |

*Significant at 10% level.
**Significant at 5% level.
***Significant at 1% level.

positive momentum stocks, long-term post-addition abnormal returns may be a result of price momentum rather than the effects of addition.[12] We follow Ikenberry et al. (1996) to explore this issue. If the long-term cumulative abnormal returns noted for additions are due to price momentum, the momentum (or price run-up) in the

[12] We thank the referee for this suggestion.

pre-addition period and the cumulative abnormal returns should be positively related. We form a measure of price run-up in the pre-addition period [−250, −30] as the buy and hold return for each stock in our sample in excess of the buy and hold return for the market as proxied by the CRSP value-weighted index for the same time period. The mean price run-up for the sample is 18.98%. This suggests S&P 500 additions are, on average, positive momentum stocks. We then divide the sample into quartiles based on price run-up and compute the mean of each. The results are reported in Panel B of Table 4. It is noteworthy that the 1st quartile has a mean excess return of 72.53% while the 4th quartile has a mean excess buy and hold return of −21.41% suggesting that not all of the stocks in the addition sample are positive momentum stocks. The mean cumulative abnormal returns (CARs) are reported for the long-term event intervals for each price run-up quartile. The results in Panel B, without exception, support an inverse relationship between pre-addition price run-up and the abnormal returns in the event intervals [−4, +10], −4, +20], [−4, +30], and [−4, +40]. The observed inverse relationship suggests the long-term abnormal returns are not attributable to momentum.

### 5.3. Measures of trading activity, analyst coverage, and institutional ownership

In Table 5, we evaluate changes in daily standardized trading volume, daily trade frequency, average trade size, standard deviation of order flow, daily return volatility, proportion of crossing trades, proportion of block trades, institutional ownership, and analyst coverage from the pre-addition to the post-addition period. To analyze long-term changes in each of these measures following index addition, we form a "Post/Pre ratio" of each of these variables in the post-addition period [+10, +60] to the same variable in the pre-addition period [−10, −60] for each stock. The mean (median) Post/Pre ratio of standardized trading volume is 1.27 (1.15), and 64% of the stocks have ratios greater than 1.0. Both the standard *t*-test and the sign test indicate that the increase in standardized trading volume is statistically significant at least at the 5% level. Consistent with the findings of previous studies, the inclusion of stocks in the S&P 500 index appears to permanently increase its standardized trading volume.

Moreover, Table 5 provides evidence that the character of trading in the listed security improves in other ways. Specifically, the stock trades more frequently on average but with fewer shares per trade after the addition. The mean (median) daily trade frequency is 50% (31%) greater, and 89% of the sample firms experience an increase in trade frequency. In contrast to the increase in the event widow (noted in the last section), the daily average (median) trade size (measured in number of shares) Post/Pre ratio is 0.93 (0.90), and only 32% of the sample exhibits increased trade size. Both the changes in trading frequency and trade size are highly significant. Furthermore, the proportion of block trades (defined as the number of trades involving at least 10,000 shares divided by the total number of trades) registers a significant decline in the post-addition period. The observed increase in trade frequency and the decreases in trade size and the proportion of block trades are contrary to what we would expect from an increase in institutional ownership (given

Table 5

Long-term impact on trading activity, institutional ownership, and analyst coverage

The sample consists of 74 NYSE listed firms that were added to the S&P 500 index during the period 1993–1998. Standardized trading volume, daily trade frequency, average trade size, proportion of block trades, proportion of crossing trades, standard deviation of signed order flow, and standard deviation of daily returns measure different aspects of trading activity. Standardized trading volume is defined as daily trading volume in shares divided by the total NYSE volume for the same day. Trade frequency is defined as the number of NYSE trades for the day. Trade size is for NYSE reported trades only. The proportion of NYSE block trades is the number of trades greater than or equal to 10,000 divided by the total number of NYSE trades for the added stock. The proportion of crossing trades is the number of trades that occur between bid and ask prices divided by the total number of trades for the added stock. Signed order flow is defined as trade size multiplied by a trade indicator variable (i.e., +1 for a public buy and −1 for a public sell). Order flow is signed using the Lee and Ready (1991) algorithm. The measures of trading activity are computed for each stock for the pre-S&P 500 index addition period [−60, −10] and the post-addition period [+10, +60]. Institutional ownership denotes the number of adjusted shares held by institutions as provided by Value Line Investment Survey. Analyst following is the number of earnings estimates as reported in the I/B/E/S summary data file. Shares outstanding are obtained from Center for Research in Securities Prices (CRSP) data. Percent institutional ownership and analyst following are obtained are for each stock for the quarter prior to and following index addition. The standard deviation of daily returns is computed using the CRSP daily master file. The sign test and a standard $t$-test are used to test the null hypothesis that the measure is unchanged in the post-addition period as compared to the pre-addition period.

| | Standardized volume | Daily trade frequency | Trade size (shares) | Proportion of block trades | Proportion of crossing trades | Std. dev. of order flow (shares) | Std. dev. of daily returns | Inst. ownership (millions of adj. shares) | Analyst following |
|---|---|---|---|---|---|---|---|---|---|
| Mean | | | | | | | | | |
| Pre-addition | 0.1149% | 162.1 | 2,771 | 0.0660 | 0.2654 | 9,164 | 2.18% | 139.6 | 17.50 |
| Post-addition | 0.1248% | 217.5 | 2,463 | 0.0569 | 0.2707 | 8,061 | 2.13% | 155.7 | 17.00 |
| Median | | | | | | | | | |
| Pre-addition | 0.0741% | 126.0 | 2,690 | 0.0656 | 0.2664 | 8,151 | 2.13% | 99.8 | 17.64 |
| Post-addition | 0.0855% | 177.5 | 2,282 | 0.0480 | 0.2818 | 7,081 | 2.00% | 106.5 | 18.00 |
| Post/Pre ratio | | | | | | | | | |
| Mean | 1.2647 | 1.4996 | 0.9259 | 0.9265 | 1.0364 | 0.9692 | 0.9774 | 1.2084 | 1.0246 |
| Median | 1.1491 | 1.3118 | 0.9003 | 0.8490 | 1.0068 | 0.8865 | 0.9397 | 1.0617 | 1.0000 |
| Proportion increased | 0.64 | 0.89 | 0.32 | 0.31 | 0.54 | 0.35 | 0.50 | 0.73 | 0.45 |
| Sign test ($z$) | 2.32** | 6.74*** | −2.56** | −3.25*** | 0.70 | −2.56** | 0.00 | 3.95*** | 0.93 |
| $t$-Test | 2.61*** | 5.82*** | −2.51** | −1.57 | 1.53 | −0.56 | −0.52 | 3.35*** | 1.30 |

**Significant at 5% level.

***Significant at 1% level.

their buy-and-hold behavior). Also, we find little evidence of any change in the proportion of crossing trades (defined as the number of trades occurring within the bid–ask quotes divided by the total number of trades) and the standard deviation of daily returns of the listed stock.

Next we examine the variability of total order flow to further investigate the change in the character of order flow. The aggregate signed order flow is constructed by multiplying trade size by a trade indicator variable that equals $+1$ for a public buy and $-1$ for a public sell (based on the Lee and Ready (1991) algorithm). We use the standard deviation of total signed order flow as a measure of variability in order flow.[13] There is evidence to suggest the aggregate order flow is less variable in the post-addition period. The mean (median) Post/Pre ratio is 0.9692 (0.8865) and the sign test ($z = -2.56$) finds the change significant. The fall in order flow variability is consistent with the buy-and-hold behavior of index investors. To the extent fewer block orders and more frequent and predictable order flow allow market-makers to more easily manage their inventory, these changes in the character of order flow should reflect increased liquidity, due primarily to a decrease in the direct cost of trading.

It is not surprising that the number of shares of equity owned by institutional investors increases following S&P 500 index addition. Our measure of aggregate institutional ownership is the number of shares owned by institutions as provided by Value Line Investment Survey. We compare the number of shares (adjusted for splits) as reported in the post-addition quarter to the pre-addition quarter. The next to last column of Table 5, the mean (median) Post/Pre ratio for institutional ownership is 1.2084 (1.0617). The change is significant using either the $t$-test ($t = 3.35$) or the sign test ($z = 3.95$). Further, the change in institutional ownership is widespread as 73% of the additions sample experiences an increase in institutional ownership following addition. The results are consistent with our expectations as well as the findings of Pruitt and Wei (1989). They likely reflect the purchase of added stocks by mutual fund managers and other institutional index investors. What is surprising, however, is the insignificant increase in the number of analysts following the added stocks. The fact that the mean number of analysts per stock in the pre-addition period is quite high (approximately 17) suggests that analyst coverage and thus the information environment for added stocks is rich to begin with. Recall also that the added firms are, on average, large with substantial trading volume. The finding of no significant change in analyst coverage suggests that the source of any long-term improvement in liquidity is less likely to be related to changes in analyst coverage.

### 5.4. *Quoted spread and depth*

In Table 6, we notice a significant decrease in time-weighted quoted spread from the pre- to the post-addition period, measured either in absolute or relative terms.

[13] We also use the standard deviation of trade size and of the innovation in order flow from a fifth order auto-regressive specification and find qualitatively similar results.

