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Stock-Split Post-Announcement Returns: Underreaction or Market Friction?

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Abstract

We explore the relationship between stock splits and subsequent long-term returns during the period from 1950 to 2000. We find that, contrary to much previous research, firms do not exhibit positive long-term post-split returns. Instead, we find that significant positive returns after the announcement date do not persist after the actual date of the stock split. We also observe that abnormal returns are correlated with the price-delay or market friction. We conclude that the stock-split post-announcement "drift" is only of short duration, and it is attributable to trading frictions rather than behavioral biases.

Keywords: stock splits, market efficiency, behavioral finance, long-run performance

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1. Introduction

The long-run performance of equities after stock splits is the subject of a vigorous academic debate between the behavioral finance and the efficient markets schools of thought. It is by now well accepted that stock splits signal favorable news about

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the fundamental value of a corporation, but if markets are semi-strongly efficient, the present value of such news should be fully priced during the narrow event window around the announcement date. An observed underreaction to such a simple corporate event, which leaves the corporation materially unchanged, calls into question the market's ability to quickly digest other more complex or ambiguous information.

The genesis of stock splits' significant role in the behavioral versus rational markets debate begins with Ikenberry, Rankine and Stice (1996) and Desai and Jain (1997). They report a positive price drift during the one-year period after the announcement of stock splits from 1975 to 1991 and from 1976 to 1992, respectively. These results seem so inconsistent with the semi-strong efficient markets paradigm that Daniel, Hirshleifer and Subrahmanyam (1998), in motivating their model of underreactions and overreactions based on psychological biases, cite stock splits (and these papers) as their first example of underreaction to public news events.[1]

Fama (1998) provides a vigorous defense of market efficiency and a critique of long-term return anomalies that purport to challenge the efficiency paradigm. One of Fama's arguments is that the reported anomalies are not sufficient to refute the efficient markets paradigm, because they have not been tested out of sample. Fama observes, "some anomalies do not stand up to out-of-sample replication. Foremost (in my mind) is the stock-split anomaly observed after 1975, which is contradicted by the earlier Fama, Fisher, Jensen and Roll (1969) study" (p. 304).

Seemingly in response to Fama's critique, two papers have emerged independently and almost simultaneously, Ikenberry and Ramnath (2002) and Byun and Rozeff (2003). Neither paper cites the other work, and both were accepted in different journals in early 2002. Ikenberry and Ramnath (2002) re-examine the stock-split anomaly over a long sample period, 1927–1997. They report significantly positive abnormal returns after stock splits throughout the sample period and generalize their findings by abstracting "these results are consistent with the notion of market underreaction to the information in corporate news events." Their results provide support for the behavioral theories of Daniel, Hirshleifer and Subrahmanyam (1998) and Barberis, Shleifer and Vishny (1998).

Commenting on Ikenberry and Ramnath (2002), Titman (2002) concurs that the study seems to "provide strong support for the overconfidence/underreaction hypothesis. Given the consistency of this evidence [their results] should probably tilt our beliefs toward some sort of overconfidence explanation" (p. 530). Nevertheless, Titman concedes puzzlement over what prevents people from trading on knowledge of the anomaly and making it disappear over time, and calls for further research on the matter.

Using almost the same sample period (1927–1996), Byun and Rozeff (2003) also study long-run performance after stock splits. They confirm the findings of

[1] Barberis, Shleifer and Vishny (1998), who provide another leading behavioral model, cite the post-announcement drift in Ikenberry, Rankine and Stice (1996) as an area where long-term overreaction may be building.

Ikenberry, Rankine and Stice (1996) and Desai and Jain (1997) in that they report long-run positive performance after two-for-one splits from 1975 to 1990. However, the magnitude of the upward drift is much smaller, 3.06% versus 7% or 8% reported in the other studies.

Here, Byun and Rozeff's (2003) only important point of accord with Ikenberry and Ramnath (2002) ends. Unlike Ikenberry and Ramnath, Byun and Rozeff find no robust evidence of out-of-sample post-split returns. They state (p. 1,066), "based on all the evidence in this paper, we conclude that investors have not systematically under-reacted (or over-reacted) to stock splits." They interpret their results as "new evidence that the stock market is efficient with respect to stock splits."

Thus, we find a debate over the existence, or absence, of post-split abnormal returns as a focal point of contention in an ongoing battle of competing price formation paradigms. Two simultaneously produced studies, investigating virtually the same event set, both relying upon state-of-the-art, long-run performance methodologies, arrive at conclusions that are in diametric opposition. Ikenberry and Ramnath (2002) report evidence in support of behavioral models of price formation; Byun and Rozeff (2003) argue that the evidence supports market efficiency. Clearly, a reconciliation of these conflicting findings is needed.

The purpose of this paper is to offer the needed reconciliation for the conflicting empirical findings and to offer an alternative rational explanation for the persistent long-run drift reported by Ikenberry and Ramnath (2002).

Our analysis demonstrates that security prices do, in fact, experience positive abnormal returns after stock-split announcements, but these abnormal returns are short-lived. Almost all of the abnormal returns attributable to stock-split announcements have been incorporated into the stock's price as of the ex-date. Post-ex-date abnormal returns are not substantially different from zero.

