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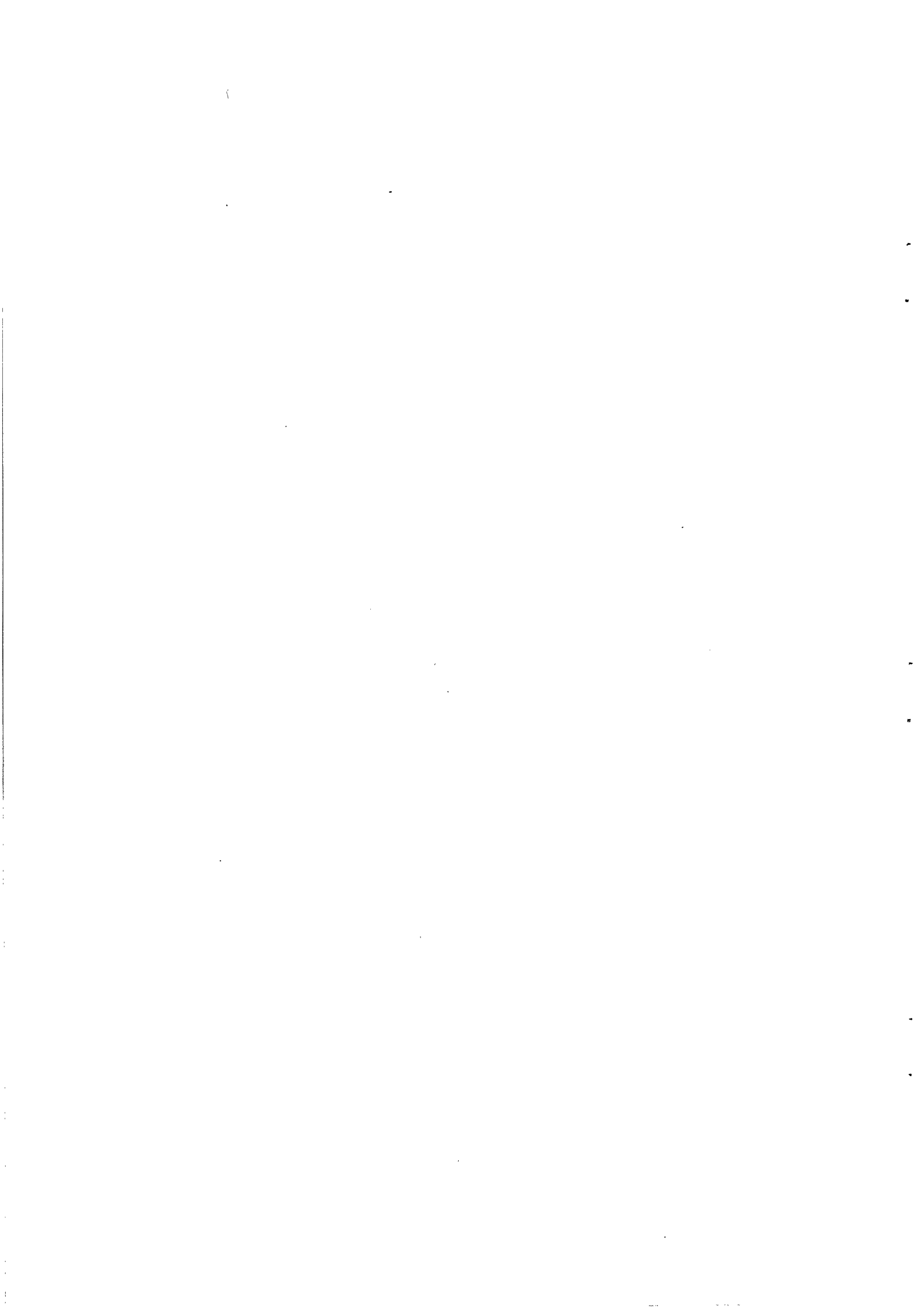
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**TRANSACTION COSTS AND PRICE VOLATILITY:  
EVIDENCE FROM COMMISSION DEREGULATION**

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**Transaction Costs and Price Volatility:  
Evidence from Commission Deregulation**

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## Transaction Costs and Price Volatility: Evidence from Commission Deregulation

**ABSTRACT:** Recent studies have suggested that lower transaction costs induce small (or noise) traders to trade more actively, thus increasing both the noise component and total volatility of asset prices. We empirically evaluate this conjecture by examining changes in volatility surrounding the abolition of fixed brokerage fees on the NYSE. We find that the lowering of transaction costs is associated with a significant *reduction* in volatility for the aggregate NYSE portfolio and five size-ranked portfolios, even after accommodating market-wide conditional heteroskedasticity. Ancillary tests indicate that prices are better characterized as random walks following the fixed fee abolition.

## **Transaction Costs and Price Volatility: Evidence from Commission Deregulation**

### **I. Introduction:**

As noted by Schwert and Seguin (1993), a new annual rite has been introduced into the Washington legislative calendar: the proposal of some sort of securities transaction tax (STT). Broadly-based STTs have been under consideration since 1987, with more recent proposals more narrowly focused, including “user fees” for exchange traded derivatives. Proponents of STTs have suggested benefits beyond generating additional tax revenues. For example, Stiglitz (1989) argues that an STT taxes investors with short-term trading horizons more frequently than those with longer horizons. This disproportionate charge reduces the trading and influence of short-term investors on management, leaving managers free to engage in more “proper” long-run strategies like research and development. Second, proponents of an STT argue that the tax would reduce a negative societal externality due to “excessive” speculation by “throwing sand in the gears” of financial markets (Summers and Summers (1989, especially p.231) and others).

