THE JOURNAL OF FINANCE • VOL. XLVIII, NO. 1 • MARCH 1993

# A Simple Measure of Price Adjustment Coefficients

ASWATH DAMODARAN*

**ABSTRACT**

One measure of market efficiency is the speed with which prices adjust to new information. We develop a simple approach to estimating these price adjustment coefficients by using the information in return processes. This approach is used to estimate the price adjustment coefficients for firms listed on the NYSE and the AMEX as well as for over-the-counter stocks. We find evidence of a lagged adjustment to new information in shorter return intervals for firms in all market value classes, and some support for the proposition that prices adjust much more slowly and with more noise for smaller firms.

WHILE THERE IS FAIRLY widespread agreement among financial economists that an efficient market is one in which information is reflected quickly in prices, there is much more dissension on how to measure the speed of the price adjustment process and whether financial markets are efficient on this measure. While some of the early studies of block trading (Dann, Mayers, and Rabb (1977), Kraus and Stoll (1972)) suggested that stock prices adjusted very quickly, if not instantaneously, to new information, these findings have been challenged in recent years by studies which have taken advantage of the increased availability of daily returns and transactions data. There are two broad findings that emerge from these studies. First, there is evidence of a lagged adjustment[1] to news on an intraday basis. Second, there are differences across firms in the speed[2] with which prices incorporate information.

The primary contribution of this paper is the development of a simple approach to estimating price adjustment coefficients for different return intervals as functions of the return variances in these intervals. The derivation of the measure draws on the description of the price process contained in Amihud and Mendelson (1987) as a combination of an intrinsic value adjustment factor and a noise term. The approach developed here is used to estimate price adjustment coefficients for all firms listed on the CRSP daily

* Stern School of Business, New York University. I would like to thank an anonymous referee for helpful comments.

[1] Patell and Wolfson (1984) and Woodruff and Senchack (1988) come to this conclusion with earnings reports while Hasbrouck and Ho (1987) use serial correlations in prices to confirm it.

[2] Woodruff and Senchack (1988) conclude that prices adjust much more quickly to earnings announcements which contain good news. Jennings and Stark (1981) find that prices adjust much more quickly to news for firms with options listed on them.

returns tape from 1977 to 1986. The estimates are made separately for two five-year periods, the first extending from January 1977 to December 1981 and the second covering January 1982 to December 1986. We find evidence of a lagged adjustment to new information for short return intervals (one to three days) in both time periods, and some evidence of an overreaction in longer return intervals in the latter period.

The price adjustment coefficients developed here are an intuitively appealing way of measuring the information efficiency of a particular stock, a group of stocks or the overall market. In an illustration, we classify firms on the NYSE and AMEX on the basis of market value into five classes and estimate price adjustment coefficients for each class and compare these coefficients with estimates obtained for NASDAQ stocks with data available on the CRSP NMS tape. We find that NASDAQ firms have lower price adjustment coefficients, higher intrinsic variance and more noise in their price processes than stocks listed on the major exchanges, but we do not find significant differences in price adjustment coefficients across market value classes for NYSE and AMEX stocks.

## I. A Partial Price Adjustment Model

### *A. The Estimation of Price Adjustment Coefficients*

We start with the simple model of price behavior described in Amihud and Mendelson, which distinguishes between the intrinsic value $V_t$ and the observed price $P_t$ by allowing for both market-structure-related and information noise as well as imperfect price adjustments to value changes.

$$P_t - P_{t-1} = g[V_t - P_{t-1}] + u_t \tag{1}$$

where $V_t$ and $P_t$ are in logarithms and $g$ is the price adjustment coefficient $[0 < g < 2]$. The last term $u_t$ is a noise term, the magnitude of which is determined by information-related factors (such as liquidity trading and noisy information) and market-structure-related factors (such as the bid-ask spread, dealer inventory positions, and the discreteness of stock prices). If we define $v^2$ to be the variance of the intrinsic value process and $\sigma^2$ to be variance of the noise term, the observed return variance can be decomposed into three components:

$$\mathrm{Var}(R_t) = \underset{\text{Intrinsic variance}}{v^2} + \underset{\text{Noise}}{2\sigma^2} + \underset{\text{Price adjustment effect}}{\left[\left(\frac{g}{2-g} - 1\right)v^2 + \left(\frac{2}{2-g} - 2\right)\sigma^2\right]} \tag{2}$$

If prices adjust slowly to information ($g < 1$), the price adjustment effect will be negative and lead to lower observed return variances, whereas market overreaction to news ($g > 1$) will have the opposite impact. Concurrently, higher (lower) bid-ask spreads will lead to more (less) noise and higher (lower) return variances.