The empirical explanation for the differing results that Ikenberry and Ramnath (2002) and Byun and Rozeff (2003) obtain seems to rest on the differing horizon windows chosen for study. Ikenberry and Ramnath measure long-term returns beginning with the announcement date, whereas Byun and Rozeff begin their measurements after the effective split date.

That stock prices do not immediately and fully respond to management's split announcement, but that the information seems to be incorporated within the relatively brief period between the announcement and the split, is suggestive of the presence of some market friction that impairs the market's ability to fully price new information rapidly.

Hou and Moskowitz (2005) "parsimoniously characterize the severity of market frictions affecting a stock using the delay with which its price responds to information" (page 981). We calculate the Hou-Moskowitz market friction metric for splitting firms in our sample, and we find a strong correlation between the magnitude of price adjustments both at announcement and subsequently through the ex-date period.

The results in this paper are consistent with the post-announcement drift for splitting stocks being better explained by previously documented market frictions that impede the immediate incorporation of new information into the price of the security.

We note that although the abnormal returns that Ikenberry, Rankine and Stice (1996) and Ikenberry and Ramnath (2002) report are frequently referred to as "one-year post-split abnormal returns," they could more accurately be referred to as "one-year post-announcement abnormal returns," a semantic distinction that now proves to be of some importance.

2. Data and sample selection

We obtain the stock-split sample from the 2002 CRSP database. The primary focus of our paper is on the 12-month period after the announcement and ex (effective)-dates of stock splits. We use the calendar-time portfolio approach (e.g., Mitchell and Stafford, 2000) to examine the 612-month period from January 1950 to December 2000. To accomplish this task, we identify firms that announce two-for-one or greater stock splits during January 1949 to November 2000. We look for any noncash and nontaxable common share distribution of at least one share per existing common share. Both ordinary stock splits and large stock dividends (CRSP distribution codes 5,523 and 5,533, respectively) can qualify for the sample. For an event to qualify, the firm must meet the following additional criteria:

1. The common stock must trade on the NYSE, Amex, or Nasdaq and must be an ordinary common stock of a U.S. domiciled corporation (CRSP share codes 10 and 11). American Depository Receipts, Real Estate Investment Trusts and closed-end funds are not included in the sample.
2. The stock-split announcement date must appear in the CRSP database; data for the stock must start at least 24 months before the month of the split announcement.
3. The firm's ending stock price for the announcement month must be $2.00 per share or greater. This constraint is imposed to mitigate econometric biases induced by the bid-ask bounce of small price firms (Conrad and Kaul, 1993).

The sample contains 5,550 ordinary stock splits and 556 large stock dividends, for a total sample of 6,106 stock-split events. We assign NYSE market capitalization size decile rankings based on each sample firm's market capitalization for the month ending before the split announcement month. We calculate NYSE market capitalization size decile breakpoints monthly from the entire universe of U.S.-domiciled firms having common shares trading on the NYSE. We assign the sample firms, including Amex- and Nasdaq-listed ones, to size deciles based on the NYSE breakpoints. Descriptive statistics appear in Table 1.

Table 1

Descriptive statistics for stock-split sample

Qualifying stock splits must have a CRSP split factor of 1.0 (two-for-one stock split) or greater. Stock dividends of 100% (CRSP split factor of 1.0) or greater also are included. All NYSE-listed common stocks on CRSP are ranked by market capitalization each calendar month and assigned to ten size deciles, with decile ten containing the largest firms. All NYSE-, Amex- and Nasdaq-listed splitting firms are assigned to the appropriate NYSE decile based on market capitalization for the month ending before the stock-split announcement month.

Panel A: Frequency of stock splits by year of announcement

Year	Number	Year	Number	Year	Number
1950	33	1967	99	1984	138
1951	45	1968	139	1985	180
1952	28	1969	108	1986	333
1953	19	1970	35	1987	238
1954	43	1971	64	1988	67
1955	71	1972	97	1989	123
1956	79	1973	60	1990	94
1957	27	1974	22	1991	110
1958	18	1975	76	1992	178
1959	88	1976	114	1993	196
1960	47	1977	109	1994	142
1961	51	1978	148	1995	192
1962	52	1979	99	1996	227
1963	43	1980	236	1997	243
1964	80	1981	212	1998	231
1965	98	1982	100	1999	215
1966	93	1983	368	2000	198

Panel B: Stock-split sample categorized by NYSE size decile and exchange

NYSE decile	Number	Exchange	Number
1	813	NYSE	3,224
2	524	Amex	615
3	499	Nasdaq	2,267
4	562	Total	6,106
5	572		
6	572		
7	603		
8	583		
9	636		
10	742		
Total	6,106		

3. Methods

3.1. Calendar-time portfolios

We measure post-announcement abnormal returns primarily using the calendar-time method that Mitchell and Stafford (2000) propose, as Boehme and Sorescu

(2002) use in their paper concerning long-run returns after dividend initiations and resumptions. Mitchell and Stafford argue that the calendar-time approach is generally superior to the buy-and-hold abnormal return (BHAR) and cumulative abnormal return (CAR) approaches that earlier long-run performance studies use.