Finally, and of greatest pertinence to this study, proponents have also suggested that the imposition of a transactions tax may be an effective weapon in the war on “excess” volatility. The reasoning underpinning this claim stems from the belief that there are “noise traders:” traders who submit orders for reasons other than fundamental information.<sup>1</sup> Because noise trades are not based on information about underlying values, these trades may move prices away from their intrinsic value, reducing price informativeness while increasing volatility. Proponents of an STT suggest that the tax would diminish the activities of noise traders. As a result, their impact on the quality and stability of prices would decline.

However, as pointed out by Grundfest and Shovin (1991), despite a voluminous literature of rhetoric on the benefits or costs of such taxes, there exists little empirical research measuring effects. Arguably, the best empirical evidence to date is provided by Umlauf (1993), who “create(s) a single data point for the transaction tax debate” by evaluating Swedish transaction taxes during the 1980s. Umlauf finds that the introduction of or increase in the Swedish tax led to significant increases in volatility; daily variances were highest during the period of greatest tax, for example. Further, Swedish transaction

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<sup>1</sup>We define noise traders broadly here to include portfolio insurers and other so-called “positive feedback” traders (DeLong, Shleifer, Summers and Waldman (1990)), since their trades are based only on changes in reported prices, rather than on intrinsic value, as well as “dentists and doctors in the midwest and the retired individuals in the sunbelt...” (Stiglitz, (1989)) who “believe (irrationally) that trading systems, horoscopes, etc., are beneficial in forecasting prices.” (Roll (1989)).

taxes affected the amount of price noise. Weekly-to-daily variance ratios varied inversely with the tax, with ratios significantly below unity during the high tax regime.

This study complements Umlauf (1993) by providing an additional clinical data point to the debate. In contrast to Umlauf, we examine the effects on volatility and noise of an across the board *reduction* in transactions costs. Specifically, we investigate the introduction of lower, negotiated commissions on May 1, 1975 (May Day). Regulated commissions are similar to transaction taxes since both are fixed in amount and levied on parties whenever a securities transaction occurs. Thus, our event is analogous to a one-time reduction in a tax on equity transactions.

An obvious deficiency in our experimental design is that we look at an analogous event rather than a change in a governmental securities transactions tax *per se*. Nonetheless, our experimental design offers at least two advantages. First, we use American data. This feature enhances the pertinence of our results to the ongoing debate about the role of transaction taxes in U.S. markets, since our experiment takes place in the same microstructure, regulatory and competitive environments. Second, the change from fixed to negotiated commissions occurs for only a subset of securities: those listed on the New York Stock Exchange (NYSE). Since fixed commissions on all Over-the-Counter (OTC) securities were explicitly forbidden under the Maloney Act of 1938, their commission structure was not affected by the legislative change. Therefore, we can use a portfolio of OTC issues as a control sample. Using this control, we can examine changes in volatility associated with the decline in transaction costs relative to changes in volatility of an unaffected portfolio. In this way, we are able to increase the statistical power of our tests by controlling for changes in aggregate volatility.

Our empirical results uniformly reject the hypothesis that the abolition of fixed commissions increases volatility in general, and the noise component of returns in particular. Indeed, consistent with Umlauf, we find that a reduction in transaction costs is associated with a decline in all measures of volatility. We find lower volatility at both the firm and size-ranked portfolio level, even after correcting for firm size and changes in conditional heteroskedasticity of the "market factor." Further, we find that returns to both individual firms and the five size-ranked portfolios follow a random walk more closely after the introduction of negotiated commissions. Although the links between trading noise, market efficiency and random walks are still debated, we interpret this last result as being consistent with the belief that the reduction in commissions and trading costs decreased the noise component of NYSE stock returns.

Aside from their relevance to the debate on the efficacy of transaction taxes, our findings are of interest in a broader context. Studies have often argued that the deleterious

impact of noise traders on financial markets has been exacerbated by lower transaction costs. The belief that the impact of noise traders could be altered via changes in transaction costs has prompted some to advocate increasing these costs through avenues other than transaction taxes, including increased margin requirements or reductions in the availability of low cost substitutes like equity derivatives.<sup>2</sup> However, our results suggest an alternative course of action. Both Umlauf's and our results find a positive relation between volatility and transaction costs. Therefore, increasing the costs of trading through transaction taxes, increasing margin requirements or reducing the availability of low cost substitutes like equity futures may have the effect opposite of that intended.

In the following section, we review the events surrounding the abolition of fixed commissions in 1975, and review existing empirical work surrounding that event. In Section 3, we discuss data and methods. Changes in volatility surrounding the deregulation of commissions are examined in Section 4, followed by our conclusions.

## **II. A Brief Regulatory Background:**

In 1968, the NYSE petitioned the Securities and Exchange Commission for an increase in the nominal levels of fixed or mandated minimum commission rates. Surprisingly (at the time), the Department of Justice objected to an increase and instead suggested that the entire system of mandated minimum commissions be reconsidered. Following years of controversy, the fixed commission framework was eliminated with the passage of Section 6(e)(1) of the Securities Exchange Act, and negotiated commissions were introduced on May 1, 1975.