There are three critical assumptions that we make that enable us to arrive at our formulations for the price adjustment coefficients. The first is that the intrinsic value $\{V_t\}$ process is i.i.d. and follows a random walk. This is not an uncommon assumption in other papers in this area, but it is difficult to justify stationarity in the process over very long time periods. While the longest return interval that we use in our empirical tests is twenty days, we still assume return stationarity over five-year time periods—1977 to 1981 and 1982 to 1986. To the degree that this assumption is violated, our estimates are likely to be flawed. The second is that the noise terms and intrinsic value processes are independent, an assumption of convenience that may be set aside if one is willing to explicitly model the covariance between the two processes. The third is that the price adjustment coefficient approaches one as the return interval is extended. This assumption is justifiable on both intuitive and empirical grounds. It seems reasonable to assume that giving financial markets more time to adjust to new information will result in a more complete price adjustment process. Studies of earnings announcements also seem to indicate that any lags due to incomplete price adjustment last a few days at the most (and can generally be measured in hours for larger firms). Furthermore, the price adjustment coefficient for longer intervals approaches one rapidly even for very low or very high values of $g_1$.

To derive our measure of $g$, we use the variances in different return intervals. We define $R_{jt}$ to be the returns in time period $t$ where each interval is of length $j$. The variance in these returns can be written as:

$$\operatorname{Var}(R_{jt}) = \left[\frac{g_j}{2 - g_j} jv^2 + \frac{2}{(2 - g_j)}\sigma^2\right] \tag{3}$$

where,

$\operatorname{Var}(R_{jt})$ = Variance in observed return time series of returns of interval $j$
$g_j$ = Price adjustment coefficient for $j$-interval returns

Note that $v^2$ is linear in $j$ because $\{\partial V_t\}$ is an i.i.d. random variable, and the intrinsic variance and noise processes are assumed to be independent of each other. The noise variance becomes a smaller component of the total return variance as the return interval is increased because of the negative serial correlation induced by the bid-ask spread.

Assume that a sequence of $n$ observed unit-interval returns $\{R_{11}, R_{12} \ldots R_{1n}\}$ is used to estimate the variances for different return intervals ($j = 1, 2, \ldots k$). We assume that $k$ is of sufficient length to allow $g_k = 1$. Then the variance in $j$-interval returns and the equivalent variance[3] in $k$-interval

[3] If $g_k = 1$, equation (3) simplifies to the following: $\operatorname{Var}(R_{kt}) = kv^2 + 2\sigma^2$.

returns can be written as follows:

$$\mathrm{Var}(R_{jt}) - \frac{\mathrm{Var}(R_{kt})}{k} = v^2\left[\frac{g_j}{(2-g_j)} - 1\right] + 2\sigma^2\left[\frac{1}{(2-g_j)} - \frac{1}{k}\right] \tag{4}$$

where the noise and the intrinsic variance components can be written as functions of the covariance and variance in $k$-interval returns,

$$\sigma^2 = -\mathrm{Cov}(R_{kt}, R_{kt-1}) \tag{5}$$

$$v^2 = \frac{\mathrm{Var}(R_{kt}) + 2\,\mathrm{Cov}(R_{kt}, R_{kt-1})}{k} \tag{6}$$

Substituting for $v^2$ and $\sigma^2$ in (4), and solving through for $g$, we get

$$g_j = \frac{2\left[\dfrac{\mathrm{Var}(R_{jt})}{j} + \dfrac{\mathrm{Var}(R_{kt})}{k}(j-1) + \dfrac{\mathrm{Cov}(R_{kt}, R_{kt-1})}{j}\right]}{\dfrac{\mathrm{Var}(R_{jt})}{j} + \dfrac{\mathrm{Var}(R_{kt})}{k}(2j-1) + \dfrac{2\,\mathrm{Cov}(R_{kt}, R_{kt-1})}{k}} \tag{7}$$

Equation (7) helps us estimate the price adjustment coefficients for return intervals until $j = k - 1$, and requires no information beyond that available in the time series of unit-interval returns data.

Amihud and Mendelson (1989) extend their basic model of the return process described in equation (1) to examine the price adjustment coefficients of portfolios. They note that both the spread and idiosyncratic information components can be diversified away in a portfolio, and that the serial correlation in portfolio returns is a direct function of the price adjustment coefficient for systematic information.[4] They show that,

$$g_s = 1 - \rho_{t,t-1} \tag{8}$$

where $g_s$ is the price adjustment coefficient relative to systematic information and $\rho_{t,t-1}$ is the serial correlation in returns. In the special case, where the price adjustment coefficient is the same for systematic and firm-specific information, this approach will yield estimates that can be compared to those obtained from equation (7). In the more general case, where the price adjustment coefficients are different, this approach will provide a measure of the speed with which prices respond to new market-wide information while

[4] Amihud and Mendelson assume that the price adjustment coefficient is the same for both market-wide as well as firm-specific information. Their general formulation for portfolio variance can be written as follows:

$$\mathrm{Var}(R_{pt}) = \frac{g}{2-g}\beta_p^2 t^2 + \frac{g}{2-g}\sum_{i=1}^{N} w_i^2 v_i^2 + \frac{g}{2-g}\sum_{i=1}^{N} w_i^2 \sigma_i^2$$