Table 6
Long-term impact on bid-ask spread and depth
The sample consists of 74 NYSE listed firms that were added to the S&P 500 index over the period 1993–1998. The absolute quoted spread is the ask price less the bid price. The absolute effective spread is twice the absolute value of the difference between trade price and the prevailing quote midpoint. The relative spread measures denote the corresponding absolute measures divided by the quote midpoint. The depth is measured as the sum of quoted depth at bid and ask prices in shares and in dollar value. For each stock, the time-weighted daily average for each liquidity measure is calculated for both the pre-S&P 500 addition period [−60, −10] and the post-S&P 500 addition period [+10, +60]. The time-weight is equal to the length of time that a quote is valid as a fraction of the time between the first quote and close of the trading day. Then for each stock in the sample, the equally weighted average of the daily liquidity measures is computed for both periods. The Post/Pre liquidity ratio for each stock is defined as the ratio of the average of daily time-weighted averages of each liquidity measure for the post-S&P addition period to the corresponding average in the pre-S&P addition period. The sign test and a standard $t$-test are used to test the null hypothesis that the liquidity measure is unchanged over the post-addition period as compared to the pre-addition period.

| | Liquidity measures | | | | |
|---|---|---|---|---|---|
| | Quoted spread (%) | Quoted spread ($) | Effective spread (%) | Depth (00 shares) | Depth ($ 00) |
| Pre-addition mean | 0.4307 | 0.1588 | 0.3140 | 155.8 | 4,658 |
| Post-addition mean | 0.3853 | 0.1510 | 0.2795 | 155.0 | 5,163 |
| Pre-addition median | 0.4100 | 0.1507 | 0.2854 | 96.7 | 3,831 |
| Post-addition median | 0.3510 | 0.1454 | 0.2392 | 98.6 | 4,231 |
| Post/Pre ratio | | | | | |
| Mean | 0.8980 | 0.9587 | 0.8954 | 1.0849 | 1.1596 |
| Median | 0.8715 | 0.9559 | 0.8881 | 1.0163 | 1.0922 |
| Proportion increased | 0.20 | 0.31 | 0.26 | 0.51 | 0.62 |
| Number of stocks | 74 | 74 | 74 | 74 | 74 |
| Sign test ($z$) | −5.11*** | −3.25*** | −4.18*** | 0.23 | 2.09** |
| $t$-test | −5.61*** | −3.83*** | −4.65*** | 1.62* | 2.94*** |

*Significant at 10% level.
**Significant at 5% level.
***Significant at 1% level.

The median absolute quoted spread declines from 15.07 cents to 14.54 cents. The median proportional quoted spread drops from 0.41% to 0.35% of quote midpoint. Of the 74 stocks in the sample, 59 (51) experience a decrease in relative (absolute) quoted spread. The decrease in both measures of quoted spread is highly significant using the sign test or a standard $t$-test. In addition, we evaluate the change in the relative effective spread, which is defined as twice the absolute value of the difference between transaction price and the prevailing quote midpoint expressed as a percentage of the quote midpoint. Unlike the quoted spread, the effective spread accounts for trades that occur inside the best quotes and, therefore, is a more reliable measure of trading costs.[14] Its median value declines from 0.2854% to 0.2392%. The median relative effective spread Post/Pre ratio is 0.8881, indicating a significant decrease in the effective spread. Further, we examine the time-weighted quoted depth measured in both the number of shares offered and in dollars (i.e., depth in shares multiplied by the period's average trade price) to reflect the capital committed by market makers. While both the mean (1.0849) and median (1.0163) of the depth ratio in shares indicate an increase in depth, only the *t*-test is significant at the 10% level. However, when depth is measured in dollar terms the increase is significant. The post-addition median is $423,100 as compared to $383,100 for the pre-addition period. Overall, the univariate test results up to this point suggest that stocks added to the S&P 500 index experience a sustained increase in overall market liquidity.

### 5.5. Multivariate analysis of long-term changes in market liquidity

It is possible that the above univariate tests are biased by factors unrelated to the addition of a stock to the S&P 500 index. To control for these extraneous factors and improve the power of our tests, we perform a pooled time series cross-sectional multivariate analysis of quoted spread and depth. It is well established in the literature that quoted bid–ask spread increases with return volatility and decreases with stock price and trading volume, but depth varies directly with volume and inversely with price and volatility (e.g., Demsetz, 1968; Copeland and Galai, 1983). We use the following conventional log-linear specification under which the regression coefficients provide estimates of elasticities:

$$\begin{aligned}\log \mathit{Liquidity}_{jt} = {} & \beta_0 + \delta_0 DS\&P_t + \beta_1 \log \mathit{Volume}_{jt} + \delta_1 \log \mathit{Volume}_{jt} DS\&P_t \\ & + \beta_2 \log \mathit{Price}_{jt} + \beta_3 \log \mathit{StdDev}_{jt} + \varepsilon_j \\ & \text{for } j = 1, 2, \ldots, 74 \text{ and } t = 1, 2, \end{aligned} \tag{2}$$

where $t = 1$ represents the pre-S&P 500 listing period, $[-60, -10]$, and $t = 2$ represents the post-listing period, $[+10, +60]$. The dependent variable $\log \mathit{Liquidity}_{jt}$ is represented by either $\log \mathit{Spread}_{jt}$ (quoted spread), $\log \mathit{Spread}_{jt}$ (effective spread), or $\log \mathit{Depth}_{jt}$ ($ depth). The variables denote natural logarithms of time-weighted daily averages of spreads or depth, closing price, trading volume in shares, and return volatility for stock $j$ in time period $t$. The independent variable $DS\&P_t$ takes on a value of one in the post-S&P 500 addition period and zero otherwise. We are primarily interested in the change in the intercept, $\delta_0$, and the change in the slope coefficient of $\log \mathit{Volume}_{jt}(\delta_1)$. If the proportion of noise trading increases after a stock is included in the index, then we expect market makers to be less sensitive to order flow and to lower (raise) the spread (depth) on average. Jegadeesh and Subrahmanyam (1993) use a similar econometric specification in their study of the

[14] The effective spread for a single trade occurring at the quote midpoint is zero by definition. To avoid the obvious problem this creates in measuring percentage changes, we use the average effective spreads in the sub-periods in constructing the Post/Pre ratios. In our sample, the minimum average effective spreads are 3.12 cents and 3.32 cents in the pre- and post-addition periods, respectively.

impact of the introduction of S&P 500 futures on the liquidity of underlying stocks. Our analysis differs in two ways. First, our sample of stocks is subject to minimal clustering, but their observations are clustered in time leading to cross sectional dependence, auto-correlation and heteroskedasticity. Second, we do not constrain the slope coefficient for volume to be equal across time periods. We follow Greene (1993) in estimating a pooled cross-sectional time series regression with cross-sectional heteroskedasticity and first order auto-correlation.

The regression estimates of Eq. (2) are presented in Panel A of Table 7. Entries in the last row indicate the regressors explain around 90% of cross-sectional variation in market liquidity using the Buse (1973) version of $R^2$ for generalized least squares regressions. Consistent with previous studies, the coefficient estimates of $\log Volume_{jt}$, $\log Price_{jt}$, and $\log StdDev_{jt}$ have the predicted signs and are highly significant. More importantly, there are significant decreases in the intercepts of the two spread equations and a significant increase in the intercept of the depth equation after controlling for variations in price, trading volume, and standard deviation of returns. This indicates that liquidity as measured by spread (quoted and effective) and dollar depth both improve following S&P 500 index addition even after controlling for the effects of changes in volume, price, and volatility. For instance, the estimates of the $DS\&P_t$ coefficient (i.e., the change in intercept from the pre- to the post-addition period) show the (log) effective spread declines on average by 0.38%, while the (log) dollar depth increases by 0.35% in the post-listing period. Also, the coefficient estimates associated with the interaction term $\log Volume_{jt} DS\&P_t$ are significantly positive for the two spread regressions and negative for the depth regression. For instance, a one percent increase in mean trading volume (log *Volume*) is associated with a decrease at the margin of 0.0786% in the average effective spread in the pre-listing period. But this drop is only 0.0520% ($= -0.0786\% + 0.0266\%$) in the post-listing period. Similarly, the sensitivity of average quoted dollar depth to volume decreases from 0.5642% to 0.5339% for a 1% increase in mean trading volume. These results are consistent with the argument that bid–ask quotes, depths and trade prices are less sensitive to order flow after stocks are added to the S&P 500 index.

As a simple check of robustness of our results, we follow Kumar et al. (1998) and construct a summary measure of liquidity, $DepSpr_{jt}$, by dividing the average dollar depth by either the average quoted or effective relative spread for each security $j$ in time period $t$.[15] We estimate Eq. (2) with the two resulting measures of $\log DepSpr_{jt}$ as the dependent variable. As expected, for both summary liquidity measures, the positive and significant coefficient on the $DS\&P_t$ dummy variable for changes in the intercept indicates that liquidity has improved, on average, in the post-listing period for our sample of added stocks. Further, the coefficient estimates associated with $\log Volume_{jt} DS\&P_t$ indicate that market makers are less sensitive to changes in volume following the index addition. Our multivariate analysis also suggests a general improvement in market liquidity. The reduced sensitivity of the spread to

[15] This simple ratio specification ignores potential nonlinearity and simultaneity issues in the relation between spread and depth.