Contemporary long-run studies now commonly employ a four-factor model, specifically the three-factor Fama and French (1993) model augmented by a momentum factor introduced by Carhart (1997).[2] In this study, we first form calendar-time portfolios: for each calendar month, we calculate the monthly return to both equally weighted and value-weighted portfolios of firms announcing a stock split during the period $[t - h, t - 1]$, where t is the calendar month and h is equal to the length of the post-event investment horizon of interest, for example, 12 months.[3] To reduce idiosyncratic noise, we omit months containing fewer than ten firms. We estimate the four-factor regression:

$$R_{p,t} - R_{f,t} = \alpha_p + \beta_p(R_{m,t} - R_{f,t}) + s_p SMB_t + h_p HML_t + u_p UMD_t + e_{p,t}, \quad (1)$$

where $R_{p,t}$ represents the calendar-time portfolio of split-announcement firms, and $R_{f,t}$ is the return of one-month Treasury bills. The four independent variables are the excess return on the CRSP value-weighted market portfolio ($R_{m,t} - R_{f,t}$), the difference between returns of value-weighted portfolios of small and big firm stocks (SMB_t), the difference between returns of value-weighted portfolios of high and low book-to-market stocks (HML_t), and the difference between returns of portfolios of high and low prior year momentum stocks (UMD_t).[4]

[2] The explanatory power of the three-factor model is usually improved when augmented with the momentum factor. Many researchers argue that each of the three factors is a proxy for systematic risk, and the theoretical model of Johnson (2002) posits that momentum is a proxy for cash flow growth rate risk and represents a macroeconomic or systematic risk factor.

[3] Calendar-time portfolios are rebalanced each month to reflect the changing portfolio composition. Value-weighted returns use the prior month market capitalization as the weighting vector. Fama (1998) specifically calls attention to value weighting because several previous studies that report significant abnormal returns by using equally weighted portfolios are shown to lack robustness when reexamined with value-weighted methods. Fama posits that small firms are the most susceptible to the misspecified model problem. Therefore, if the sample is overpopulated with small firms, empirical results based on equally weighted portfolios are more likely to be driven by the misspecified model problem. Value weighting mitigates the misspecified model problem by giving a higher weight to the larger firms, for which the problem is likely to be less severe. Value weighting also is potentially more representative of the aggregate wealth generating aspects of any presumed anomaly.

[4] We thank Kenneth French for providing the three Fama and French (1993) factors and the one-month T-Bill returns. We calculate the momentum factor (UMD_t) in a procedure similar to Carhart (1997) as follows: for each calendar month t, we rank all NYSE, Amex and Nasdaq stocks based on their holding period return (prior year "momentum") for months $t - 12$ to $t - 2$. We calculate the returns to a zero investment portfolio UMD for calendar month t, that is, long and short on portfolios comprised firms in the top 30% and bottom 30% momentum categories, respectively. We calculate both equally and value-weighted portfolio versions of this momentum factor.

3.2. Size- and momentum-matched calendar-time regressions

Stock splits announcement generally follow a period of strongly positive stock price performance or momentum.[5] Failing to control for this pre-event momentum can lead to misleading inferences concerning long-run abnormal returns. For example, Lyon, Barber, and Tsai (1999) find that firms with high pre-event momentum yield positively biased *t*-statistics in random samples over one-year horizons. Fama and French (1996) also report a momentum bias for their three-factor model. In addition to the momentum bias, Fama and French (1996) and Mitchell and Stafford (2000) show that the traditional Fama and French (1993) three-factor model does not completely explain the cross-section of stock returns. When the three-factor model is estimated in randomly chosen samples of small firms that have low book-to-market ratios, the null hypothesis of zero abnormal performance is over-rejected.

Similar to Mitchell and Stafford (2000), we control for potential biases or model misspecifications by estimating a matched four-factor model. We construct a hedge (zero-investment) calendar-time portfolio consisting of long positions on splitting firms and short positions on nonsplitting control firms.[6] We regress the returns of this hedge portfolio on the four-factor model:

$$R_{p,t} - R_{c,t} = \alpha_p + \beta_p(R_{m,t} - R_{f,t}) + s_p SMB_t + h_p HML_t + u_p UMD_t + e_{p,t}. \quad (2)$$

The "matched" intercept (α_p) we obtain in this manner represents a measure of long-run abnormal performance that specifically corrects for the size bias that is inherent in the traditional three-factor model and any remaining momentum bias that the Carhart momentum factor captures insufficiently. We report both the unmatched (traditional) and matched intercepts.[7]

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[5] For example, calendar-time pre-announcement abnormal returns (not reported in a table) are 2.15% ($t =$ 21.09) and 3.91% ($t = 21.52$) per month, for equally weighted portfolios during 1950–1974 and 1975–2000, respectively.

[6] To match control firms to each sample firm, we select as candidates all firms that have market value of equity as of month $m - 1$ (m is the split announcement month) between 60% and 140% of that of the sample firm and did not announce a two-for-one or greater stock split in the prior 12 months. We identify the three candidates that have one-year pre-announcement performance closest to the sample firm, where pre-announcement performance is the holding period return over $[m - 12, m - 1]$. The candidate closest in pre-announcement performance is designated the primary control firm; the remaining two become the secondary and tertiary control firms. If the primary control is delisted before the end of the horizon under investigation, then the secondary control match is substituted for the remainder of the horizon; if the secondary control is delisted, then the tertiary is used. As in the unmatched regression, months containing fewer than ten sample firms are omitted.