They note that as $N$ gets larger the second and third terms drop out, leaving the price adjustment coefficient as the only source of serial covariance.

equation (7) measures the speed of adjustment to all new information.[5] This statistic has to be interpreted with some caution, given the findings in Lo and MacKinlay (1990) that returns on larger stocks systematically lead returns on smaller stocks, which in turn can induce cross-auto-correlations in security returns and positive autocorrelations in portfolio returns. By assuming that all firms in the index adjust to market-wide information at the same rate "$g_s$," equation (8) is likely to result in an underestimation of market-wide price adjustment coefficients.[6]

### *B. Issues Related to Estimation*

While equation (7) provides us with a measure of the price adjustment coefficient, there are several issues related to estimation that need to be addressed prior to empirically testing its usefulness. The first relates to the unit measure for unit returns since the empiricist has a choice of transactions data, daily returns, or even longer return periods. While longer period returns may not meet our requirements of stationarity in the intrinsic value process, returns over very short intervals (such as fifteen minutes) may be too noisy to provide precise estimates.

The second issue relates to the limiting interval. Since we assume full price adjustment by this interval, the assumption is a key one and determines the magnitude of the estimated price adjustment coefficients. We choose a limiting interval of twenty days to conduct our tests, but examine the sensitivity of our results to other limiting intervals.[7]

The third issue relates to the precision and unbiasedness of the estimates obtained from equation (7). Harris (1989) notes that the covariances obtained from daily returns are biased upwards, probably because of positive correlation in the intrinsic value process. Another problem arises from the use of overlapping data to estimate the variances in different return intervals. By taking covariances from the limiting interval returns, we mitigate the first problem somewhat, but the estimates for individual firms remain noisy. To avoid the second problem, we report the average price adjustment coefficients

[5] The price adjustment coefficient $g$ estimated in equation (9) can be written as a weighted average of the market-wide adjustment coefficient $g_s$ and the firm-specific adjustment coefficient $g_e$.

$$\frac{g}{2-g} = \frac{\frac{g_s}{2-g_s}\beta^2\sigma_m^2 + \frac{g_e}{2-g_e}\sigma_e^2}{\sigma^2}$$

where $\beta$ is the firm's beta, $\sigma_m^2$ is the variance in market returns, $\sigma_e^2$ is the variance in residual returns and $\sigma^2$ is the variance in the firms' returns.

[6] Note that if larger firms adjust faster to market-wide information than smaller firms, as Lo and MacKinlay conclude, $\rho_{t,t-1}$ will be positively biased and the price adjustment coefficient will be downward biased.

[7] If $g_{20} \neq 1$, the estimates we obtain for $g_1$ through $g_{19}$ will be biased, downwards if $g_{20} < 1$, and upwards if $g_{20} > 1$.

for fairly large subsamples of stocks, since the estimators are much more reliable when large cross-sectional samples are used.

The fourth issue relates to the nontrading of securities. Researchers (Scholes and Williams (1977) and Dimson (1979)) note that infrequent trading can lead to biased estimates of the auto-covariances and variances in returns, which in turn can impact the estimates of price adjustment coefficients obtained from equation (7). To be more specific, the price adjustment coefficients will be biased downwards for firms which do not trade frequently (such as smaller firms).

The efficacy and robustness of this approach to estimating price adjustment coefficients was tested, using a series of simulations, each of which generated 1250 price observations from known values of $\sigma^2$, $v^2$, and $g_1$. Equations (5), (6), and (7) were used to obtain the corresponding estimates from the generated price data with 20 days used as the limiting interval. While the estimates were noisy for individual simulations, the averages obtained over 1000 simulations were close to true values.[8]

## II. Empirical Methods and Results

### A. Sample and Methodology

To test the efficacy of the model in estimating price adjustment coefficients, we use daily returns from January 1, 1977 to December 31, 1986. Our sample includes all firms which are listed on the CRSP NYSE/AMEX daily returns tape and have data available on them for every trading day from January 1, 1977 to December 31, 1986. We do not, at this stage, screen returns for nontrading but use all returns in making our firm-specific estimates of the variances with return frequencies ranging from one day to twenty days[9]—

$$\operatorname{Var}(R_{jt}) = \frac{1}{\left(\frac{n}{j} - 1\right)} \left[ \sum_{t=1}^{n/j} \left(R_{jt} - \hat{R}_j\right) \right]$$

where $n$ is the number of daily returns, $j$ is the intervaling frequency ($j = 1, 2, \ldots 20$), $R_{jt}$ is the return in interval $t$ and $\hat{R}_j$ is the average return for $j$-interval returns over time period. Using the daily returns from January 1, 1977 to December 31, 1986, we compute successive nonoverlapping twenty-

[8] We ran 1000 simulations, each of which generated 1250 price observations, with $\sigma^2 = 0.02$, $v^2 = 0.05$ and $g_1 = 0.7$. Estimated values for these variables were obtained using equations (6), (7), and (9) after each simulation. The averages across all 1000 simulations were very close to the true values: $\sigma^2 = 0.0196, v^2 = 0.0505$, and $g_1 = 0.6910$.