Table 7
A multivariate analysis of the long-term impact on stock market liquidity

*Panel A: Conventional regressions*
The sample consists of 74 NYSE listed firms that were added to the S&P 500 index over the period 1993–1998. A log-linear pooled cross-sectional time series regression specification is used to determine if the average market liquidity of stocks improves following their inclusion in the S&P 500 index after controlling for average share price, trading volume, and volatility of stock returns. Also, $\delta_1$ tests if the slope coefficient on log *Volume* has changed. The specification is as follows:

$$\log Liquidity_{jt} = \beta_0 + \delta_0 D\&SP_t + \beta_1 \log Volume_{jt} + \delta_1 \log Volume_{jt} DS\&P_t + \beta_2 \log Price_{jt} + \beta_3 \log StdDev_{jt} + \varepsilon_j \quad \text{for } j = 1, 2, \ldots, 74 \text{ and } t = 1, 2,$$

where $t = 1$ represents the pre-S&P 500 addition period and $t = 2$ represents the post-addition period. The dependent variable log $Liquidity_{jt}$ represents average relative spread or depth or a summary liquidity ratio of mean spread to depth for security $j$ in time period $t$. The spread measures, log $Spread_{jt}$, are the natural logs of average daily quoted and effective time-weighted relative bid–ask spreads. The quoted relative spread is the quoted dollar spread divided by the quote midpoint. The relative effective spread is twice the absolute value of the difference between trade price and the prevailing quote midpoint as a percentage of quote midpoint. The depth measure, log $Depth_{jt}$, is the natural log of daily average time-weighted dollar depth in time period $t$ for security $j$. The quoted depth is defined as the sum of depth (in shares) at bid and ask prices. The average depth in shares is multiplied by the average trade price to obtain depth in dollar terms. The independent variables, log $Volume_{jt}$, log $Price_{jt}$, and log $StdDev_{jt}$, represent average daily closing price, average daily trading volume in shares, and daily return volatility in time period $t$ for security $j$. The independent variable $DS\&P_t$ equals 1 in the post-S&P 500 addition time period and 0 otherwise. The Parks (1967) method as detailed in Greene (1993) is used to estimate a cross-sectionally heteroscedastic and first order autoregressive model. Estimated coefficients are reported with $t$-ratios in parentheses

| | Dependent variables | | | | |
|---|---|---|---|---|---|
| Independent variables | log *Spread* (quoted) | log *Spread* (effective) | log *Depth* ($ depth) | log *DepSpr* (quoted) | log *DepSpr* (effective) |
| Constant | −0.8727*** | −0.8154*** | 0.1033 | 1.0381** | 0.9167** |
| | (−6.09) | (−2.32) | (0.22) | (2.19) | (2.05) |
| *DS&P* | −0.3214*** | −0.3780** | 0.3523** | 0.6185*** | 0.6502*** |
| | (−3.55) | (−2.37) | (2.01) | (25.71) | (26.31) |
| log *Volume* | −0.0826*** | −0.0786*** | 0.5642*** | −0.0309* | −0.0786*** |
| | (−9.97) | (−4.49) | (23.31) | (−1.84) | (−3.98) |
| log *VolumeDS&P* | 0.0222*** | 0.0266** | −0.0303** | 0.5672*** | 0.6206*** |
| | (3.14) | (2.19) | (−2.33) | (14.56) | (15.97) |
| log *Price* | −0.8174 | −0.9596*** | −0.2444*** | −0.7441*** | −0.7112*** |
| | (−65.09) | (−34.84) | (−7.17) | (−19.39) | (−18.93) |
| log *StdDev* | 0.1460*** | 0.1274*** | −0.5172*** | 0.3935* | 1.0427*** |
| | (15.65) | (4.43) | (−13.51) | (1.79) | (4.03) |
| Buse (1973), $R^2$ | 0.98 | 0.91 | 0.97 | 0.86 | 0.89 |

Table 7 (*continued*)

*Panel B: Additional regressions*

In addition to the independent variables detailed in Panel A, the following additional explanatory variables are added to the pooled cross-sectional time series regression specification: log $TradeSize_{jt}$, log $TradeFreq_{jt}$, log $StdDevOF_{jt}$, log $Analyst_{jt}$, and log $InstOwn_{jt}$. $TradeSize_{jt}$ is the average trade size in shares. $TradeFreq_{jt}$ is the average number of trades per day. $StdDevOF_{jt}$ is the standard deviation of signed order flow divided by the average trade size. $Analyst_{jt}$ is the number of analysts providing earnings estimates as reported in the I/B/E/S summary history file. $InstOwn_{jt}$ is the number of shares owned by institutions as reported by the Value Line Investment Survey. Estimated coefficients are reported with *t*-ratios in parenthesis

| Regression no. | 1 | 2 |
|---|---|---|
| Independent variables | Dependent variable | |
| | log *Spread* (effective) | |
| Constant | −2.1909*** | −1.2025*** |
| | (−11.73) | (−3.81) |
| *DS&P* | −0.1764*** | 0.0576 |
| | (−2.99) | (0.33) |
| log *Analyst* | 0.0323 | |
| | (1.00) | |
| log *InstOwn* | −0.0367** | |
| | (−2.35) | |
| log *InstOwnDS&P* | 0.0278** | |
| | (2.35) | |
| log *TradeSize* | | 0.0533** |
| | | (2.25) |
| log *TradeSizeDS&P* | | 0.0141 |
| | | (0.78) |
| log *TradeFreq* | | −0.2349*** |
| | | (−10.77) |
| log *TradeFreqDS&P* | | −0.0210 |
| | | (1.29) |
| log *StdDevOF* | | 0.0574 |
| | | (1.20) |
| log *StdDevOFDS&P* | | −0.0311 |
| | | (−0.59) |
| log *Price* | −0.8991*** | −0.7977*** |
| | (−38.19) | (−27.27) |
| log *StdDev* | 0.0726*** | 0.2681*** |
| | (3.46) | (10.62) |
| Buse (1973) $R^2$ | 0.92 | 0.94 |

*Significant at 10% level.
**Significant at 5% level.
***Significant at 1% level.

aggregate trading volume suggests that the proportion of noise-trading increases after the index addition of stocks.

Panel B of Table 7 presents additional regression analyses to shed light on the role of institutional ownership and analyst following of the listed stocks in improving their liquidity. In our analysis of liquidity effects, the entry of index funds into the market for added stocks plays a central role, and the long-term change in the composition of aggregate trading volume is primarily attributable to the increased stock holdings of these passive institutions and to the potential increase in analyst coverage. To assess the impact of institutions and analysts, we introduce two additional variables into the analysis. Institutional ownership is denoted by $InstOwn_{jt}$ for each security $j$ in time period $t$ and it represents the adjusted shares of equity owned by institutions as reported by Value Line Investment Survey. For $InstOwn_{jt}$, time period $t=1$ represents the quarter prior to addition while $t=2$ represents the quarter following addition. This variable also proxies for the visibility effect—the increase in the proportion of individual liquidity traders due to the greater visibility of the added stock. Analyst following, $Analyst_{jt}$, denotes the number of analysts providing a quarterly earnings estimate as reported in the I/B/E/S summary history file. For $Analyst_{jt}$, time period $t=1$ corresponds to the month prior to addition while $t=2$ corresponds to the month following addition. The pair-wise correlation between the logs of trading volume and institutional ownership in our sample is high at 0.52. To minimize the problem of multicollinearity, we replace log *Volume* and log $Volume_{jt}DS\&P_t$ with the logs of $InstOwn_{jt}$ and $Analyst_{jt}$ and the associated slope-change dummy variables. The corresponding regression estimates are reported under the column titled Regression no. 1 in Panel B. The coefficient estimate of –0.0367 for log $InstOwn_{jt}$ indicates the average effective spread varies inversely with institutional equity ownership. Also, consistent with the notion that the increase in institutional ownership (i.e., index investors) improves liquidity, the positive slope change of 0.0278 (the coefficient of log $InstOwn_{jt}DS\&P_t$) signifies the effective spread is less sensitive to cross-sectional variations in institutional ownership in the post-listing period.[16]

Our measures of analyst following and institutional holdings are stale as they are observed only once a month or a quarter. As an alternative, we use changes in the composition of order flow in terms of trade size, trade frequency and variance of order flow as proxies for the post- to pre-change in analyst coverage and institutional ownership. With intraday data, we can measure these three trade-based variables much more accurately than ownership structure and analyst following. To investigate the effects of these changes in the character of order flow, we substitute log $TradeSize_{jt}$, log $TradeFreq_{jt}$ (representing average daily trade frequency), and log $StdDevOF_{jt}$ (denoting standard deviation of signed order flow) and the

[16] The interpretation of the coefficient estimates associated with the institutional ownership variable is complicated by the presence of multicollinearity. It is also possible these estimates are biased due to the exclusion of the volume variable. If we retain log *Volume* in Regression no. 1, the two ownership-related variables turn insignificant and the analyst coefficient becomes significant with a (counterintuitive) positive sign.

corresponding slope-change variables for the institutional ownership and analyst following variables. The regression estimates in the 2nd column of Panel B in Table 8 indicate the average effective spread varies directly with trade size and inversely with trade frequency, consistent with our expectation. But the coefficients on the standard deviation of order flow are insignificant in both the pre- and post-listing periods. Since we reported in Table 5 that the average level of trade size decreases but that of trade frequency increases over the [+10, +60] interval, the effective spread decreases on average. But the lack of significance of the two associated slope-change coefficients implies there is no drop in the sensitivity of effective spread in the post-listing period.

In summary, we find a significant long-term improvement (up to three calendar months) in the liquidity of stocks—as reflected by a decrease in spreads, an increase in trade frequency, but a decrease in both trade size and order flow variability—after their inclusion in the S&P 500 index. In addition, our estimates uncover a significant decrease in the sensitivity of spreads to cross-sectional variation in aggregate trading volume after index reconstruction. These findings suggest a long-term improvement in market liquidity for added stocks. It is still not clear, however, what percent of the observed decrease in the average effective spread is due to an increase in passive institutional ownership rather than the result of increased competition among informed institutions or the improvement in the character of the order flow. It is difficult to sort out this issue because we cannot observe ownership structure (i.e., the proportion of informed and uninformed investors). A better way to distinguish between these potential sources of improvement in liquidity is to examine the fixed and variable components of the effective spread. We take up the decomposition of spread in the next section.