In this section, we contrast the commission structures before and after May Day, and review studies that document changes in market and brokerage firm performance subsequent to the change. We do not review the arguments made in favor or against a fixed commission structure. For details on these arguments, see reviews by Tiniç and West (1980) or Stoll (1979).

### *2.1. The Deregulation of Commissions*

The abolition of fixed commissions is perhaps best viewed as the most important and visible discrete event in a gradual process of commission reform dating back to December, 1968. At that time, discounts of mandated proportions were first allowed for transactions in excess of a half-million dollars. Over the following years, the size of the discount and

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<sup>2</sup>See Stiglitz (1989) and Summers and Summers (1989) for discussions linking noise trading with total transaction costs. Similar arguments in the context of margin requirements are summarized in Chance (1990), in the context of exchange traded options by Skinner (1989), and in the context of equity index futures by Bessembinder and Seguin (1992).

the minimum qualifying market value were both altered and the relative cost of transacting large versus small blocks continually fell. By 1975, the per share cost of transacting a 10,000 share block was roughly one-third the cost of transacting a round lot.

As anticipated, the introduction of negotiated commissions instigated a permanent decline in commissions. According to Stoll (1979), surveys conducted straddling the introduction of negotiated commissions found that commissions fell between 31% to 44% for institutions and by up to 47% for individuals. Commission rates for small transactions fell even more for frequent traders, according to Ofer and Melnik (1978). They find declines in commissions of 22%, even for orders of 100 shares or less.

Such declines understate the true declines available to traders for at least two reasons. As Stoll (1979) observes, with the advent of negotiated commissions came the arrival of discount brokers to NYSE listed securities. Jarrell (1984) reports that discount brokers, who, by rule, were excluded from discounting before May Day, captured 3% of Big Board volume by 1977 and 6% by 1980. Thus, all traders willing to unbundle research and other ancillary services from order execution services could obtain still lower commission rates. Further, Tiniç and West (1980) note that, for this period of relatively high inflation, calculated nominal reductions in commissions understate real and proportional reductions. To conclude, the abolition of fixed commission rates led to an immediate, permanent and economically significant decline in the cost of trade execution, although the magnitude of the reduction varied across order size, share price, and the identity and trading frequency of the trader.

## *2.2. Other Effects of Negotiated Commissions: A Review*

The most intently studied fallout of the reduction in commission rates has been its effect on the profitability and competitiveness of the brokerage industry. Schwert (1977), Stoll (1979) and Jarrell (1984) review measures of the profitability of the brokerage industry such as revenues, NYSE seat prices and industry concentration. Though seat prices fell significantly in anticipation of the regime shift, these studies found little evidence that profitability was adversely affected. Jarrell reports that the equity of National Full Line brokers significantly outperformed the market after controlling for changes in seat prices and changes in NYSE volume. Stoll (1979 p. 55) finds that despite declines in revenues, institutional firms remained profitable, indicating "a remarkable degree of adaptability by the institutional firms remaining in business." Some firms failed to make a successful transition, however. Jarrell concludes that medium-sized, research-oriented firms suffered the most. Though measures of concentration indicated a slight increase in concentration,



Stoll argues that this increase is due to factors other than the introduction of negotiated commissions.

Of greater relevance to the issues considered here, Jarrell (1984) examines the effect of the regime shift on equity volume. He finds increases in volume ranging from 30% to 100% depending on the estimation period and the technique used. Thus, lower commission rates (or lower transaction costs in general) incited more active and frequent trading by all agents. This is consistent with the belief that transaction costs can affect trading volume. Of course, the key issue is whether this additional volume is associated with increased volatility.

The effects of the introduction of negotiated commissions on volatility have not been studied explicitly, to our knowledge. The only available evidence is indirect and concerns block sale underpricing, average price change per transaction and average bid-ask spread<sup>3</sup>. No evidence of a deterioration of market quality could be detected using these measures.

### *2.3. The Relevance for Transaction Taxes*

The relevance of the results and conclusions of this study to the debate on the efficacy of transactions taxes depends crucially on the degree to which the repeal of mandated commissions is analogous to a reduction in a transaction tax. Perhaps the most obvious appeal of this experiment as an approximation to a decline in transaction taxes is that our event is a one-time, across-the-board significant reduction in transaction costs.

To estimate the magnitude of this reduction, we recall that past studies have found no evidence of changes in the bid-ask spread around the regime change. If we assume an average bid-ask spread of 25¢, a reduction in commissions from 28¢ per share to 19¢ per share represents a 17% reduction in total transaction costs for institutional orders between 1,000 and 10,000 shares.<sup>4</sup> Across all transaction types and sizes, percent reductions in total transaction costs range from 4% (for individual orders under 200 shares) to 19.7% (for institutional orders between 200 and 1,000 shares).

However, changes in commissions may not be perfectly analogous to changes in an STT for at least two reasons. The decline in trading costs examined here was not uniform, with institutional and active traders enjoying greater reductions. Broad-based transaction taxes, in contrast, are designed, in theory, to be fair. However, as Heaton and Lo (1983), McConnell (1983) and Schwert and Seguin (1993) point out, it is virtually impossible to derive any sort of transaction tax that is borne equally by all agents and does not induce distortions.

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<sup>3</sup>See NYSE (1975), SEC (1976) or Stoll (1979, pp. 70-71).