[9] The arbitrary choice of twenty days as the limiting return interval may trouble the reader. To evaluate the effect of this choice on the coefficient estimates, the coefficients are re-estimated using alternative limiting intervals, ranging from ten to forty periods, and the results are encouraging. The coefficient estimates stabilize for limiting intervals beyond fifteen days.

day returns for each firm.[10] We estimate the covariances in these twenty-day returns:

$$\text{Cov}(R_{kt}, R_{kt-1}) = \frac{1}{\left(\frac{n}{k} - 2\right)} \left[ \sum_{t=1}^{n/k} \left(R_{kt} - \hat{R}_k\right)\left(R_{kt-1} - \hat{R}_k\right) \right]$$

### *B. Coefficient Estimates for Five-Year Periods: 1977 to 1981 and 1982 to 1986*

We estimate the price adjustment coefficients for all the firms in our sample over two five-year time periods, the first covering 1977 to 1981 and the second 1982 to 1986. The cross-sectional averages[11] and standard deviations of price adjustment coefficients for all firms in our sample, as well as the medians of the distributions of the coefficients are reported in Table I for the 1977 to 1981 and 1982 to 1986 subperiods.

Several interesting results emerge from this table. First, there seems to be evidence of a lagged adjustment to new information, since the price adjustment coefficients are significantly less than one for return intervals up to five days, using a simple median *t*-statistic[12] based upon a binomial test. For instance, our estimates of the price adjustment coefficients for the 1977 to 1981 period suggest that approximately 67% of new information is reflected in prices by the end of the day[13] on which it is released and almost 95% by the end of the next day. This puts us in agreement with Patell and Wolfson (1984), Woodruff and Senchack (1988), and Hasbrouck and Ho (1987) who all find evidence of lagged price adjustment in transactions data. Furthermore, since delayed price adjustment processes induce positive autocorrelation in returns, these results would be consistent with Roll's findings (1984) of

[10] The starting point and the successive time intervals for calculating twenty-day returns are the same for all the firms in the sample. While this means that there will be some nontrading days included in return calculation, it also means that the sample size is the same for all the firms in the sample.

[11] While these are unweighted estimates, the weighted averages and standard errors of the price adjustment coefficients, with market values as weights, were also estimated. The one-day price adjustment coefficients increase to 0.69 for the 1977 to 1981 time period and to 0.70 for the 1982 to 1986 time period.

[12] The cross-sectional distributions of the price adjustment coefficients tend to be positive skewed and deviate from normality significantly, as evidenced by low Kolmogorov D statistics. The price adjustment coefficients for the lower (higher) return intervals tend to have much lower (higher) kurtosis than would a normal distribution. Since these deviations from normality suggest that *t*-statistics may not be reliable, we report nonparametric statistics for testing hypotheses wherever possible.

[13] The low one-day price adjustment coefficients are cause for some concern, since they are much lower than our priors would lead us to expect. There are a couple of reasons for this finding. First, these coefficients are extremely noisy on a firm-specific basis. When cross-sectional averages of the variances and covariances are used to estimate the coefficients, we eliminate this problem, but we cannot estimate standard errors. Second, nontrading on stocks will reduce the adjustment coefficients, especially for the short intervals.

## Table I
## Estimates of Price Adjustment Coefficients: 1977 to 1981 and 1982 to 1986 Cross-Sectional Averages of Firm-Specific Estimates

The price adjustment coefficients are estimated from daily returns for the two time periods by setting $g_j = 2[\mathrm{Var}(R_{jt})/j + (\mathrm{Var}(R_{kt})/k)(j-1) + \mathrm{Cov}(R_{kt}, R_{kt-1})/j]/[\mathrm{Var}(R_{jt})/j + (\mathrm{Var}(R_{kt})/k)(2j-1) + 2\,\mathrm{Cov}(R_{kt}, R_{kt-1})/k]$, where $\mathrm{Var}(R_{jt})$ is the variance in returns with interval $j$ and $\mathrm{Cov}(R_{jt}, R_{jt-1})$ is the serial covariance in returns with interval $j$. (We set $g_j = 0.01$ if $g_j \le 0$, and $g_j = 2$ if $g_j \ge 2$.) The standard deviations are estimated from the cross-sectional distribution of the coefficients.