## 6. Analysis using alternative measures of the effective spread

While the analysis based on the quoted spread is useful in understanding dealers' quote-setting behavior in response to the inclusion of a stock in the market index, Petersen and Fialkowski (1994) as well as Lee and Ready (1991) note that quoted spread is an imperfect measure of liquidity due to the fact that trades frequently occur within the quoted spread. In our sample, for example, approximately 20% of trades occur at the prevailing quote mid-point. The quoted spread also overstates the round trip cost of transacting because it fails to account for the tendency of prices to rise following a purchase and fall following a sale. Therefore, we analyzed in the last section a standard measure of the effective spread, defined as twice the absolute value of the difference between trade price and the prevailing quote midpoint. In this section, we examine alternative measures of the effective spread as a robustness check of our results. The price formation models of Glosten and Harris (1988), Foster and Viswanathan (1993) adaptation of Hasbrouck (1991), as well as Madhavan et al. (1997) allow estimation of the effective spread that does not rely on quotes but rather on the time series of transactions data. As a word of caution, however, the estimates of effective spread derived from the price formation models are also subject to criticism. Accordingly, these time series-based measures of effective spread are offered merely as alternatives to the standard measures used in Section 5. We begin with a brief discussion of the Glosten and Harris (GH), Foster-Viswanathan-Hasbrouck (FVH), and Madhavan, Richardson, and Roomans (MRR) methodologies. In the following discussions, the subscript $t$ indexes transaction time.

### 6.1. The Glosten-Harris (GH) model

GH model the change in transaction price, $p_t$, as

$$\Delta p_t = \lambda q_t + \psi[Q_t - Q_{t-1}] + \varepsilon_t, \tag{3}$$

where $q_t$ represents signed order flow (the product of trade size and the trade indicator variable, $Q_t$ ($Q_t = +1$ for a buy and $-1$ for a sell)) and $\varepsilon_t$ is the public information signal. The asymmetric information, $\lambda_{\mathrm{GH}}$, and the direct cost, $\psi_{\mathrm{GH}}$, parameters are estimated by ordinary least squares with an intercept term to account for possible misspecification. Following GH, the effective spread can be estimated from the empirical estimates of $\lambda_{\mathrm{GH}}$ and $\psi_{\mathrm{GH}}$ as $(2\psi_{\mathrm{GH}} + n\lambda_{\mathrm{GH}})$, where $n$ denotes average trade size· The GH estimate of effective spread assumes the asymmetric information cost of transacting is an increasing linear function of trade size with a negligible fixed component. It also assumes the direct cost of transacting is fixed per trade and does not vary with trade size.

### 6.2. The Foster-Viswanathan-Hasbrouck (FVH) model

The GH model does not account for serially correlated order flow, the tendency of buy (sell) orders to follow buy (sell) orders. As a check for robustness of the GH estimates, we also examine the Foster and Viswanathan (1993) adaptation of Hasbrouck (1991), which provides for auto-correlation in the order flow.[17] Retaining the previous notation, the FVH model estimates the innovation in order flow, $\tau_t$, as

$$q_t = \alpha_q + \sum_{j=1}^{5} \beta_{t-j}\Delta p_{t-j} + \sum_{j=1}^{5} \gamma_j q_{t-j} + \tau_t. \tag{4}$$

The time series of changes in transaction prices is now expressed as a function of the innovation in order flow (i.e., $\tau_t$) rather than the order flow itself as in the Glosten-Harris model. Specifically,

$$\Delta p_t = \alpha_p + \psi[Q_t - Q_{t-1}] + \lambda\tau_t + \upsilon_t. \tag{5}$$

The term, $\upsilon_t$, represents the innovation in the price process due to publicly available information. The asymmetric information cost, $\lambda_{\mathrm{FVH}}$, and the direct cost, $\psi_{\mathrm{FVH}}$, parameters are estimated by ordinary least squares with an intercept term to allow for misspecification. In order to estimate the effective spread with the FVH estimates of $\lambda$ and $\psi$ in a manner similar to the GH measure, it is necessary to know the

[17] We consider 5 lags as in Foster and Viswanathan (1993) and Brennan and Subrahmanyam (1996).

average innovation in trade size. Since this is not directly observable, we use the standard deviation of trade size as a proxy for the average innovation in order flow.[18] The corresponding estimate of the effective spread is equal to $2\psi_{\mathrm{FVH}} + n\lambda_{\mathrm{FVH}}$, where $n$ denotes the standard deviation of trade size.

### 6.3. The Madhavan-Richardson-Roomans (MRR) model

Since the last two models fail to explicitly account for the probability of a trade taking place inside the quotes, we also employ the model of MRR. In contrast to the previous models in which trade size is important, the MRR model assumes a constant order size, implying that trade direction may have more explanatory power than signed order flow.[19] The dealer incorporates the anticipated price effect of an incoming trade, $\lambda$, and the direct transaction costs, $\psi$, in setting a quote. Order flow is assumed to exhibit first order auto-correlation, $\rho$, and trades can take place within the quotes (i.e., at the quote midpoint) with probability equal to $\theta$. MRR show the price changes, $\Delta p_t$, can be written as

$$\Delta p_t = p_t - p_{t-1} = \alpha + (\psi + \lambda)Q_t - (\psi + \rho\lambda)Q_{t-1} + \mu_t, \tag{6}$$

where $Q_t$ and $Q_{t-1}$ are the trade indicator variables (i.e., $Q_t = +1$ for a public order to buy, $Q_t = -1$ for a public order to sell, and $Q_t = 0$ for a mid-quote transaction) for trades at time $t$ and $t-1$ respectively, $\alpha$ represents constant drift (if any) in prices, and $\mu_t$ is a composite error term that embodies both innovations in price changes due to new public information as well as errors due to price discreteness. The parameter vector $(\alpha, \psi, \lambda, \rho, \theta)$ can be estimated using Hansen's (1982) generalized method of moments (GMM). The five moment equations implied by the model exactly identify the parameter vector. Specifically, the moment conditions are

$$\mathrm{E}\begin{bmatrix} Q_tQ_{t-1} - Q_t^2\rho \\ |Q_t| - (1-\theta) \\ \Delta p_t - \alpha \\ (\Delta p_t - \alpha)Q_t \\ (\Delta p_t - \alpha)Q_{t-1} \end{bmatrix} = 0. \tag{7}$$

The first moment condition defines the first-order correlation in order flow, the second defines the probability of a mid-quote transaction, the third defines the expectation of the drift term as the average pricing error, and the last two are the OLS normal equations. In GMM estimation, the parameter vector is chosen so that the sample moments most closely approximate the population moments according to a specified weighting matrix. Since the system is just-identified, the choice of

[18] We also conduct the FVH effective spread analysis using average trade size as a proxy for innovation in order flow. The results of the analysis are unchanged.

[19] MRR motivate a constant order size in three ways. First, it permits closed form solutions to the parameters of interest. Second, large orders often originate in the upstairs market, not on the floor of the exchange. Third, informed traders may desire to break large orders into smaller ones.

weighting matrix is irrelevant. Hansen (1982) shows the GMM estimates of the parameter vector are consistent and asymptotically normally distributed. MRR demonstrate that a consistent estimate of the quoted and effective spread can be obtained from the estimated parameter vector. In their model, the (implied) quoted spread is given by $2(\psi_{\mathrm{MRR}}+\lambda_{\mathrm{MRR}})$ and the effective spread by $(1-\theta)(2\psi_{\mathrm{MRR}}+\lambda_{\mathrm{MRR}})$.

### 6.4. Empirical results using alternative measures of the effective spread

We retain TAQ transactions coded as "regular way" with a correction indicator of "0" or "1" indicating "no correction" and "corrected trade", respectively, to estimate the three models. The evidence in Table 8 indicates a significant decrease in all three alternative measures of the effective spread from the pre- to the post-addition period [+10, +60].[20] For instance, the GH median effective spread estimate drops from 0.22% to 0.18%, while the MRR median estimate declines from 0.16% to 0.13%. The finding of lower MRR estimates as compared to the GH and FVH estimates is not surprising because the MRR model explicitly accounts for the probability of a transaction occurring at the quote midpoint. The results for the effective spread measures using the FVH models are similar. Thus, the analysis of alternative effective spread measures derived from the time series properties of transactions prices and trades confirms our previous result that the inclusion of stocks in the S&P 500 index leads to a long-term improvement in their market liquidity.[21]

### 6.5. Long-term effects on the direct and asymmetric information costs of transacting

We now turn to an investigation of the sources of improvement in trading volume and market liquidity—information versus liquidity-motivated trading—associated with the inclusion of a stock in the market index. Since we cannot directly observe the type of trades, we infer the price effects of order flow from the time series properties of transaction prices. Liquidity-motivated trades are unrelated to the underlying value of securities and lead to transaction prices that bounce between the bid and ask prices. In contrast, orders submitted by informed traders will be correlated with future stock price changes. They have a permanent effect on all future prices, and they do not cause negative serial correlation in transaction prices.

[20] In dollar terms, the mean (median) GH effective spread decreases from 8.60 cents (9.23 cents) in the pre-addition period to 8.06 cents (8.83 cents) in the post-addition period. The drop is highly significant using the sign test. A $t$-test of the null hypothesis that the mean of the Post/Pre ratio is equal to unity is also rejected. Similar findings are observed with the MRR estimates. The mean (median) effective spread decreases from 6.62 cents (6.99 cents) in the pre-addition period to 6.16 cents (6.61 cents) in the post-addition period. The changes in effective spread are again highly significant.