<sup>4</sup>These numbers are those reported by the SEC (1976) as average commission rates measured as of April 1975 and December 1975.

Second, Grundfest and Shovin (1991) and Summers and Summers (1989) argue that volatility and noise are affected by the ratio of destabilizing to stabilizing traders. In this situation, it is conceivable that the way the elimination of fixed commissions would alter this ratio differs from the way a reduction in a transaction tax would. It is, of course, impossible to refute this possibility. However, given the similarities in the incidence and nature of commissions and transaction taxes noted above, we believe the two are closely related.

### **III. Data and Methods:**

#### *3.1. Data description*

We use the Center for Research in Security Prices (CRSP) daily data for both NYSE and NASDAQ securities. We include all firms that traded on the NYSE without interruption from one year before to one year after May 1, 1975, and that have an unbroken series of daily closing prices over this interval in the NYSE sample. Data requirements for the NASDAQ sample are similar. The choice of one year for the pre- and post-event volatility analysis is somewhat arbitrary but reflects a balance between power considerations and a desire to focus on the commission reduction alone.

We next sort the 1,349 securities in the NYSE sample into five roughly equal-sized portfolios based on the market value of common equity outstanding (size, hereafter) on April 30, 1975. We then assign NASDAQ securities to portfolios using the same size breakpoints as the NYSE sample. We remove NASDAQ firms smaller than the smallest NYSE firm in the sample (approximately \$2.1 million in outstanding equity). We control for the size of the NYSE issues since we believe that size is an easily compiled proxy for many unobservable factors that affect volatility, including information quality and quantity, volume and the bid-ask spread. Finer controls are examined below. All returns are continuously compounded.

#### *3.2 Sample Description*

Table 1 reports cross-sectional means of market values, stock prices, and daily return variances for NYSE and NASDAQ portfolios, as well as average closing bid-ask spreads for the NASDAQ portfolios. Since NASDAQ firms are small relative to NYSE firms, 1,035 of 1,457 or 71% of the NASDAQ securities are assigned to the smallest portfolio, while only 2.5% of the NASDAQ firms qualify for the largest portfolio. While this facilitates the analysis for small firms, the small number of firms in the large-NASDAQ-firm portfolio yields less statistical power to detect changes in the volatility of larger stocks.

A comparison of the size-sorted portfolios indicates that our simple control is effective: within each portfolio, features of the NYSE and NASDAQ firms are quite similar, and average market values are almost identical for the middle portfolios. Since NYSE firms in the smallest portfolio are somewhat larger than their NASDAQ counterparts, we require a more thorough control which we accomplish below in a regression context.

### 3.2.1 Comparisons of Volatility

From Table 1, NYSE stocks appear to be more volatile than NASDAQ stocks of similar size, especially for the smaller stocks. These differences are illusory, however, and are entirely attributable to CRSP's reporting conventions for the two markets. Specifically, returns for NYSE stocks are calculated using closing transaction prices, while during this period, NASDAQ prices are calculated as the average of closing bid and ask prices. Because of bid-ask bounce, returns for NYSE stocks calculated from the CRSP data will overstate the true variance in daily returns. Roll (1984) shows that this spurious variance in transaction returns can be as high as  $s^2/2$ , where  $s$  is the security's proportional spread.<sup>5</sup> In contrast, returns on NASDAQ stocks will not reflect this spurious volatility.

According to the point estimates in Table 1, the difference in reporting convention creates biases large enough to explain the differences in volatility between markets. For example, using the average proportional spread of the third portfolio,  $s^2/2 = 10.58$  ( $\times 10^{-4}$ ), which exceeds the difference between the two samples.

Because only the NYSE returns contain some bid-ask bounce, we cannot use variance ratios to compare volatilities before and after May Day. To demonstrate, suppose that securities in both markets have true return volatilities of  $U$  before the event and  $V$  afterward, with  $U \neq V$ . Further suppose that NYSE security returns are measured with error due to bid-ask bounce causing spurious volatility of  $s$  throughout the time period. With these assumptions, the calculated variance ratio becomes:

$$VR_{NYSE} \equiv \frac{(U + s)}{(V + s)} \neq \frac{U}{V} \equiv VR_{NASDAQ}$$

where  $VR_j$  is the ratio calculated for market  $j$  = NYSE, NASDAQ. The addition of noise to both the numerator and denominator biases the ratio towards one and, according to Jones

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<sup>5</sup>If adverse selection exists, transactions at the bid or ask price will be informative and the spurious variance will be lower. If the bid-ask spread is due solely to adverse selection, there will be no spurious variance; transaction prices are then true prices (see Glosten and Milgrom (1985)). Empirical estimates of the fraction of the spread attributable to adverse selection vary from about 10% in George, Kaul and Nimalendran (1991) to 40% in Stoll (1989). Spurious variance will also be lower if there is no trading on some days, because on such days CRSP records the average of the closing bid and ask prices (see Jones, Kaul and Lipson (1993)).

and Kaul (1993), causes Type I error in a test of the null hypothesis that volatility changes across the two markets are identical.