| Interval | No. of Firms | 1977–81 Mean | 1977–81 SD | 1977–81 Median | 1977–81 t (Median)[a] | 1982–86 Mean | 1982–86 SD | 1982–86 Median | 1982–86 t (Median)[a] | Matched Pair t-Statistic[b] | Serial Corr. NYSE Index[c] |
|---|---|---|---|---|---|---|---|---|---|---|---|
| 1 day | 1453 | 0.6706 | 0.8440 | 0.6409 | 6.54** | 0.6795 | 0.8468 | 0.6476 | 6.42** | 0.30 | 0.1557 |
| 2 days | 1453 | 0.9468 | 0.1140 | 0.8248 | 5.86** | 0.9568 | 0.1160 | 0.8387 | 5.60** | 1.18 | 0.0913 |
| 3 days | 1453 | 0.9838 | 0.1000 | 0.9420 | 5.55** | 0.9945 | 0.1020 | 0.9517 | 4.79** | 1.35 | 0.0588 |
| 4 days | 1453 | 0.9948 | 0.0600 | 0.9765 | 4.84** | 1.0018 | 0.0600 | 0.9837 | 3.46** | 1.92 | 0.0423 |
| 5 days | 1453 | 0.9967 | 0.0360 | 0.9874 | 4.54** | 1.0043 | 0.0400 | 0.9950 | 1.83 | 2.84** | 0.0682 |
| 6 days | 1453 | 1.0000 | 0.0280 | 0.9953 | 2.31* | 1.0049 | 0.0280 | 1.0000 | 0.03 | 1.75 | 0.0441 |
| 7 days | 1453 | 1.0008 | 0.0200 | 0.9979 | 1.75 | 1.0055 | 0.0240 | 1.0025 | 1.83 | 2.31* | −0.0068 |
| 8 days | 1453 | 1.0003 | 0.0160 | 0.9980 | 1.66 | 1.0034 | 0.0160 | 1.0013 | 1.35 | 2.14* | 0.0430 |
| 9 days | 1453 | 1.0008 | 0.0140 | 0.9996 | 0.62 | 1.0048 | 0.0160 | 1.0032 | 4.40** | 3.38** | 0.0904 |
| 10 days | 1453 | 1.0005 | 0.0130 | 0.9997 | 0.68 | 1.0028 | 0.0120 | 1.0019 | 3.60** | 3.15** | −0.0294 |
| 11 days | 1453 | 1.0021 | 0.0120 | 1.0014 | 1.80 | 1.0039 | 0.0110 | 1.0030 | 4.82** | 1.80 | 0.0139 |
| 12 days | 1453 | 1.0003 | 0.0110 | 0.9996 | 0.59 | 1.0028 | 0.0100 | 1.0022 | 4.40** | 3.21** | −0.0351 |
| 13 days | 1453 | 1.0009 | 0.0100 | 1.0002 | 0.28 | 1.0020 | 0.0090 | 1.0012 | 2.82** | 1.77 | −0.0106 |
| 14 days | 1453 | 1.0007 | 0.0100 | 0.9998 | 0.37 | 1.0017 | 0.0080 | 1.0013 | 2.87** | 1.44 | 0.0018 |
| 15 days | 1453 | 1.0000 | 0.0090 | 0.9995 | 1.30 | 1.0007 | 0.0070 | 1.0017 | 0.53 | 1.35 | −0.0081 |
| 16 days | 1453 | 1.0007 | 0.0080 | 1.0002 | 0.62 | 1.0009 | 0.0070 | 1.0004 | 1.27 | 0.48 | 0.0762 |
| 17 days | 1453 | 1.0007 | 0.0060 | 1.0007 | 1.10 | 1.0021 | 0.0060 | 1.0007 | 4.90** | 3.04** | 0.0410 |
| 18 days | 1453 | 0.9999 | 0.0050 | 0.9997 | 1.46 | 1.0011 | 0.0050 | 1.0008 | 3.13** | 2.76** | −0.0315 |
| 19 days | 1453 | 1.0009 | 0.0040 | 1.0006 | 1.03 | 1.0013 | 0.0040 | 1.0011 | 3.92** | 1.18 | 0.0177 |
| 20 days | 1453 | 1.0000 | 0.0040 | 1.0000 | 0.00 | 1.0000 | 0.0040 | 1.0000 | 0.00 | 0.00 | −0.0072 |

** Significant at 99% level; * Significant at 95% level.

[a] With large samples, the actual number of firms with $g_j > 1$ $(x)$ can be compared to the expected number $(0.5n)$ to arrive at a $t$-statistic: $t_{med} = (x - 0.5n)/0.5\sqrt{n}$ (where $n$ is the number of firms in the sample).

[b] This is a nonparametric test of differences in means with paired samples: $t_{par} = (y - 0.5n)/0.5\sqrt{n}$, where $y$ is the number of firms with $g_j$ in period 1 $> g_j$ in period 2.