[7] The matched portfolio procedure differs from that shown in Equation (6) of Mitchell and Stafford (2000) but conforms to Boehme and Sorescu (2002). The procedure of Boehme and Sorescu produces an intercept identical to Mitchell and Stafford's, but the *t*-statistic of the intercept accounts for the covariance between the event firm and control firm calendar-time returns.

3.3. Industry momentum-matched calendar-time regressions

Unexpectedly good (bad) economic times within an industry obviously coincide with positive (negative) stock price momentum. Moskowitz and Grinblatt (1999) focus on the role of industry momentum in explaining individual firm returns. They observe that industry momentum has greater explanatory power for subsequent returns than firm-specific momentum and conclude that "industry momentum may be the key element in understanding return persistence anomalies" (p. 1,287). Against this backdrop, we also conduct industry momentum-matched regressions. Specifically, we use a hedge portfolio approach made up of long positions in each split stock, and corresponding short positions in industry-matched control stocks that do not split within the prior 12 months. We select the control stock using the first two digits of the sample firm's (Standard Industry Classification) (SIC) code, and because many SIC are notoriously indiscriminant in their inclusion criteria, we attempt to select a firm within the SIC that closely resembles the sample firm's true "industry" by choosing the SIC-matched firm that most closely matches the sample firm's pre-split momentum.[8]

4. Long-run post-announcement versus post-split abnormal returns

4.1. Post-announcement abnormal returns

To preserve comparison with other studies, we report, in Table 2, post-announcement abnormal returns from calendar-time regressions for the entire 1950–2000 period, and for 1950–1974, 1975–1987, and 1988–2000. We report the results for each of the first three post-announcement years in separate panels.

The pattern of abnormal returns for the full 1950–2000 period is generally consistent with the findings of Ikenberry and Ramnath (2002). For the one-year post-announcement horizon in Panel A, all equally weighted intercepts are positive and statistically significant at the 5% level or better. Although the value-weighted results are less robust across subperiods, there is evidence of positive abnormal performance over the full 1950–2000 period. The magnitude and statistical significance of the coefficient estimates are greatest for 1975–1987. This subperiod most closely corresponds to the study periods of Ikenberry, Rankine and Stice (1996) and Desai and Jain (1997).[9] Compounding Panel A equally weighted matched intercept for 1975–1987 over 12 months yields an annual abnormal return of 7.33%, roughly equivalent to the one-year post-announcement BHARs of 7.94% and 7.05% in Ikenberry, Rankine and Stice and Desai and Jain, respectively.

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[8] Imposing size or market capitalization bounds for size matching often fails to select a control firm within the size bounds, due to the often limited number of firms in a given industry.

[9] Ikenberry, Rankine and Stice (1996) analyze splits (only two-for-one splits) announced between 1975 and 1991, whereas Desai and Jain (1997) study splits (1.25-for-1 or greater splits) announced from 1976 to 1992.

Table 2

Long-run abnormal returns following stock-split announcements using calendar-time regressions

Regression intercepts representing estimated monthly abnormal return. Equal- and value-weighted (monthly rebalanced) calendar-time portfolio returns (in decimal, not percent) are calculated each month from sample firms that announce two-for-one or greater stock splits in the previous one to 12, 13–24, or 25–36 calendar months. Months containing fewer than ten firms are omitted. For the unmatched regression, monthly excess returns to the calendar-time portfolios, $R_{p,t} - R_{f,t}$, are regressed on the four-factor model:

$$R_{p,t} - R_{f,t} = \alpha_p + \beta_p(R_{m,t} - R_{f,t}) + s_p SMB_t + h_p HML_t + u_p UMD_t + e_{p,t},$$

where, $R_{f,t}$ is the return of one-month T-Bills. $(R_{m,t} - R_{f,t})$ is the excess return of the CRSP value-weighted market index. SMB_t is the difference in returns between value-weighted portfolios of small and big firm stocks. HML_t is the difference in returns between value-weighted portfolios of high and low book-to-market ratio stocks. UMD_t is the difference in returns between portfolios of high and low prior year momentum stocks (equal- and value-weighted for equal- and value-weighted regressions, respectively). For the matched regressions, two control firm-matching methods are used: (1) pre-announcement size and momentum (matched on both pre-announcement size and momentum for the month before the month containing the announcement date) and (2) industry and momentum (from among the firms in each sample firm's industry (first two digits of SIC code), the firm closest in pre-announcement momentum is chosen as the control form). Control portfolios are subtracted from the event portfolios, and the difference $R_{p,t} - R_{c,t}$ regressed on the four factors as shown above: Ordinary and weighted least squares (OLS and WLS) time-series regressions are estimated. Months in the WLS model are weighted by the square root of the number of firms contained in the month. Heteroskedasticity-consistent t-statistics (White, 1980) are in brackets.