[21] Three observations are lost for the MRR measures due to a negative direct cost component in the post-addition period. Otherwise, the minimum values for the average of alternative measures of absolute effective spread in either sub-period are 4.10 cents for GH, 4.40 cents for FVH, and 2.73 cents for MRR.

Table 8

Long-term impact on stock market liquidity: alternative measures of liquidity

Alternative measures of the relative effective spread are obtained by applying the models of GH (Glosten and Harris, 1988), MRR (Madhavan, Richardson, and Roomans, 1997), and FVH (Foster-Viswanathan (1993)and Hasbrouck (1991)). The sample consists of 74 NYSE listed firms that were added to the S&P 500 index during the period 1993–1998. The GH measure of relative effective spread is defined as $(2\psi + n\lambda)/P$ where $P$ is the average daily closing price, $\psi$ and $\lambda$ represent the direct and asymmetric information cost of transacting, respectively, and $n$ is the average trade size in shares. For the FVH measure, $n$ represents the innovation in order flow which we proxy with the standard deviation of order flow. For the MRR estimates of $\psi$, $\lambda$, and $\theta$, the relative effective spread is defined as $(1-\theta)(2\psi+\lambda)/P$, where $\theta$ is the probability of a mid-quote trade. For each stock, the three measures of effective relative spread are calculated for both the pre- $[-60, -10]$ and the post-S&P 500 addition periods $[+10, +60]$. The Post/Pre ratio for each stock is defined as the ratio of relative effective spread in the post-period to its counterpart in the pre-S&P addition period. The sign test and a standard $t$-test are used to test the null hypothesis that the effective spread is unchanged in the post-addition period as compared to the pre-addition period.

| | Relative effective spread measures | | |
|---|---|---|---|
| | Glosten-Harris | Madhavan-Richardson-Roomans | Foster-Viswanathan |
| Pre-addition mean | 0.2472% | 0.1966% | 0.2566% |
| Post-addition mean | 0.2193% | 0.1785% | 0.2295% |
| Pre-addition median | 0.2182% | 0.1592% | 0.2269% |
| Post-addition median | 0.1760% | 0.1294% | 0.1868% |
| Post/Pre ratio | | | |
| Mean | 0.8884 | 0.8803 | 0.8883 |
| Median | 0.8742 | 0.8552 | 0.8771 |
| Proportion increased | 0.23 | 0.24 | 0.24 |
| Number of stocks | 74 | 71[a] | 74 |
| Sign test | −4.65*** | −4.39*** | −4.42*** |
| $t$-Test | −4.73*** | −4.36*** | −4.36*** |

[a] The number of stocks for the MRR estimates is reduced due to three negative estimates of the direct cost parameters in the post-addition period.

*** Significant at 1% level.

In the GH, FVH, and the MRR price formation models, liquidity-motivated trades result in a transitory change in transaction prices, while information-motivated trades lead to a permanent price change that is contemporaneously correlated with the order flow. Temporary changes in transaction prices reflect direct costs of supplying liquidity unrelated to information (such as inventory carrying costs, order processing costs, clearing fees, compensation for dealer's time and capital invested, etc.), and permanent price changes account for asymmetric information costs of transacting. Our objective is to investigate whether the overall improvement in liquidity in the post-addition period is due to a decrease in the direct or asymmetric information cost of trading or both.

Following Brennan and Subrahmanyam (1996), we estimate the relative asymmetric information and direct costs of transacting as $\lambda n/P$ and $\psi/P$, where $P$ is the average daily closing price and $n$ is the average trade size in shares. For the constant order size MRR model, the relative asymmetric information cost is simply $\lambda/P$. In addition, we examine the absolute (i.e., $\lambda$ and $\psi$) costs of transacting to account for a potential upward trend in prices after the addition.

The results presented in Table 9 indicate a significant decrease in the direct cost of transacting. For example, the median Post/Pre ratio of the GH measure of the relative direct cost of transacting is 0.8723, and only 24% of sample firms have ratios greater than 1.0. The results for the FVH and MRR estimates of the direct cost of transacting are very similar. Further, the estimates suggest a decline in the asymmetric information cost of transacting, although the evidence is not as strong as that for the direct cost. Only the FVH and MRR Post/Pre ratios suggest a significant decrease in the variable relative cost of transacting from the pre- to the post-addition period. The results based on the absolute measure of spread components (i.e., cents per share; not reported in Table 9) are similar suggesting the liquidity improvements are independent of trends in stock prices subsequent to index addition. Thus, the declines in both the direct and asymmetric information costs are support for the liquidity-enhancing effects of changes in ownership composition, order flow character, analyst coverage, and arbitrage activity. Moreover, the weak reduction we notice in the average adverse selection cost suggests that the increase in the proportion of passive index traders dominates the decrease in the variance of liquidity trades due to the buy-and-hold tendency of index investors.

## 7. Analysis of abnormal returns and long-term changes in liquidity

Following the theoretical work of Amihud and Mendelson (1986), previous research has speculated that the abnormal returns associated with the inclusion of a stock in the S&P 500 index may arise from increased liquidity. Typically the assertion is based on the observed permanent increase in average trading volume and the positive and significant average cumulative abnormal return (ACAR). Others have suggested that an increase in informational efficiencies due to arbitrage trading in index derivatives leads to an improvement in liquidity and the abnormal returns on the added stock (see Erwin and Miller, 1998). However, none has performed a direct test of the linkage between the improvement in liquidity for the newly listed stocks and their abnormal returns over the event period. As noted earlier in the liquidity-wealth-effect hypothesis in Section 2, if the abnormal returns are at least in part due to an improvement in liquidity, we should observe a significant negative relationship between the cumulative abnormal returns and the long-term change in liquidity costs.

Table 9
Long-term impact on stock market liquidity: spread components
The sample consists of 74 NYSE listed firms that were added to the S&P 500 index during the period 1993–1998. The direct and asymmetric information costs of transacting are computed using the MRR (Madhavan, Richardson, and Roomans, 1997), GH (Glosten-Harris, 1988) and FVH (Foster and Viswanathan (1993) and Hasbrouck (1991)) methodologies. The absolute direct cost ($\psi$) and the absolute asymmetric information cost ($\lambda$) parameters for each stock in the sample are estimated separately for the pre-S&P 500 addition period [−60, −10] and the post-addition period [+10, +60]. Dividing the absolute direct cost parameter by the average daily closing price ($P$) yields the relative direct ($\psi/P$) cost parameter. The relative asymmetric information cost parameter is ($\lambda n/P$), where $n$ denotes average trade size for the GH, standard deviation of order flow for the FVH, and unity for the MRR estimates. The Post/Pre ratios for each stock in the sample are formed by dividing the estimate of each parameter in the post-addition period by its counterpart in the pre-addition period. The sign test and a standard $t$-test are used to test the null hypothesis that the spread components are unchanged over the test period.

| | Direct cost of trading | | Asymmetric information cost of trading | |
|---|---|---|---|---|
| | Post/Pre ratio ($\psi/P$) | Post/Pre ratio ($\psi$) | Post/Pre ratio ($\lambda n/P$) | Post/Pre ratio ($\lambda n$) |
| Glosten-Harris (GH) | | | | |
| Mean | 0.8881 | 0.9387 | 1.0563 | 1.1326 |
| Median | 0.8723 | 0.9638 | 0.8979 | 0.9485 |
| Foster-Viswanathan-Hasbrouck (FVH) | | | | |
| Mean | 0.8895 | 0.9403 | 0.9661 | 1.0308 |
| Median | 0.8716 | 0.9625 | 0.8540 | 0.9355 |
| Madhavan-Richardson-Roomans (MRR) | | | | |
| Mean | 0.8846 | 0.9240 | 0.9380 | 1.0072 |
| Median | 0.8465 | 0.9334 | 0.8829 | 0.9787 |
| Number of stocks | | | | |
| GH | 74 | 74 | 74 | 74 |
| FVH | 74 | 74 | 74 | 74 |
| MRR | 71[a] | 71[a] | 74 | 74 |
| Proportion increased | | | | |
| GH | 0.24 | 0.36 | 0.42 | 0.46 |
| FVH | 0.24 | 0.39 | 0.41 | 0.45 |
| MRR | 0.27 | 0.34 | 0.35 | 0.45 |
| Sign test ($z$) | | | | |
| GH | −4.42*** | −2.32** | −0.23 | −0.23 |
| FVH | −4.42*** | −1.86* | −1.63* | −0.93 |
| MRR | −3.92*** | −2.73*** | −2.56** | −0.93 |
| $t$-Test | | | | |
| GH | −4.63*** | −4.02*** | 0.74 | 1.67* |
| FVH | −4.56*** | −4.02*** | −0.57 | 0.54 |
| MRR | −2.11** | −1.51 | −1.75* | 0.20 |

[a] The number of observations for MRR is reduced due to three negative estimates of the direct cost parameter in the post-addition period.

*Significant at 10% level.

**Significant at 5% level.

***Significant at 1% level.

### 7.1. Abnormal returns and changes in effective spread

We use cumulative abnormal returns (CAR) over the [−4, +4] event window as our measure of abnormal returns due to S&P 500 index addition.[22] The [−4, +4] interval serves to capture the announcement date for the majority of our sample firms and the effective date for the entire sample. It also permits the unraveling of excess returns in the post-addition period due to the "S&P game." The selected event window also includes each individual day for which the average abnormal return (AAR) is statistically significant (see Table 2).