However, the use of volatility differences rather than a ratio eliminates this problem. In the example above,

$$\Delta\sigma_{\text{NYSE}} \equiv (U + s) - (V + s) = U - V \equiv \Delta\sigma_{\text{NASDAQ}}.$$

Thus, if we assume that the samples are identical except for constant but unknown measurement error, then the strategy is straightforward: compare changes in volatility for NYSE firms to changes in volatility for the control NASDAQ firms. Therefore, for each security, we calculate the standard deviation of daily returns for one year prior to ( $\sigma_{\text{pre}}$ ) and one year after ( $\sigma_{\text{post}}$ ) May 1, 1975 and calculate the change in standard deviation as:  $\Delta\sigma \equiv \sigma_{\text{post}} - \sigma_{\text{pre}}$ . We acknowledge that this method is predicated on the assumption that volatility is constant within each of the two regimes. We explicitly accommodate market-wide heteroskedasticity in section 4.2.

We also calculated daily volatility using a procedure similar to one in Schwert (1990). This procedure captures short-term movements in conditional expected returns. Results of are unchanged.

#### **IV. Results**

Volatility changes for stocks in the two markets are compared in Table 2, which reports cross-sectional averages of individual security volatilities. Bootstrapped standard errors that accommodate cross-sectional dependence are reported in parentheses<sup>6</sup>.

Evidence of unequal volatility responses is strongest for the smallest firm portfolios, where NYSE-listed firm volatilities fell by 2.5 times as much as the corresponding NASDAQ-listed firm volatilities. This decline is both economically and statistically significant. For larger firms, the evidence is less clear. Although average volatilities were lower for all portfolios, the declines associated with NYSE portfolios 3 and 5 are statistically indistinguishable from their NASDAQ counterparts, with volatilities for NYSE portfolio 4 actually falling less than their NASDAQ counterparts. The relatively small size of this control portfolio (65 firms) may be to blame. This supposition is supported by the observation that  $\sigma_{\text{pre}}$  for this NASDAQ portfolio is high relative to that of the other portfolios.

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<sup>6</sup>Specifically, in each of 500 iterations, 250 days are chosen at random and with replacement from the pre-event year, and 250 days are similarly chosen in the post-event year. We use these days for every firm in the calculation of a  $\Delta\sigma_i$ . This method accommodates cross-sectional correlation, since  $\Delta\sigma_i$  is calculated using the same trading days for each security  $i$ . Next, we calculate  $\Delta\sigma$ , the cross-sectional average of  $\Delta\sigma_i$ , in each iteration. We calculate standard errors as the standard deviation of estimated  $\Delta\sigma$ 's across all iterations. The percent of all iterations where  $\Delta\sigma$  is positive is reported in the final column. See Efron and Tibshirani (1986) for a survey.

#### 4.1. Cross-sectional regression analysis

We test for changes in volatility in a cross-sectional regression framework for three main reasons. First, this framework controls for firm size more accurately. Second, since the analyses reported in Table 2 control for size by grouping NYSE and NASDAQ stocks into five market value portfolios, the information in the data is reduced. A regression approach mitigates this problem. Finally, the small number of firms in the larger NASDAQ portfolios may be responsible for a lack of power in statistical tests for larger firms. By combining all qualifying firms, we circumvent this power problem.

Using all 2,958 securities which meet the two year data requirements, we estimate the equation:

$$\Delta\sigma_i = 0.255 - 0.133 I_{NY,i} - 0.055 \ln(\text{SIZE}_i) + \varepsilon_i,$$

(0.025)                      (0.009)

where  $\Delta\sigma_i$  is the change in the standard deviation of daily returns from the pre-event to post-event period, expressed in percent per day;  $I_{NY,i}$  is an indicator variable equal to one for NYSE securities and zero for NASDAQ-listed securities; and  $\text{SIZE}_i$  is the market value of outstanding common equity (in thousands of dollars) on April 30, 1975. The sample includes the 2,806 securities in the Table 2 analysis as well as 152 additional NASDAQ securities that were smaller than the smallest NYSE stock and thus excluded from the first analysis. The results are essentially identical when these 152 securities are excluded. Due to the severe non-normality of the dependent variable (and, therefore, of the residuals) and cross-sectional correlations, the parentheses contain standard errors that are based on bootstrapping.

Estimates of the coefficients confirm and strengthen the results reported in Table 2. Specifically, we find that the abolition of fixed commissions for NYSE-listed securities is unambiguously associated with a statistically and economically significant decline in the daily volatility of NYSE stocks relative to NASDAQ stocks, even after precisely controlling for differences in firm size. Note that the negative coefficient on  $\text{SIZE}_i$  indicates that volatility declines were most pronounced for large firms.

As a robustness check, we estimated non-linear and polynomial versions of the above equation. These specifications would accommodate any non-linearities in the relations between size and volatility changes. The results reported above are the most conservative. As a second robustness check, we decompose  $\text{SIZE}_i$  into its components:  $\text{SHARES}_i$ , the number of shares outstanding for security  $i$  (expressed in thousands), and  $\text{PRICE}_i$ , the price per share for security  $i$ , both as of April 30, 1975. Estimating this specification yields:

$$\Delta\sigma_i = 0.523 - 0.104 I_{NY,i} - 0.106 \ln(\text{SHARES}_i) + 0.007 \ln(\text{PRICE}_i) + \varepsilon_i$$

(0.023)                      (0.007)                      (0.017)

Again, volatility declines are significantly greater for the NYSE-listed firms after controlling for firm size and its components. The two components of firm size differ in their significance: share price is irrelevant, while the number of shares outstanding is significantly related to the magnitude of the volatility decline. As above, we estimated non-linear and polynomial versions of this specification. Again, the results reported in the text are the most conservative.