[c] The last column reports the serial correlation in the value-weighted NYSE index for the period 1977 to 1986. $[g_{ts} = 1 - r_{t,t-1}]$.

positive autocorrelations for shorter return intervals (because of delayed price adjustment) and negative autocorrelations for longer return intervals (when the bid-ask spread effect dominates). Second, the standard deviations of the price adjustment coefficients are greatest for the daily returns and decline as the return interval is extended. Third, the estimates are very similar for both time periods in our study. The 1982 to 1986 time period has slightly higher price adjustment coefficients than the 1977 to 1981 period, and the coefficients are significantly below one for only the first four days. This is borne out by an examination of the matched-pair $t$-tests which indicate that the latter period had significantly higher price adjustment coefficients for eight out of the twenty time intervals. A puzzling phenomenon of the 1982 to 1986 time period is the appearance of coefficients greater than one at longer intervals (between nine and twelve days). This would either denote a tendency for markets to overreact to new information over these longer intervals, which may explain why DeBondt and Thaler (who use longer period returns) find evidence of market overreaction to earnings announcements, or it could indicate that our model does not fit the data, i.e., that our simplifying assumptions were not met. Fourth, our estimates of price adjustment coefficients for one-day returns are significantly lower than those estimated using the portfolio approach developed by Amihud and Mendelson (1989). The last column of Table I reports the serial correlation in returns on the value-weighted NYSE index for the period 1977 to 1986 for different intervaling periods, and the market-wide price adjustment coefficients for different return intervals can be estimated using equation (8).

$$g_{ts} = 1 - \rho_{t,t-1}$$

These systematic adjustment coefficients are much higher than the firm-specific coefficients for the shorter return intervals, suggesting that prices adjust much more quickly to new market-wide information than to new firm-specific information. (For instance, the one-day adjustment coefficient is 0.8443 for market-wide and about 0.67 for the firm-specific information.) It should be noted that if prices adjust faster to market-wide information for larger firms than for smaller firms these market-wide coefficients[14] will be significantly overstated.

### C. Price Adjustment Coefficients and Information Quality

The results in Table I are consistent with a market that adjusts slowly to new information, but do not really cast much light on the reasons or the slow adjustment process. The price adjustment coefficient for an individual stock should be a function of the quality and amount of information that comes out about the firm, and is an intuitively appealing measure of information

[14] The cross-autocorrelations induced by systematic leads and lags in security returns will result in positive autocorrelations in portfolio returns. Lo and MacKinlay find evidence of such systematic leads and lags and correlate them to size (the returns of larger firms lead the returns of smaller firms) and explain a significant portion of the positive correlation in index returns.

efficiency. Firms which are information rich should have speedier (and more complete) price adjustment processes, on average, than information poor firms. To test propositions relating the speed of the price adjustment coefficient to size, we compute price adjustment coefficients for firms classified by market value of equity into five classes. In addition, we estimate a covariance measure of the noise, the variance ratio and the "intrinsic variance" for different market value classes.

$$\text{Noise} = -2(\text{Cov}_{20})$$
$$\text{Intrinsic Variance} = (\text{Var}_{20} + 2\,\text{Cov}_{20})/20$$

where $\text{Cov}_{20}$ is the covariance and $\text{Var}_{20}$ is the variance in twenty-day returns. Note that in the special case where the spread is the only source of noise, the square root of the noise term is the average spread. We also extend this estimation approach to stocks listed[15] on the CRSP NMS tape for two reasons. First, these stocks are much smaller in market value than the smallest stocks in the NYSE/AMEX sample used in Table I and hence provide us with more cross-sectional dispersion in market value and a much stronger test of the relationship between size and the speed of price adjustment. Second, the data for these stocks includes bid and ask prices, the midpoints of which can then be used to estimate price adjustment coefficients without the noise induced by the bid-ask bounce. The cross-sectional medians and standard deviations of price adjustment coefficient estimates for NMS securities[16] (using bid prices, ask prices, midpoint of the bid and ask prices and transactions prices) and firms in different market value classes in the NYSE/AMEX sample are reported in Table II.

When transactions prices are used, the noise term is greater for the over-the-counter stocks than it is for stocks on the major exchanges, and declines with market value for stocks on the NYSE and the AMEX. When the bid or ask prices are used for NASDAQ stocks, the noise term is lower since the spread is eliminated as a source of noise. Consistent with expectations, the intrinsic variance declines with market value for NYSE/AMEX stocks and is higher, on average, for the over-the-counter stocks. On the price adjustment coefficients, the results are mixed. The one-day price adjustment coefficients are not dramatically different across market value classes[17] for the NYSE/AMEX stocks, with values ranging from 0.61 for the smallest firms to 0.72 for the largest firms, and there are only marginal differences in coefficients for longer return intervals. However, the NASDAQ stocks have, on average, significantly lower price adjustment coefficients than the stocks

[15] The sample was restricted to those stocks which had data available on the NMS tape for the entire sample period of 1982 to 1986.

[16] There are 458 over-the-counter firms with data available on the CRSP NMS tape for the 1982 to 1986 period.

[17] These results may at least partially be explained by our sample characteristics and data needs. Since the sample is restricted to firms which had daily returns data available between 1977 and 1986, we introduce a strong bias into our sampling since it is unlikely that many of the smallest firms on the exchanges would survive this screen.

on the major exchanges for one-day, two-day, and five-day return intervals. While nontrading is a problem for some of the smaller stocks, corrections[18] for it do not result in material changes in our conclusions.