To test the relation between abnormal returns and the change in liquidity costs, we divide the sample into two groups based on the long-term change, $\Delta\%EffSpd_j$, in the standard as well as GH, FVH, and MRR measures of the relative effective spread.[23] Firms with values of $\Delta\%EffSpd_j$ below the sample median are placed in one sub-group while those with values above the median are placed in the other. We perform a standard *t*-test for the difference in means and the non-parametric Mann–Whitney U-test to test for a difference in location.[24] In Panel A of Table 10, the mean and median CAR are greater for the below median $\Delta\%EffSpd_j$ sub-group for all four measures of the effective spread. For example, firms with the below-median MRR $\Delta\%EffSpd_j$ (i.e., firms with a larger decline in the effective spread) have a mean (median) CAR of 7.97% (6.40%), while those above the median (i.e., firms with a smaller decrease or an increase in effective spread) have a mean (median) CAR of only 3.09% (3.78%). For the remaining three measures of the effective spread, the differences in means are in the predicted direction but are not as strongly significant.

In Panel B we present estimates of the cross-sectional ordinary least squares (OLS) regressions of abnormal returns on the three measures of percent change in effective spread. Consistent with our *liquidity-wealth-effect hypothesis*, all the slope coefficient estimates are negative and significant. For example, a 1% drop in the MRR measure of effective spread leads to a 0.12% increase in abnormal returns. Further, the percentage change in the MRR effective spread explains 16% of cross-sectional variation in abnormal returns. Thus, the evidence in Table 10 suggests that the improvement in overall liquidity costs is a significant, but not the only source of abnormal returns due to the inclusion of a stock in the S&P 500 index. The highly significant intercept estimates coupled with the low regression $R^2$'s, however, suggest that the improvement in liquidity provide but only a partial explanation for abnormal returns accompanying rebalancing of the S&P 500 index.

The evidence in Table 10 suggests a significant statistical relation between the average cumulative abnormal returns and the change in trading costs. We now examine the economic significance of the reduction in trading costs. The mean

[22] The results of the abnormal returns analysis are unchanged when the [−3, +3] interval is used.

[23] It is important to make clear that the measures of relative effective spread used in calculating the percent changes represent an average of all trades in each of the two sub-periods. Thus, while in theory and practice, the effective spread for a single trade can be zero, the average effective spread over a great number of trades is not zero as indicated by the minimum values of effective spread noted in footnote 14 and footnote 21.

[24] The U-test is a variant of a rank-sum test, see Hogg and Craig (1995) for a thorough discussion.

Table 10

Abnormal returns and long-term improvement in stock market liquidity

The sample consists of 74 NYSE listed firms that were added to the S&P 500 index during the period 1993–1998. The cumulative abnormal returns (CARs) are measured using the market model and the standard event study methodology over the nine trading day interval, [−4, +4], centered on the effective day of index addition. In addition to the standard measure of relative effective spread (which is equal to twice the absolute value of the difference between trade price and prevailing quote midpoint as a percentage of quote midpoint), we use three additional measures of the effective spread. These are estimated using the price formation models of Glosten and Harris, GH (1988), Foster and Viswanathan, FVH (1993), and Madhavan, Richardson, and Roomans, MRR (1997). The additional effective spread measures are expressed in relative (i.e., percent of average daily closing price) terms. $\Delta\%EffSpd_j$ measures the percent change in effective spread of stock $j$ over the post-addition period [+10, +60] relative to the pre-addition period [−60, −10].

*Panel A: Difference in CARs sorted by effective relative spread*

The sample is divided into two sub-groups based on the median value of $\Delta\%EffSpd_j$. The difference in mean and median CARs [−4, +4] between the two sub-groups is tested using a standard $t$-test and the large sample approximation for the Mann–Whitney U-test

| | Above-median $\Delta\%EffSpd_j$ sub-group | | Below-median $\Delta\%EffSpd_{jd}$ sub-group | | | |
|---|---|---|---|---|---|---|
| Effective spread measure | Mean CAR (%) | Median CAR (%) | Mean CAR (%) | Median CAR (%) | $t$-Test $H_0$: means equal | Mann–Whitney U-test ($z$ value) |
| Standard | 4.86 | 5.00 | 6.41 | 6.30 | −0.90 | −0.98 |
| GH | 3.92 | 5.00 | 7.35 | 6.30 | −2.10** | −1.68* |
| FVH | 4.08 | 5.00 | 7.19 | 6.30 | −1.85* | −1.45 |
| MRR | 3.09 | 3.78 | 7.97 | 6.40 | −2.85*** | −2.11** |

*Panel B*: *Simple linear regression of CARs [−4, +4] on $\Delta\%EffSpd_j$*

The dependent variable in the OLS regressions is CARs [−4, +4] and the independent variables are the four measures of $\Delta\%EffSpd_j$. The White (1980) heteroscedastic consistent covariance matrix is used to compute standard errors. The $t$-ratios are reported in parentheses below the parameter estimates

Dependant variable: CAR [−4, +4]

| Regression no. | 1 | 2 | 3 | 4 |
|---|---|---|---|---|
| | Measures of effective relative spread | | | |
| Independent variables | Standard | GH | FVH | MRR |
| Constant | 0.0455*** | 0.0409*** | 0.0407*** | 0.0362*** |
| | (4.48) | (4.21) | (4.14) | (3.74) |
| $\Delta\%EffSpd_j$ | −0.0817* | −0.1070*** | −0.1093*** | −0.1191*** |
| | (−1.93) | (−3.17) | (−3.13) | (−3.92) |
| Adjusted $R^2$ | 0.04 | 0.10 | 0.10 | 0.16 |
| No. of obs. | 74 | 74 | 74 | 71 |

*Significant at 10% level.

**Significant at 5% level.

***Significant at 1% level.

decrease in effective spread related in Table 6 is approximately 0.035% of share price. If we assume a normalized share price of \$1, an 8% real discount rate, and further assume that the share turns over each year in perpetuity, the present value of these cost savings is \$0.00035/0.08 = \$0.004375 or approximately 0.44% of share price.[25] Given the [−4, +4] interval abnormal return of 5.63%, the observed decrease in trading costs can explain only a fraction of the observed abnormal return from an economic standpoint.[26] Clearly, the liquidity improvement we document is only a small part of the wealth effect.

### 7.2. Abnormal returns and changes in direct and asymmetric information costs

In the last section we established the decrease in effective spread results primarily from a drop in the direct cost of transacting, which proxies for an increase in the volume of liquidity-motivated trading as a stock enters the market index. Now we use the same framework as in the last sub-section to examine whether the event window abnormal returns are attributable to the permanent (long-term) improvements in the direct and/or asymmetric information components of effective spread. In Table 11, the mean and median CAR for the sub-group with below-median values of $\Delta\%Direct_j$ (i.e., a larger decrease in the direct cost of transacting) are greater than those for the sub-group with above-median values (i.e., firms with a smaller decrease or an increase in the direct cost of transacting). For example, the mean (median) CAR for the below-median sub-group for the MRR relative direct cost measure is 9.13% (7.23%), while the mean (median) CAR for the above-median sub-group is 2.07% (2.86%). The results are consistent across our three measures of change in the direct cost of transacting, confirming a statistically significant negative relationship between the S&P addition abnormal returns and long-term changes in the direct cost of transacting.

The test results with respect to the asymmetric information cost of transacting are, however, somewhat surprising. In Table 9 we saw weak evidence of a decline in the asymmetric information cost of transacting in the post-addition period. If the abnormal returns stem also from a drop in the asymmetric information cost of trading, then we should observe higher abnormal returns for firms with below-median values of $\Delta\%Info_j$. While the $t$-tests and U-tests based on the GH and FVH measures of $\Delta\%Info_j$ suggest that the drop in the asymmetric information cost has little effect on abnormal returns, the tests based on the MRR measure of $\Delta\%Info_j$ provides confounding results. The stocks with a smaller decrease (or an increase) in the asymmetric information cost of trading in the post-addition period have significantly higher (at the 10% level) mean and median CARs (7.19% and 6.29%, respectively) than those stocks with below-median values of $\Delta\%Info_j$. A potential explanation lies in the strong negative correlation ($\rho = -0.51$) between the $\Delta\%Direct_j$ and $\Delta\%Info_j$ for the MRR measures in our sample.

[25] The basis for this discussion is Amihud and Mendelson (1988).

[26] We are grateful to the referee for suggesting this discussion.

Table 11

Abnormal returns and long-term improvements in spread components

The sample consists of 74 NYSE listed firms that were added to the S&P 500 index during the period 1993–1998. The cumulative abnormal returns (CARs) are measured using the market model and the standard event study methodology over the nine trading day interval, [−4, +4], centered on the effective day of addition. The measures of direct and asymmetric information costs of trading are from the models of Glosten and Harris, GH (1988), Madhavan, Richardson, and Roomans, MRR (1997) and Foster and Viswanathan (1993) adaptation of Hasbrouck (1991), FVH. The direct and asymmetric information costs are measured in relative (i.e., percent of average daily closing price) terms. $\%\Delta Direct_j$ and $\%\Delta Info_j$ measure the percent change in direct and asymmetric information costs of transacting in stock $j$, respectively, over the post-addition period [+10, +60] relative to the pre-addition period [−60, −10]. The sample is then divided into sub-groups based on the median values of $\%\Delta Direct_j$ and $\%\Delta Info_j$. The difference in mean of CARs [−4, +4] for the two sub-groups is tested using a standard $t$-test and the large sample approximation to the Mann–Whitney U-test.

| | Above-median sub-group | | Below-median sub-group | | | |
|---|---|---|---|---|---|---|
| %Δ in –transaction cost measure | Mean CAR (%) | Median CAR (%) | Mean CAR (%) | Median CAR (%) | $t$-Test $H_0$: means equal | Mann–Whitney U-test |
| Δ in direct cost | | | | | | |
| GH | 3.92 | 5.00 | 7.35 | 6.30 | −2.04** | −1.68* |
| FVH | 3.92 | 5.00 | 7.35 | 6.30 | −2.04** | −1.68* |
| MRR | 2.07 | 2.86 | 9.13 | 7.23 | −4.43*** | −3.39*** |
| Δ in asymmetric information cost | | | | | | |
| GH | 5.57 | 5.53 | 5.70 | 5.87 | −0.07 | −0.17 |
| FVH | 5.93 | 5.53 | 5.34 | 5.87 | +0.34 | +0.26 |
| MRR | 7.19 | 6.29 | 4.09 | 4.53 | +1.84* | +1.89* |

*Significant at 10% level.