#### 4.2 Time series regression analysis

We next examine the volatility of NYSE size-ranked portfolios controlling for changes in conditional heteroskedasticity unrelated to the regulatory change. Rather than assuming that the volatility in the two markets is constant, as we did above, we instead allow the volatility of a portfolio to vary over time in response to a single factor.

Our methods closely follow those of Schwert and Seguin (1990), who model conditional heteroskedasticity for a portfolio as being linear in the predicted standard deviation of some aggregate portfolio or factor.<sup>7</sup> We implement their model by forming five equally-weighted NYSE portfolios using the same size breakpoints that we used in the earlier analysis. As above, we exploit the fact that NASDAQ commissions were beyond the scope of the new regulations, and so use the equally-weighted NASDAQ portfolio as our proxy for the single volatility factor.

Finally, to capture changes in the level of volatility associated with the regime change, we add an indicator variable to the specification. As a result, we can examine changes in volatilities of size ranked NYSE portfolios associated with the deregulating of commissions while controlling for conditional heteroskedasticity using the aggregate NASDAQ portfolio.

We estimate the following regression from one year before to one year after May 1, 1975:

$$\Delta\sigma_{pt}^{NY} = \gamma_0 + \gamma_1 \Delta\sigma_t^{OTC} + \gamma_2 I_{POST,t} + \varepsilon_t,$$

where  $\Delta\sigma_{pt}^{NY}$  is the standard deviation of a NYSE size-ranked portfolio  $p$  on day  $t$ ,  $I_{POST,t}$  is an indicator variable that equals one beginning on May 1, 1975 and zero otherwise, and

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<sup>7</sup>Schwert and Seguin (1990) find that a single factor model of standard deviations describes the cross-sectional and time series behavior of portfolio volatility as well as more complicated GARCH-type specifications, and better than linear variance-based specifications. Models similar in spirit can be found in Ng(1991) and Bodurtha and Mark (1991).

$\hat{\sigma}_t^{OTC}$  is the estimated standard deviation of returns on the equal weighted NASDAQ portfolio for day  $t$  conditional on 12 previous daily returns.<sup>8</sup>

The results, reported in Table 3, suggest that the single-factor volatility model works well in this context. Reported  $R^2$ 's are lower than those in Schwert and Seguin (1990, Table VI), indicating that daily volatility is more difficult to predict than the monthly variances they examined. As with Schwert and Seguin, portfolios of large firms experience a better fit to the single factor model ( $R^2$ ), although, unlike Schwert and Seguin, large firms have a greater sensitivity to aggregate heteroskedasticity ( $\gamma_1$ ) than do portfolios of small firms.

Of principal interest, however, are the estimates of  $\gamma_2$ . These numbers estimate the change in NYSE portfolio variance after conditioning on changes in our NASDAQ proxy for aggregate volatility. These estimates support our earlier findings, and are remarkable in their uniformity: volatility for every NYSE portfolio fell, with changes ranging from 0.0025 to 0.0035. Based on the average volatilities of the relevant NYSE portfolios, these estimates represent reductions in portfolio volatility following May Day of between 18% and 24%.

#### 4.3 Autocorrelations and variance ratios

The results of subsection 4.1 indicate that volatilities for individual NYSE listed stocks fell more than the volatilities of their NASDAQ counterparts following the deregulation of commissions after controlling for firm size. In the previous subsection, we confirmed the volatility decline while controlling for conditional heteroskedasticity. Further, we observed volatility declines even for large, diversified portfolios. This last finding indicates that declines in volatility are not entirely attributable to changes in firm-specific factors. In this section, we calculate variance ratios for both individual firms and portfolios to assess how closely stock prices follow a random walk both before and after the deregulation of commissions (see Lo and MacKinlay (1988) for the relevant statistical foundations).

Random walks are not a necessary condition for market efficiency and deviations from a random walk do not necessarily imply any sort of noise. For example, in numerous

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<sup>8</sup>Specifically,  $\hat{\sigma}_t^{OTC}$  is the fitted value from a regression of the daily portfolio return standard deviation on its lagged values:

$$\hat{\sigma}_t^{OTC} = \sum_{i=1}^{12} \beta_i \sigma_{t-i}^{OTC} + \varepsilon_t$$

where, again following Schwert and Seguin (1990),  $\sigma_t^{OTC}$  is the unsigned return to the aggregate (NASDAQ) portfolio multiplied by  $\sqrt{\pi/2}$ .

general equilibrium models expected returns vary over time.<sup>9</sup> However, if intertemporal marginal rates of substitution or other relevant parameters of the economy change sufficiently slowly, predictability in returns at short intervals may be associated with either market frictions, especially non-synchronous trading, or inefficiencies.<sup>10</sup> Thus, we use the random walk as a benchmark, and examine the autocorrelations during 1974-1976.

We investigate the variability of a five day return relative to the variability of daily returns using the following variance ratio statistic:

$$VR = \frac{\sigma_w^2}{5\sigma_d^2},$$

where  $\sigma_d^2$  is the variance of daily returns and  $\sigma_w^2$  is the variance of five (trading) day returns. As Lo and MacKinlay (1988) show, if returns are uncorrelated, the variance ratio should be approximately one, negative autocorrelation leads to estimated variance ratios below one, and positive autocorrelation generates estimated variance ratios above one.