The approach used here to estimate price adjustment coefficients can be used to examine the price adjustment process around any "event" that satisfies the following conditions. First, the event should be hypothesized to increase or decrease the speed of price adjustment. Second, a sufficient number of firms should have experienced the event to allow the researcher to draw on the cross-sectional distribution of price adjustment coefficients to accept or reject the hypothesis. Third, there should be sufficient return data available (500 trading days or more if daily return data is used) before and after the event available on each firm to make reasonable estimates of variances in longer return intervals. Damodaran and Lim (1991) use the approach developed in this paper to examine price adjustment coefficients around option listing and conclude that prices adjust much more quickly to new information after the listing of options. Other events that seem to satisfy these conditions include stock splits, first-time listing on the NYSE and dual listing on international exchanges. If transactions data is available, this set will expand to include earnings and dividend announcements.

## III. Conclusions

Studies of market efficiency have generally focused on the question of how quickly information is reflected in prices, and the consensus of those studies that use short-period returns data (daily and transactions data) is that prices adjust slowly to new information. Extending the basic model of returns developed in Amihud and Mendelson, we derive a simple measure of firm-specific price adjustment coefficients that draws on the information in return processes. The approach is used to estimate the price adjustment coefficients of all firms listed on the CRSP daily returns tape between 1977 to 1986 for two five year time periods—1977 to 1981 and 1982 to 1986. We find evidence of lagged price adjustment for short return intervals (one to three days) in both estimation periods. When the price adjustment coefficients for firms in the NYSE/AMEX sample are compared to those for a NASDAQ subsample, we find that the over-the-counter stocks have much lower price adjustment coefficients for return intervals up to five days. We also find evidence that the firms in the latter group have higher intrinsic variance and noise components in the price process than firms on the major exchanges.

[18] By restricting our sample to stocks on the New York Stock Exchange which survived from 1977 and 1986 and had no missing data on the CRSP tape and by using daily returns rather than transactions data, the nontrading problem is reduced significantly. As a further test, the sample was restricted to 740 firms which had less than ten days of nontrading during the 1982 to 1986 time period and the return process parameters were reestimated, with no change in the conclusion.

**Table II**

**Price Adjustment Coefficients for NMS and NYSE / AMEX Stocks: 1982 to 1986: All Firms in Sample With Data Available From 1982 to 1986**

The variance statistics and price adjustment coefficients were estimated for 458 over-the-counter firms and 1290 NYSE and AMEX stocks using daily returns from 1982 to 1986. The NYSE/AMEX firms were classified into five market value classes based upon the average market value of equity between 1982 and 1986. The medians (with the cross-sectional standard deviations in brackets below) are reported for the variance ratio, intrinsic variance, and noise, as well as for the price adjustment coefficients. Since the distribution of these statistics is not normal, we test for differences between classes (NMS vs NYSE/AMEX and across NYSE/AMEX market value classes) with a nonparametric statistic—a chi-squared rank sum statistic.

| | | | | Price-Adjustment Coefficients | | | | |
|---|---|---|---|---|---|---|---|---|
| Size Class | Variance Ratio[a] | Intrinsic Variance[b] | Noise[c] (20-Day Ret.) | One-Day | Two-Day | Five-Day | Ten-Day | Fifteen-Day |
| NMS securities | | | | | | | | |
| Bid price | 0.832365 | 0.001033 | 0.000550 | 0.503742 | 0.808228 | 0.980576 | 1.000419 | 1.000647 |
| | (0.21293) | (0.00063) | (0.00087) | (0.21672) | (0.16739) | (.02574) | (0.00756) | (0.00579) |
| Ask price | 0.823204 | 0.000991 | 0.000530 | 0.502765 | 0.811655 | 0.981638 | 1.000456 | 1.000674 |
| | (0.21144) | (0.00057) | (0.00091) | (0.23331) | (0.17225) | (0.02817) | (0.00766) | (0.00572) |
| (Bid + Ask)/2 | 0.830537 | 0.001002 | 0.000543 | 0.504491 | 0.810043 | 0.980770 | 1.000435 | 1.000734 |
| | (0.21078) | (0.00059) | (0.00090) | (0.22440) | (0.17109) | (0.02776) | (0.00736) | (0.00530) |
| Trade price | 1.138141 | 0.000854 | 0.000871 | 0.510333 | 0.809721 | 0.978904 | 1.000435 | 1.000518 |
| | (0.25322) | (0.00068) | (0.00133) | (0.26867) | (0.19885) | (0.03132) | (0.00771) | (0.00587) |