	Jan. 1950–Dec. 1974 ($N = 300$ months)		Jan. 1975–Dec. 1987 ($N = 156$ months)		Jan. 1988–Dec. 2000 ($N = 156$ months)		Jan. 1950–Dec. 2000 ($N = 612$ months)	
Calendar-time model	OLS	WLS	OLS	WLS	OLS	WLS	OLS	WLS
Panel A: Year 1 of post-announcement horizon (months 1–12)								
Equally weighted,	0.204	0.170	0.664	0.566	0.454	0.353	0.396	0.390
unmatched	[2.23]**	[2.05]**	[4.73]***	[4.14]***	[3.66]***	[2.69]***	[5.82]***	[5.27]***
Equally weighted,	0.312	0.330	0.591	0.554	0.340	0.251	0.375	0.360
size/momentum matched	[3.17]***	[3.31]***	[4.62]***	[4.35]***	[2.95]***	[2.05]**	[5.65]***	[5.14]***
Equally weighted,	0.230	0.291	0.496	0.485	0.276	0.232	0.303	0.299
industry matched	[2.18]**	[2.86]***	[3.54]***	[3.62]***	[1.43]	[1.14]	[3.66]***	[3.04]***
Value weighted,	0.121	0.033	0.328	0.420	0.264	0.234	0.187	0.218
unmatched	[1.10]	[0.30]	[1.87]*	[2.57]**	[1.99]**	[1.47]	[2.40]**	[2.46]**
Value weighted,	0.197	0.236	0.366	0.556	0.089	−0.055	0.200	0.223
size/momentum matched	[1.39]	[1.69]*	[1.61]	[3.32]***	[0.58]	[−0.30]	[2.07]**	[2.25]**
Value weighted,	0.110	0.104	0.618	0.625	0.178	0.133	0.245	0.248
industry matched	[0.72]	[0.70]	[2.84]***	[2.87]***	[0.66]	[0.32]	[2.10]**	[1.54]

(*continued*)

Table 2 (*continued*)

Long-run abnormal returns following stock-split announcements using calendar-time regressions

	Jan. 1950–Dec. 1974 ($N = 300$ months)		Jan. 1975–Dec. 1987 ($N = 156$ months)		Jan. 1988–Dec. 2000 ($N = 156$ months)		Jan. 1950–Dec. 2000 ($N = 612$ months)	
Calendar-time model	OLS	WLS	OLS	WLS	OLS	WLS	OLS	WLS
Panel B: Year 2 of post-announcement horizon (months 13–24)								
Equally weighted,	0.073	0.109	0.013	0.020	−0.146	−0.130	0.013	−0.007
unmatched	[0.89]	[1.20]	[0.11]	[0.14]	[−0.94]	[−0.76]	[0.17]	[−0.10]
Equally weighted,	0.028	0.060	−0.083	−0.188	−0.010	0.074	−0.016	−0.038
size/momentum matched	[0.30]	[0.61]	[−0.62]	[−1.44]	[−0.08]	[0.61]	[−0.24]	[−0.56]
Equally weighted,	0.109	0.172	−0.065	−0.064	−0.054	0.085	0.012	0.042
industry matched	[1.17]	[1.63]	[−0.40]	[−0.53]	[−0.26]	[0.36]	[0.15]	[0.41]
Value weighted,	0.116	0.236	0.157	0.115	−0.097	−0.127	0.057	0.031
unmatched	[1.13]	[2.06]**	[0.88]	[0.63]	[−0.66]	[−0.77]	[0.72]	[0.32]
Value weighted,	0.117	0.220	0.396	0.098	−0.177	−0.201	0.120	0.014
size/momentum matched	[0.91]	[1.63]	[1.96]**	[0.64]	[−0.92]	[−0.83]	[1.28]	[0.14]
Value weighted,	0.209	0.399	−0.209	−0.155	0.164	0.085	0.065	0.071
industry matched	[1.41]	[2.48]**	[−0.92]	[−0.64]	[0.79]	[0.36]	[0.57]	[0.53]
Panel C: Year 3 of post-announcement horizon (months 25–36)								
Equally weighted,	0.038	0.030	0.218	0.180	0.133	0.137	0.108	0.100
unmatched	[0.44]	[0.33]	[1.64]	[1.17]	[0.99]	[0.98]	[1.59]	[1.14]
Equally weighted,	0.013	0.005	0.063	0.003	0.206	0.175	0.063	0.053
size/momentum matched	[0.13]	[0.05]	[0.55]	[0.03]	[1.64]	[1.39]	[0.91]	[0.77]
Equally weighted,	−0.005	0.017	0.221	0.057	−0.162	−0.246	0.007	−0.094
industry matched	[−0.05]	[0.17]	[1.62]	[0.39]	[−0.91]	[−1.40]	[0.09]	[−0.97]
Value weighted,	−0.107	−0.134	−0.019	−0.068	0.220	0.261	−0.018	0.023
unmatched	[−1.05]	[−1.37]	[−0.14]	[−0.44]	[1.60]	[1.84]*	[−0.24]	[0.28]
Value weighted,	−0.127	−0.194	−0.020	−0.132	0.375	0.406	0.005	0.040
size/momentum matched	[−0.84]	[−1.34]	[−0.11]	[−0.71]	[2.03]**	[2.11]**	[0.05]	[0.37]
Value weighted,	0.021	0.018	0.100	−0.061	−0.137	−0.196	−0.005	−0.088
industry matched	[0.14]	[0.13]	[0.53]	[−0.30]	[−0.56]	[−0.72]	[−0.04]	[−0.67]