Because we do not employ overlapping observations to compute variance ratios, the caveats of Richardson and Stock (1989) do not apply. Nonetheless, we calculate simulation-based standard errors to ensure that asymptotic theory does not lead to erroneous inference in our finite sample.

Table 4 presents results for both portfolio returns and individual security returns. After May 1, 1975, every portfolio behaves more like a random walk: weekly to daily variance ratios fall toward unity. However, because of the short sample period, we do not have the power to distinguish these changes from zero. Notice also that all variance ratios remain substantially above one, a consequence of substantial positive index autocorrelation.

The results for individual securities require more care in interpretation. As it did for the portfolios, the mean cross-sectional variance ratio falls for all five quintiles, though the decline is, again, not statistically significant. For stocks in the larger quintiles, this movement is towards unity, indicating behavior closer to a random walk.

For smaller stocks, however, mean cross-sectional variance ratios move away from one. We believe that this result is spurious and attributable to bid-ask bounce. Kaul and Nimalendran (1990) demonstrate that individual security weekly-to-daily variance ratios are typically less than one when returns are calculated using transaction prices and typically greater than one when bid-ask averages are used. The difference is due to the negative autocorrelation induced by the bid-ask bounce that reduces the weekly to daily variance

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<sup>9</sup> Examples include the consumption CAPM of Breeden (1979) and Cox, Ingersoll, and Ross (1985), and the differential information model of Wang (1993).

<sup>10</sup>Lo and MacKinlay (1988), Mech (1993) and others argue that non-synchronous trading can explain return predictability, while Lehmann (1990) and Jegadeesh and Titman (1993) claim that predictability is due to inefficiencies. See Boudoukh, Richardson and Whitelaw (1993) for an excellent review.



ratio. Because CRSP does not provide closing bid and ask quotes for NYSE securities, we cannot determine with certainty whether variance ratios of returns purged of bid-ask bounce would increase above one for the smaller quintiles. However, the magnitude of the effect reported by Kaul and Nimalendran for NASDAQ-NMS securities would be sufficient. Thus we conclude that both portfolios and individual firms behave more like random walks following the abolition of fixed commissions.

## **V. Conclusions and Policy Implications:**

The empirical results of this study are straightforward. Jarrell (1984) reports that following the reduction in the commission portion of transaction costs, volume in NYSE listed shares dramatically increased. The contribution of this study is to indicate that the effect of this volume increase was to reduce all measures of volatility, regardless of whether they are estimated using individual firms, portfolios, or net of market volatility. Further, in the low cost, high volume regime, prices were better characterized as a random walk. This last finding is at least consistent with the hypothesis that any noise component in prices was mitigated by the increase in volume.

The goals of this study are intentionally modest. We believe that our results complement Umlauf (1993) and represent an additional clinical data point to the debate concerning the probable impact of the imposition of a transaction tax. Our results, when combined with those of Umlauf, suggest that the logic of increasing transaction taxes to reduce the impact of noise traders and, therefore, to reduce volatility, does not withstand empirical scrutiny. Indeed, our results indicate that increasing transaction costs through any avenue may well have an effect exactly opposite from that intended.

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**Table 1**  
**Summary Statistics**

NYSE stocks have been classified into five quintiles based on their market value of outstanding equity as of April 30, 1975. Firms in the NASDAQ control sample are then assigned using the same size breakpoints as the NYSE sample. Descriptive statistics given for both samples include the cross-sectional average price per share on April 30, 1975 and the cross-sectional average of individual security daily return variance for the 504 trading days (two years) surrounding May 1, 1975. The average proportional bid-ask spread is calculated for the NASDAQ sample uses closing quotes.

NYSE					NASDAQ				
Quintile	Number of firms	Average market value (millions)	Average share price	Average log return variance ( $\times 10^{-4}$ )	Number of firms	Average market value (millions)	Average share price	Average log return variance ( $\times 10^{-4}$ )	Average proportional spread
1 (smallest)	270	\$15.6	6.30	17.71	1,035	\$9.6	6.97	11.39	.1346
2	270	42.3	11.44	9.83	205	41.6	15.88	7.36	.0626
3	270	102.2	16.99	7.60	115	93.5	21.17	5.91	.0460
4	270	276.1	24.68	6.47	65	281.9	33.89	7.28	.0313
5 (largest)	269	1,713.4	39.63	4.95	37	2,726.7	41.96	5.48	.0263
All Firms	1,349	428.9	19.79	9.32	1,457	101.9	11.43	10.06	.1101

**Table 2**  
**Volatility before and after May 1, 1975**

Cross-sectional averages of changes in individual security volatility from the year before the deregulation of NYSE commission rates on May 1, 1975 to the year after. NYSE stocks are classified into five quintiles based on their market value of outstanding equity on April 30, 1975. The NASDAQ control sample is then assigned using the same size breakpoints as the NYSE sample. Volatility changes are measured as  $\Delta\sigma = \sigma_{\text{post}} - \sigma_{\text{pre}}$  (the change in standard deviations of daily returns over the relevant time period), and averages across firms are reported for each NYSE portfolio and each NASDAQ control portfolio. Bootstrapped standard errors based on 500 simulations are given in parentheses and take into account cross-sectional correlation. Significance levels, reported in the final column, are calculated similarly.