**Table II**—*Continued*

| Size Class | Variance Ratio[a] | Intrinsic Variance[b] | Noise[c] (20-Day Ret.) | Price-Adjustment Coefficients One-Day | Two-Day | Five-Day | Ten-Day | Fifteen-Day |
|---|---|---|---|---|---|---|---|---|
| NYSE/AMEX Stocks | | | | | | | | |
| Smallest | 1.103571 | 0.000793 | 0.000654 | 0.610644 | 0.834356 | 0.983740 | 1.000848 | 1.000864 |
| | (0.58748) | (0.00055) | (0.00262) | (0.32280) | (0.23919) | (0.03799) | (0.00993) | (0.00614) |
| MV class 2 | 1.082085 | 0.000648 | 0.000548 | 0.701191 | 0.885805 | 0.989921 | 1.000644 | 1.000650 |
| | (0.39780) | (0.00049) | (0.00169) | (0.26610) | (0.18722) | (0.02944) | (0.00715) | (0.00553) |
| MV class 3 | 1.070588 | 0.000469 | 0.000517 | 0.658075 | 0.871954 | 0.985080 | 1.000637 | 1.000651 |
| | (0.29579) | (0.00041) | (0.00129) | (0.28581) | (0.19975) | (0.03113) | (0.00789) | (0.00555) |
| MV class 4 | 1.051614 | 0.000410 | 0.000470 | 0.654707 | 0.871078 | 0.994981 | 1.000418 | 1.000455 |
| | (0.30245) | (0.00031) | (0.00091) | (0.26105) | (0.18871) | (0.03096) | (0.00836) | (0.00554) |
| Largest | 1.023801 | 0.000345 | 0.000430 | 0.715535 | 0.886713 | 0.995973 | 1.000411 | 1.000449 |
| | (0.29391) | (0.00016) | (0.00066) | (0.21179) | (0.17649) | (0.03438) | (0.00876) | (0.00563) |
| All firms | 1.060177 | 0.000433 | 0.000529 | 0.640761 | 0.868743 | 0.990003 | 1.000515 | 1.000555 |
| | (0.15573) | (0.00009) | (0.00043) | (0.10076) | (0.08576) | (0.01436) | (0.00377) | (0.00244) |
| Statistical tests—NMS vs NYSE/AMEX | | | | | | | | |
| $\chi^2$ statistic[d] | 21.55** | 277.00** | 102.40** | 75.30** | 63.00** | 20.50* | 5.50 | 6.20 |
| Statistical tests—Market Value classes—NYSE/AMEX | | | | | | | | |
| $\chi^2$ statistic[d] | 33.00** | 330.00** | 42.40** | 5.99 | 12.00* | 15.92* | 4.74 | 5.62 |

** Significant at 99% level; *Significant at 95% level.

[a] The variance ratio is defined to be the daily return variance/Implied daily variance in 20-day returns.

[b] Intrinsic variance = (Variance in 20-day returns + 21(Covariance in 20-day returns))/20.

[c] The covariances are estimated from 20-day returns. Noise = −2 (Covariance in 20-day returns).

[d] The chi-square is a rank statistic testing for differences between classes.

## REFERENCES


Amihud, Y., and H. Mendelson, 1987, Trading mechanisms and stock returns: An empirical investigation, *Journal of Finance* 42, 533–553.

———, 1989, Index and index-futures returns, Working paper, New York University.

Damodaran, A., and J. Lim, 1991, The effects of option listing on the underlying stocks, return process: A study, *Journal of Banking and Finance* 15, 647–664.

Dann, L., D. Mayers, and R. Raab, 1977, Trading rules, large blocks and the speed of adjustment, *Journal of Financial Economics* 4, 3–22.

DeBondt, W. F. M., and R. H. Thaler, 1985, Does the stock market overreact?, *Journal of Finance* 40, 793–805.

Dimson, E., 1979, Risk measurement when shares are subject to infrequent trading, *Journal of Financial Economics* 7, 197–226.

Harris, L., 1989, Statistical properties of the Roll serial covariance bid/ask spread estimator, *Journal of Finance* 45, 579–590.

Hasbrouck, J., and T. Ho, 1987, Order arrival, quote behavior and the return-generating process, *Journal of Finance* 42, 1035–1049.

Jennings, R. H., and L. T. Stark, 1986, Earnings announcements, stock price adjustment and the existence of options markets, *Journal of Finance* 41, 43–61.

Kraus, A., and H. R. Stoll, 1972, Price impacts of block trading on the New York Stock Exchange, *Journal of Finance* 27, 569–588.

Lo, A. W., and A. C. MacKinlay, 1990, When are contrarian profits due to stock market overreaction?, *The Review of Financial Studies* 3, 175–205.

Patell, J. M., and M. A. Wolfson, 1984, The intraday speed of price adjustment of stock prices to earnings and dividend announcements, *Journal of Financial Economics* 13, 223–252.

Roll, R., 1984, A simple implicit measure of the effective bid-ask spread in an efficient market, *Journal of Finance* 39, 1127–1139.

Scholes, M., and J. T. Williams, 1977, Estimating betas from nonsynchronous data, *Journal of Financial Economics* 5, 309–327.

Woodruff, C. S., and A. J. Senchack, 1988, Intraday price-volume adjustments of NYSE stocks to unexpected earnings, *Journal of Finance* 43, 467–492.