***, **, * indicate statistical significance at the 0.01, 0.05 and 0.10 level, respectively.

For the second and third post-announcement years in Panels B and C, in sharp contrast to Panel A, we find no evidence of abnormal performance, suggesting that the presumed stock-split anomaly is confined to (at most) the first post-announcement year. This result is consistent with Ikenberry and Ramnath (2002), who likewise observe that the anomaly is generally confined to the first post-announcement year. Thus, it would seem that the stock-split anomaly is relatively short-lived vis-à-vis anomalies that other researchers typically report over longer horizons of three to five years. This also suggests that the stock-split anomaly is unlikely to result from a potential misspecification of the four-factor model, because the calendar portfolios in the second and third post-announcement years do not evidence such misspecification.

For purposes of robustness, we repeat the analysis in Table 2 using equally weighted BHARs and CARs. The results are in Table 3.[10] Once again, we find that the positive abnormal returns are completely isolated to the first year; second-year BHARs and CARs are actually negative and statistically significant during 1975–1987.

Summarizing the results from Tables 2 and 3, there seems to be strong evidence of abnormal performance in the first year after announcements of stock splits. The abnormal performance is strongest during 1975–1987, but unlike findings regarding other well-publicized anomalies, the abnormal performance does not exceed one year and potentially could be confined to a much narrower horizon.

The duration of the abnormal performance after stock splits is interesting, in part, because other well-known anomalies in the financial literature are typically reported over longer horizons. For example, the anomalies (long-run abnormal returns) after dividend initiations and omissions, initial public offerings and seasoned equity offerings are reported over horizons typically lasting from three to five years.[11] Although some researchers propose rational explanations for most anomalies after other type of corporate events, the post-announcement stock-split anomaly does not seem to emanate from misspecification in the four-factor calendar-time regression model (or even the BHAR and CAR models) because we already reconcile the second and third post-announcement years to each of the asset pricing models we use.

We are left to conclude that the first-year post-announcement abnormal performance is not an artifact of pricing model misspecification. Although one plausible

[10] We calculate equally weighted BHARs and CARs following the method of Barber and Lyon (1997). We estimate the long-run raw return for each event firm (cumulative or buy-and-hold), from which we subtract the corresponding raw return of a control firm matched on size and one-year momentum (we construct the match using the same algorithm employed in the adjusted four-factor calendar-time procedure). We compute the cross-sectional averages (equally weighted) of the return differences and report the results. The BHAR t-statistics are adjusted for skewness using the bootstrapping procedure described by Lyon, Barber and Tsai (1999).

[11] Long-run abnormal returns or anomalies of three- to five-year duration are reported after initial public offerings or IPOs (Ritter, 1991), mergers (Agrawal, Jaffe and Mandelker, 1992), dividend initiations and omissions (Michaely, Thaler and Womack, 1995), stock repurchases (Ikenberry, Lakonishok and Vermaelen, 1995), new exchange listings (Dharan and Ikenberry, 1995), seasoned equity offerings or SEOs (Loughran and Ritter, 1995) and convertible-debt issuance (Spiess and Affleck-Graves, 1999).

Table 3

Buy-and-hold and cumulative abnormal returns (BHARs and CARs) following stock-split announcements

Percentage equally weighted mean BHARs and mean CARs following stock-split announcements, calculated using the Barber and Lyon (1997) procedures. BHARs are calculated by subtracting the buy-and-hold return of the control firm from that of the splitting firm. CARs are calculated by subtracting the sum of monthly returns of a control firm from that of the splitting firm. Two matching methods (same control firms as Table 2) are used: pre-announcement size and momentum (Panel A) and industry and pre-announcement momentum (Panel B). *t*-statistics are in brackets. For BHARs, *t*-statistics are skewness adjusted, and significance levels are computed using the bootstrapping procedure in Lyon, Barber and Tsai (1999).

	Jan. 1950–Dec. 1974		Jan. 1975–Dec. 1987		Jan. 1988–Dec. 2000		Jan. 1950–Dec. 2000	
	BHAR	CAR	BHAR	CAR	BHAR	CAR	BHAR	CAR
Panel A: Using control firms matched on pre-announcement size and momentum								
Year 1	3.816%	3.776%	4.276%	4.961%	6.558%	5.894%	4.937%	4.971%
(months 1–12)	[3.38]***	[3.75]***	[3.61]***	[4.87]***	[2.96]***	[3.82]***	[5.29]***	[6.95]***
Number of observations	1,533	1,533	2,320	2,320	2,008	2,008	5,861	5,861
Year 2	0.159	0.194	−2.743	−2.503	2.290	1.616	−0.349	−0.456
(months 13–24)	[−0.14]	[0.20]	[−2.20]**	[−2.36]**	[1.28]	[1.76]*	[−0.42]	[−0.65]
Number of observations	1,519	1,519	2,255	2,255	1,752	1,752	5,526	5,526
Year 3	−1.791	−0.918	−0.064	−0.380	0.186	0.960	−0.496	−0.157
(months 25–36)	[−1.44]	[−0.89]	[−0.05]	[−0.35]	[0.09]	[0.57]	[−0.56]	[−0.21]
Number of observations	1,488	1,488	2,168	2,168	1,447	1,447	5,103	5,103
Panel B: Using control firms matched on industry and momentum								
Year 1	2.608%	3.047%	2.704%	3.564%	6.835%	5.236%	4.229%	4.047%
(months 1–12)	[2.21]**	[2.99]***	[1.81]*	[3.12]***	[3.04]***	[3.17]***	[4.06]***	[5.14]***
Number of observations	1,526	1,526	1,861	1,861	2,039	2,039	5,426	5,426
Year 2	0.229	0.352	−1.673	−2.050	0.580	0.279	−0.317	−0.519
(months 13–24)	[0.19]	[0.35]	[−1.20]	[−1.75]*	[0.19]	[0.16]	[−0.31]	[−0.65]
Number of observations	1,512	1,512	1,802	1,802	1,804	1,804	5,118	5,118
Year 3	−0.893	−0.961	−0.700	−0.771	0.258	−2.764	−0.455	−1.467
(months 25–36)	[−0.71]	[−0.95]	[−0.50]	[−0.65]	[0.12]	[−1.59]	[−0.48]	[−1.90]*
Number of observations	1,482	1,482	1,739	1,739	1,510	1,510	4,731	4,731