**Significant at 5% level.

***Significant at 1% level.

Table 12 reports OLS regression estimates of abnormal returns on percentage changes in the two spread components derived from the three models of price formation. Three results are noteworthy. First, the first and second regressions for each of the models indicate that the change in the direct cost is the most significant and has the greatest explanatory power ranging from 10% under the GH and FVH measures to 15% under the MRR measure. Second, the change in the asymmetric information cost of trading provides little explanation for the cross-sectional variations in abnormal returns except for the MRR measure. Contrary to our expectation, the change in the MRR asymmetric information cost of trading is positively correlated (significant at the 10% level) with abnormal returns in the presence of $\Delta\%Direct_j$. We suspect this result is an artifact of the strong negative correlation in our sample between the two MRR measures of the spread components. Other explanations for this counter-intuitive finding are the linear

Table 12

Abnormal returns and long-term improvements in spread components: regression results

The sample consists of 74 NYSE listed firms that were added to the S&P 500 index during the period 1993–1998. The cumulative abnormal returns (CARs) are measured using the market model and the standard event study methodology over the nine trading day interval, $[-4, +4]$, centered on the effective day of addition. The measures of direct and asymmetric information costs of trading are from Glosten and Harris, GH (1988), Madhavan, Richardson, and Roomans, MRR (1997) and the Foster and Viswanathan (1993) adaptation of Hasbrouck (1991), FVH, price formation models. The direct and asymmetric information costs are measured in relative (i.e., percent of average daily closing price) terms. $\Delta\%Direct_j$ and $\Delta\%Infot_j$ measure the percent change in direct and asymmetric information costs of transacting in stock $j$, respectively, over the post-addition period $[+10, +60]$ relative to the pre-addition period $[-60, -10]$. The White (1980) heteroscedastic consistent covariance matrix is used to compute standard errors. The $t$-ratios are reported in parentheses below the parameter estimates.

| Dependant variable: CAR [−4, +4] | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| Regression No. | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 | 9 |
| Cost measures | Glosten-Harris | | | Foster-Viswanathan-Hasbrouck | | | Madhavan-Richardson-Roomans | | |
| Independent variables | | | | | | | | | |
| Constant | 0.0410*** | 0.0572*** | 0.0419** | 0.0414*** | 0.0511*** | 0.0346** | 0.0410*** | 0.0647*** | 0.0495*** |
| | (4.24) | (6.86) | (4.36) | (4.28) | (3.16) | (2.21) | (4.58) | (8.22) | (5.39) |
| $\Delta\%Direct$ | −0.1057*** | | −0.1062*** | −0.1041*** | | −0.1047*** | −0.0580*** | | −0.0473*** |
| | (−3.23) | | (−3.25) | (−3.17) | | (−3.19) | (−3.80) | | (−2.96) |
| $\Delta\%Info$ | | 0.0063 | 0.0073 | | 0.0045 | 0.0056 | | 0.0678*** | 0.0442* |
| | | (0.54) | (0.72) | | (0.48) | (0.62) | | (3.65) | (1.91) |
| Adjusted $R^2$ | 0.10 | −0.01 | 0.09 | 0.10 | −0.01 | 0.09 | 0.15 | 0.11 | 0.19 |
| No. of obs. | 74 | 74 | 74 | 74 | 74 | 74 | 71 | 74 | 71 |

*Significant at 10% level.
**Significant at 5% level.
***Significant at 1% level.

specification we impose on the relation between abnormal returns and changes in the two spread components and omission of relevant regressors correlated with the included liquidity variables (as reflected by the highly significant intercept estimates). Thus, except for this (weak) confounding result, the evidence strongly supports the *liquidity-wealth-effect hypothesis*, reflecting that a significant source of abnormal returns associated with the inclusion of a stock in the S&P 500 index is the decline in the direct cost of transacting. However, we find little support for the notion that the weak decrease in information asymmetry (noted in Table 9) is responsible for abnormal returns associated with additions to the S&P 500 index.

## 8. S&P 500 index deletions

If we find that index additions improve the liquidity of underlying stocks, then we should also observe the reverse (i.e., deterioration in liquidity) with index deletions. While this is a natural follow up question to ask, deletions present unique problems that are not easily remedied. First, there is typically a confounding event (e.g., merger, spin-off, and bankruptcy) associated with S&P 500 index deletions. Therefore, after applying the conventional criteria for sample selection, we are left with far fewer deletions than additions. Additionally, the criteria for inclusion in the S&P 500 index are well known public information that is published by Standard and Poor's.[27] So, it is plausible that the deletion of a firm from the S&P 500 index is anticipated by the market and its effects impounded in quotes, prices and volume well before the announcement of its exclusion. Notably, a firm may be deleted from the index because of low market value and/or inadequate liquidity. Therefore, it is extremely difficult to disentangle the effects on market value and liquidity due to firm-specific factors from those associated with the removal of a stock from the S&P 500 index. For these reasons, we contend that an addition offers a far less noisy event to study liquidity effects of index revisions than does a deletion. We do, however, investigate the wealth and liquidity effects on the corresponding sample of deletions primarily to highlight the contrast between additions and deletions. In this section, we present only a brief summary of our analysis with respect to deletions from the S&P 500 index.

The event study results for our sample of 27 deletions are presented in Panel A of Table 13. These results are consistent with those of Lynch and Mendenhall (1997) and generally mirror the results for index additions. Specifically, on event day –4 (the typical announcement day) the average abnormal return (AAR) is −2.40%. The AARs are negative and statistically significant for several days during the event window. On the effective day of deletion (i.e., event day 0) the mean AAR is an additional −2.40%. The average cumulative abnormal return (ACAR) for the event period [−4, 0] is −10.29% and it is widespread as ACAR[−4, 0] is negative for all stocks in the deletions sample. Following the effective day of index

[27] Beneish and Whaley (1996) provide a thorough discussion of the criteria for S&P 500 inclusion. Standard and Poor's Corporation also publishes their inclusion criteria on its website.

Table 13
Event window abnormal returns and trading volume for S&P 500 deletions

*Panel A: Abnormal returns around S&P 500 addition*

The sample consists of 27 NYSE listed firms that were deleted from the S&P 500 index during the period 1993–1998. Average abnormal return (AAR) and average cumulative abnormal return (ACAR) are calculated using the market model and the standard event study methodology. The estimation window for calculating the market model parameters is the event time interval [−250, −30]. AAR and ACAR are both tested for significance using a *t*-test and a standard sign test

| Event day/interval | AAR or ACAR (%) | Proportion with AAR or ACAR > 0 | *t*-Test $H_0$: AAR = 0 or ACAR = 0 | Sign test for AAR or ACAR |
|---|---|---|---|---|
| −5 | −0.90 | 0.30 | −0.73 | −0.19 |
| −4 | −2.40 | 0.26 | −1.96** | −2.12** |
| −3 | −1.70 | 0.22 | −1.38 | −2.50** |
| −2 | −1.26 | 0.30 | −1.03 | −2.89*** |
| −1 | −2.53 | 0.22 | −2.05** | −2.12** |
| 0 | −2.40 | 0.37 | −1.95* | −2.89*** |
| +1 | 1.85 | 0.74 | 1.50 | −1.35 |
| +2 | 0.19 | 0.52 | 0.16 | 2.50** |
| +3 | 0.50 | 0.59 | 0.41 | 0.19 |
| +4 | 0.43 | 0.52 | 0.35 | 0.96 |
| +5 | −0.22 | 0.59 | −0.18 | 0.19 |
| [−1, +1] | −3.08 | 0.33 | −0.84 | −1.73* |
| [−2, +2] | −4.08 | 0.30 | −0.68 | −2.12** |
| [−3, +3] | −5.35 | 0.33 | −0.62 | −1.73* |
| [−4, +4] | −7.32 | 0.19 | −0.66 | −3.27*** |
| [−4, 0] | −10.29 | 0.00 | −1.68* | −5.20*** |
| [+1, +4] | 2.97 | 0.63 | 0.61 | 1.35 |

Table 13 (*continued*)

*Panel B: Abnormal returns around S&P 500 deletion*

A pooled time series regression is estimated with the White (1980) covariance matrix to investigate the volume patterns around S&P 500 deletion as follows:

$$\log Volume_{jt} = \mathbf{i}\alpha_j + t\gamma + \sum_{i=-5}^{+5} D_i\beta_i + \varepsilon_j \quad \text{for } j = 1, 27 \text{ and } t = -250, +5,$$

where log $Volume_{jt}$ is the log of trading volume in shares for the stock $j$ and the day $t$, $\mathbf{i}$ is a $27 \times 27$ identity matrix, $D_i$ are dummy variables for each trading day in the event time interval $[-5, +5]$, $\varepsilon_j$ is a zero mean heteroscedastic disturbance term, and $\alpha_j$, $\gamma$, and $\beta_i$ are parameters to be estimated. The estimates for the linear drift ($\gamma$) and event period dummy variables ($\beta_i$) for the interval $[-5, +5]$ are reported with $t$-statistics in parenthesis below

Parameter estimates

| $\gamma$ | $\beta_{-5}$ | $\beta_{-4}$ | $\beta_{-3}$ | $\beta_{-2}$ | $\beta_{-1}$ | $\beta_0$ | $\beta_{+1}$ | $\beta_{+2}$ | $\beta_{+3}$ | $\beta_{+4}$ | $\beta_{+5}$ |
|---|---|---|---|---|---|---|---|---|---|---|---|
| 0.0023 | 0.2004 | 0.2128 | 0.4949 | 0.8917 | 0.9871 | 2.8898 | 1.3238 | 0.7724 | 0.5630 | 0.5328 | 0.4928 |
| (20.63) | (1.40) | (1.29) | (2.96) | (8.20) | (7.81) | (29.40) | (11.68) | (7.00) | (4.92) | (4.18) | (5.12) |

Adjusted $R^2 = 0.55$

*Significant at 10% level.
**Significant at 5% level.
***Significant at 1% level.

deletion, there is some weak evidence of reversal. The results indicate that stocks deleted from the S&P 500 index, on average, experience significant negative abnormal returns.