Quintile	NYSE			NASDAQ			Prob( $\Delta\sigma_{\text{NYSE}} > \Delta\sigma_{\text{NASDAQ}}$ )
	$\sigma_{\text{pre}}$	$\sigma_{\text{post}}$	$\Delta\sigma$	$\sigma_{\text{pre}}$	$\sigma_{\text{post}}$	$\Delta\sigma$	
1 (smallest)	.0302	.0246	-.0056 (.0005)	.0165	.0143	-.0022 (.0003)	0.000
2	.0236	.0191	-.0045 (.0004)	.0163	.0125	-.0038 (.0003)	0.002
3	.0215	.0165	-.0050 (.0004)	.0160	.0113	-.0047 (.0003)	0.120
4	.0205	.0152	-.0053 (.0004)	.0206	.0140	-.0066 (.0005)	0.996
5 (largest)	.0185	.0132	-.0053 (.0004)	.0183	.0132	-.0051 (.0004)	0.318

**Table 3**  
**Time series regression analysis of volatility around May 1, 1975**

Following the framework of Schwert and Seguin (1990), we estimate the following regression:

$$\delta_{pt}^{NY} = \gamma_0 + \gamma_1 \delta_{t-1}^{OTC} + \gamma_2 I_{POST,t} + \epsilon_t.$$

The regression extends from one year before to one year after the deregulation of NYSE commissions on May 1, 1975.  $\delta_{pt}^{NY}$  is the standard deviation of NYSE size-ranked portfolio  $p$  on day  $t$ . These size ranked NYSE portfolios are equally weighted, and are assigned using quintiles based on market value of outstanding equity on April 30, 1975.  $I_{POST,t}$  is an indicator variable that equals one for all observations on May 1, 1975 or later and zero otherwise.  $\delta_{t-1}^{OTC}$  is the estimated standard deviation of returns for day  $t$  on an equal weighted portfolio of all NASDAQ securities used in Table 2 conditional on 12 previous daily returns. It is calculated as the fitted value from a regression of the daily portfolio return standard deviation on its lagged values:

$$\sigma_{t-1}^{OTC} = \sum_{i=1}^{12} \beta_i \sigma_{t-i}^{OTC} + \epsilon_t$$

where  $\sigma_{t-1}^{OTC}$  is the unsigned return to the NASDAQ portfolio multiplied by  $\sqrt{\pi/2}$ . Standard errors are in parentheses.

Portfolio	$\gamma_0$	$\gamma_1$	$\gamma_2$	R <sup>2</sup>	$\overline{\sigma}_{NY}$
1 (smallest)	.0107 (.0034)	.5825 (.2988)	-.0035 (.0014)	0.042	.0116
2	.0074 (.0028)	.7337 (.2507)	-.0024 (.0011)	0.053	.0105
3	.0065 (.0026)	.8194 (.2338)	-.0029 (.0011)	0.081	.0102
4	.0053 (.0026)	.9826 (.2410)	-.0026 (.0011)	0.088	.0106
5 (largest)	.0040 (.0028)	.1978 (.2542)	-.0029 (.0013)	0.109	.0111

**Table 4**  
**Weekly to Daily variance ratios for NYSE portfolio and individual security returns before and after May 1, 1975**

This table presents estimates of weekly-to-daily variance ratios  $VR = \sigma_w^2 / 5\sigma_d^2$ , where  $\sigma_d^2$  is the variance of daily returns and  $\sigma_w^2$  is the variance of five (trading) day returns. Numbers reported under Portfolios are estimates using the returns to NYSE size-ranked portfolios that are equally weighted and based on the market value of common equity on April 30, 1975. Statistics reported under Individual Securities are the cross-sectional means of variance ratios across all component securities for each portfolio. Pre-event variances are calculated using data for one year prior to May 1, 1975; post-event variances use daily data for one year beginning May 1, 1975. AVR is the difference between the pre- and post-event ratios. Corresponding p-values for the test of whether  $\Delta VR \geq 0$ . Both p-values and standard errors (reported in parentheses) are calculated using Monte-Carlo simulations that accommodate cross-sectional correlations.

Portfolio	Portfolios			Individual Securities		
	VR <sub>pre</sub>	VR <sub>post</sub>	AVR	VR <sub>pre</sub>	VR <sub>post</sub>	AVR
1 (smallest)	2.320 (0.355)	2.109 (0.267)	-0.211 <i>p</i> =.275	0.904 (0.066)	0.885 (0.038)	-0.002 <i>p</i> =.454
2	2.087 (0.269)	1.764 (0.248)	-0.323 <i>p</i> =.259	1.101 (0.057)	0.925 (0.035)	-0.083 <i>p</i> =.268
3	1.939 (0.260)	1.688 (0.266)	-0.251 <i>p</i> =.207	1.041 (0.059)	0.964 (0.036)	-0.070 <i>p</i> =.142
4	1.909 (0.241)	1.584 (0.218)	-0.325 <i>p</i> =.184	1.160 (0.063)	1.048 (0.041)	-0.104 <i>p</i> =.122
5 (largest)	1.640 (0.253)	1.306 (0.209)	-0.334 <i>p</i> =.257	1.214 (0.086)	1.099 (0.051)	-0.113 <i>p</i> =.205
Equally weighted (all firms)	1.602 (0.248)	1.278 (0.198)	-0.324 <i>p</i> =.208	1.066 (0.060)	0.984 (0.037)	-0.068 <i>p</i> =.190