***, **, * indicate statistical significance at the 0.01, 0.05 and 0.10 level, respectively.

explanation for this first-year drift is that investors underreact to the split announcement, we next consider the possibility that the drift might arise from market frictions that impede the market's ability to reprice shares immediately upon announcement of an impending split.

4.2. Post-split abnormal returns

The alternative to the behavioral long-run anomaly explanation for post-announcement price drift is that market frictions impair the speed with which new information is incorporated into the securities' post-announcement prices.

We begin our exploration of a potential market friction-based explanation of the post-announcement drift by measuring the long-term abnormal returns for the period that begins immediately after the split date as opposed to the announcement date. We form calendar-time portfolios beginning with the month succeeding the split date. On average, there are 28 trading days between the announcement and ex-date in our sample. Hou and Moskowitz (2005) argue that for most stocks, we should expect that market frictions are resolved in less than one month.[12] Thus, these portfolios should be largely free of market friction-based contaminations. If the stock-split anomaly is actually a behavioral long-run anomaly, it should still exist if we only postpone the calendar-time portfolio inclusion period by approximately a month and a half after the announcement date, that is, until after the stock split effectively occurs.

Table 4 presents one-year calendar-time post-split results.[13] The once-robust post-announcement results in Panel A of Table 2 almost completely disappear in Table 4. The value-weighted abnormal returns are positive and statistically significant at the 10% level in only one of the 24 regressions. This finding strongly suggests that there is no long-term post-split anomaly on a value-weighted basis. Moreover, even on an equal-weighted basis, the results are significantly weakened in that the abnormal returns are zero except during 1975–1987.

For robustness, we compute one-year post-split BHARs and CARs and report the results in Table 5. The BHARs and CARs essentially reject the presence of a long-run post-split drift. None of the subperiod measures are significant at standard levels, not even 1975–1987. We conduct other robustness tests that are not tabulated. Most importantly, we replicate this section's various tests using daily data. Ikenberry and Ramnath (2002) conduct their tests beginning two days after the announcement date using daily data. Our tests of the 250-day post-announcement CARs and BHARs are entirely consistent with those that Ikenberry and Ramnath observe. However, when we begin cumulating returns two days after the split, the 250-day post-split CARs and BHARs are much smaller. Moreover, during 1988–2000, none of the post-split CARS or BHARs is significant at the 5% level, and the point estimate for industry

[12] Hou and Moskowitz (2005) state (p. 984), "At monthly frequencies, there is little dispersion in delay measures since most stocks respond to information within a month's time."

[13] The control firm algorithm of Section 3 is modified to select control firms for the month before the month containing the ex or effective split date, as opposed to the month before the announcement month.

Table 4

Calendar-time regressions of portfolios formed subsequent to the effective or split dates during 1950–2000

Regression intercepts representing estimated monthly abnormal return. Equal- and value-weighted calendar-time portfolio returns are calculated each month for sample firms having ex-dates of two-for-one or greater stock splits in the prior 12 months. Months containing fewer than ten firms are omitted. For the unmatched four-factor regression model, the dependent variable is the monthly excess return of the calendar-time portfolios, $R_{p,t} - R_{f,t}$, ($R_{f,t}$ is the return of the one-month T-bill); the independent variables are (1) ($R_{m,t} - R_{f,t}$), the excess return of the CRSP value-weighted market index; (2) SMB_t, the difference in returns between value-weighted portfolios of small and big stocks; (3) HML_t, the difference in returns between value weighted portfolios of high and low book-to-market ratio stocks; and (4) UMD_t, the difference in returns between portfolios of high and low prior year momentum stocks (equal- and value-weighted for equal- and value-weighted regressions, respectively). For the matched regressions, control firms are selected two ways: (1) pre-announcement size and momentum (matched on both pre-split size and momentum for the month before the month containing the ex-date) and (2) industry and momentum (from among the firms in each sample firm's industry [first two digits of SIC code], the firm closest in pre-announcement momentum is chosen as the control form). Ordinary and weighted least squares (OLS and WLS) time-series regressions are estimated. For WLS, months are weighted by the square root of the number of firms contained in the month. Heteroskedasticity-consistent t-statistics (White, 1980) are in brackets.