We analyze abnormal trading volume for deletions using the pooled time-series cross-sectional regression methodology discussed in Section 4.2 for additions. The results from estimating Eq. (1) for our sample of 27 deletions are presented in Panel B of Table 13.[28] The previously noted widespread and significant increase in trading volume in the event period for additions is also found for deletions. For example, the first trading day with abnormal volume is event day –3. The slope on the dummy ($\beta_{-3}$) is 0.4949 and is highly significant ($t = 2.96$), indicating trading volume is approximately 1.6 times its normal level after controlling for the linear drift in volume over time. The abnormal volume continues to rise and reaches its maximum on the effective day of deletion (i.e, event day 0). On this day, the estimate of the slope dummy ($\beta_0$) is 2.8898 and is highly significant ($t = 29.40$), indicating trading volume is approximately 18 times greater than normal. Following deletion, the abnormal volume subsides from its peak but remains significant throughout the event period.

In Table 14 we examine the liquidity effects over the longer horizon, that is, the post-deletion period [+10, +60] relative to the pre-deletion period [−60, −10]. The median trade frequency drops while the median trade size and standard deviation of order flow register significant increases as indicated by the sign test. Again, these results, as expected, are opposite to those for the sample of additions presented in Table 5. In contrast to the decline for the added stocks (see Table 6), the mean relative effective spread for the excluded stocks increases from 0.6802% to 0.7265% ($t = 1.35$). The increase in relative effective spread occurs for 63% of the deleted stocks (sign $z = 1.70$). As a robustness check, we also compute the relative effective spread using the Madhavan et al. (1997) methodology.[29] The mean (median) Post/Pre ratio for effective spread is 1.1726 (1.1066) and the change is found to be significant at the 10% level ($t = 1.94$). Further, the spread decomposition using the MRR methodology suggests the direct cost of trading increases while the asymmetric information cost remains unchanged. A confounding liquidity result for deletions that we observe is a significant increase in quoted dollar depth. This perhaps reflects attempts by market makers to accommodate increased trade size and trading volume following deletion. Overall, our findings support the predicted decline in market liquidity of stocks deleted from the S&P 500 index. We do, however, caution that deletions are perhaps a very noisy event to study the impact of index revisions on liquidity as firms are often deleted for reasons based on inadequate liquidity and/or low market value.

[28] We also estimate Eq. (1) separately for each stock in the deletions sample thus allowing for different drift terms and slope coefficients across stocks. The results of these regressions are consistent with the pooled approach as all 27 stocks in the sample exhibit statistically significant abnormal volume on the day of deletion.

[29] We also use the GH and FVH methods and find essentially the same results; these results are not reported to conserve space.

Table 14
Index deletions and long-term impact on trading activity and stock market liquidity
The sample consists of 27 NYSE listed firms that were deleted from the S&P 500 index during the period 1993–1998. Daily trade frequency, average trade size, standard deviation of signed order flow, as well as quoted and effective spread are analyzed. Trade frequency is defined as the number of NYSE trades in the deleted stock for the day. Trade size is for NYSE reported trades only. The signed order flow is the trade size multiplied by a trade indicator variable (i.e., +1 for a public buy and −1 for a public sell). Order flow is signed using the Lee and Ready (1991) algorithm. Relative quoted spread is the ask price less the bid price as a percentage of the quote midpoint. Relative effective spread is twice the absolute value of the difference between trade price and quote midpoint expressed as a percentage of quote midpoint. The Madhavan et al. (1997) price formation model is also used to estimate relative effective spread and its direct and asymmetric information components. The measures of trading activity and costs are computed for each stock for the pre-S&P 500 index deletion period [−60, −10] and the post-deletion period [+10, +60]. The Post/Pre trading activity and cost ratios for each stock are computed by dividing the average trading activity or cost measure over the post-deletion period by its counterpart in the pre-deletion period. The sample percentage of firms with a Post/Pre ratio greater than one are reported in the row Proportion >1.0. The sign test and a standard $t$-test are used to test the null hypothesis that the trading activity or cost measure is unchanged over the post-deletion period as compared to the pre-deletion period.

| | Daily trade frequency | Trade size | Standard deviation of order flow | Relative quoted spread | Relative effective spread | Relative effective spread (MRR) | Relative direct cost ($\phi/P$ MRR) | Relative information cost ($\lambda/P$ MRR) | Quoted depth ($ 00) |
|---|---|---|---|---|---|---|---|---|---|
| Pre-deletion mean | 69.3 | 181.45 | 629.49 | 1.5265% | 0.6082% | 0.8506% | 0.4102% | 0.1597% | 2,486 |
| Post-deletion mean | 57.3 | 281.93 | 777.63 | 1.6881% | 0.7265% | 1.0716% | 0.5115% | 0.1542% | 3,659 |
| Pre-deletion median | 63.6 | 183.20 | 532.20 | 1.0807% | 0.3752% | 0.4808% | 0.2148% | 0.1281% | 1,761 |
| Post-deletion median | 50.3 | 237.20 | 635.10 | 0.9731% | 0.3678% | 0.4727% | 0.2177% | 0.1411% | 2,288 |
| Post/Pre ratio | | | | | | | | | |
| Mean | 0.9198 | 1.7079 | 1.6337 | 1.0993 | 1.1592 | 1.1726 | 1.1470 | 1.0684 | 1.4644 |
| Median | 0.8372 | 1.5980 | 1.4024 | 0.9772 | 1.0638 | 1.1066 | 1.0480 | 0.9566 | 1.2775 |
| Proportion >1.0 | 0.30 | 0.93 | 0.63 | 0.48 | 0.63 | 0.58 | 0.65 | 0.48 | 0.67 |
| Sign test | −2.12** | 4.94*** | 3.00** | −0.19 | 1.70* | 1.94* | 1.74* | 0.91 | 3.65*** |
| $t$-Test | −1.25 | 4.93*** | 1.35 | 1.19 | 1.35 | 0.78 | 1.57 | −0.19 | 1.73* |

*Significant at 10% level.
**Significant at 5% level.
***Significant at 1% level.

## 9. Conclusions

Although many studies have found abnormal returns and a permanent increase in trading volume for stocks added to the S&P 500 index, the existing empirical literature fails to provide an in-depth analysis of the role of liquidity in generating excess returns for the newly listed stocks. The theoretical work in market microstructure suggests the liquidity impact of index revisions is complex and difficult to predict. Our empirical analysis reveals a significant and long-term improvement in market liquidity for S&P 500 stocks following index addition. Specifically, we find a decrease in the time-weighted quoted spread and an increase in the time-weighted quoted depth, not only over the first ten trading days surrounding the effective day of addition but also up to 60 trading days in the post-addition period. We also find a significant increase (decrease) in trading frequency (trade size and variability of order flow). Further, a our analysis reveals not only a significant decrease in the average effective spread but also a decrease in the sensitivity of spread to aggregate trading volume and trade frequency in the post-listing period, after controlling for the effects of stock price and volatility of stock returns.

A decomposition of the effective spread using three alternative market microstructure models of transactions costs shows significant declines in both the direct and to a lesser extent the asymmetric information costs of trading. We observe little change in analyst coverage for additions despite a significant increase in institutional equity ownership. Finally, consistent with Amihud and Mendelson (1986), we find a significant relationship between the abnormal returns associated with the entry of a stock into the index and the observed decrease in effective spread as well as the decrease in the direct cost of trading. In contrast to index additions, stocks deleted from the index incur negative abnormal returns and their market liquidity suffers over three months after deletion. Specifically, the relative effective spread rises on average in the post-deletion period, primarily attributable to an increase in the direct cost of trading.

Despite the evidence of a significant relation between market liquidity and the announcement widow abnormal returns, the wealth effect remains largely unexplained. If S&P's announcement of an index revision is void of valuation relevant information, the wealth effect associated with index revisions is inconsistent with common notions of efficient markets. In fact, the evidence presented, specifically abnormal returns that persist for at least two full weeks of trading following addition, supports the imperfect substitutes hypothesis of Scholes (1972) and Shleifer (1986) (i.e., stocks have a downward sloping (vs. horizontal) demand curve). Extensions that focus on non-liquidity-based explanations for the widely documented wealth effects are promising avenues for further research.

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