	Jan. 1950–Dec. 1974 (N = 300 months)		Jan. 1975–Dec. 1987 (N = 156 months)		Jan. 1988–Dec. 2000 (N = 156 months)		Jan. 1950–Dec. 2000 (N = 612 months)	
Calendar-time model estimated	OLS	WLS	OLS	WLS	OLS	WLS	OLS	WLS
Equally weighted,	0.165	0.100	0.359	0.290	0.042	−0.027	0.197	0.129
unmatched	[1.77]*	[1.06]	[2.46]**	[2.04]**	[0.35]	[−0.18]	[2.85]***	[1.67]*
Equally weighted,	0.130	0.103	0.238	0.266	0.098	−0.025	0.149	0.099
size/momentum matched	[1.40]	[1.05]	[1.69]*	[1.97]**	[0.83]	[−0.21]	[2.29]**	[1.42]
Equally weighted,	0.082	0.111	0.105	0.127	−0.087	−0.081	0.043	0.014
industry matched	[0.87]	[1.15]	[0.63]	[0.76]	[−0.44]	[−0.37]	[0.51]	[0.14]
Value weighted,	0.076	0.051	0.196	0.324	0.072	0.083	0.075	0.121
unmatched	[0.69]	[0.42]	[1.02]	[1.87]*	[0.54]	[0.54]	[0.92]	[1.33]
Value weighted,	0.094	0.103	0.151	0.210	0.124	0.095	0.094	0.120
size/momentum matched	[0.71]	[0.73]	[0.79]	[1.21]	[0.71]	[0.49]	[1.01]	[1.17]
Value weighted,	0.016	0.108	0.349	0.287	−0.193	−0.250	0.038	0.023
industry matched	[0.10]	[0.69]	[1.59]	[1.20]	[−0.87]	[−0.98]	[0.35]	[0.17]

***, **, * indicate statistical significance at the 0.01, 0.05 and 0.10 level, respectively.

Table 5

One-year buy-and-hold and cumulative post-effective split abnormal returns (BHARs and CARs)

Percentage equally weighted mean BHARs and mean CARs for the 12 months after the ex-date of stock splits, calculated using the procedure of Barber and Lyon (1997). The BHARs are calculated by subtracting the buy-and-hold return of the control firm from the buy-and-hold of the splitting firm. The CARs are calculated by subtracting the sum of monthly returns of a control firm from that of the splitting firm. Two matching methods (same control firms as Table 2) are used: pre-split size and momentum (Panel A) and industry and pre-split momentum (Panel B). *t*-statistics are in brackets. For BHARs, *t*-statistics are skewness adjusted, and significance levels are computed using the bootstrapping procedure in Lyon, Barber and Tsai (1999).

	Jan. 1950–Dec. 1974		Jan. 1975–Dec. 1987		Jan. 1988–Dec. 2000		Jan. 1950–Dec. 2000	
	BHAR	CAR	BHAR	CAR	BHAR	CAR	BHAR	CAR
Size and momentum	0.175%	0.736%	0.467%	0.707%	2.182%	2.002%	0.974%	1.155%
matched	[0.15]	[0.73]	[0.40]	[0.67]	[1.09]	[1.37]	[1.11]	[1.65]*
Number of observations	1,549	1,549	2,326	2,326	1,999	1,999	5,874	5,874
Industry and	0.811	1.096	−1.636	−0.266	0.136	2.000	−0.295	0.928
momentum matched	[0.72]	[1.08]	[−1.17]	[−0.23]	[0.06]	[1.22]	[−0.29]	[1.21]
Number of observations	1,543	1,543	1,864	1,864	1,837	1,837	5,244	5,244

* indicate statistical significance at the 0.10 level.

momentum-matched post-split CARs is actually negative when the Ikenberry and Ramnath (2002) cumulation technique is applied after the split date.

5. Abnormal returns between announcement and payment dates

Because of the result in Section 4 that one-year post-split returns, while positive, generally are not statistically different from zero, one can infer that abnormal returns are confined to the period between the announcement and the ex-date. In this section, we focus specifically on this short interval.

To understand the possible relationship between market frictions and post-announcement price drifts, let us review the process by which a stock split occurs. On the announcement date, the firm will announce a record date and a payment date for the split. By convention, the ex-date is the trading day that follows the payment date. On the ex-date and thereafter, trading commences in the split shares. Several weeks elapse between the announcement date and the payment date. For expositional convenience, we refer to the ex-date, when trading commences in the split shares, as the "split date."

To capture the magnitude of the potential short-term market frictions, Table 6 presents a breakdown of the abnormal returns on and around the announcement and

Table 6

Cumulative abnormal returns associated with stock-split announcement and effective (ex) dates

Mean percentage cumulative abnormal returns (CARs) for (1) the three-day stock-split announcement date window, (2) the interval between the announcement and ex-dates, (3) the three-day effective split or ex-date window and (4) the period from days 2–10 after the ex-date. Each three-day window is centered on its respective event date. The method of Brown and Warner (1985) is used to calculate the short-term CARs and significance levels. The CARs are estimated by summing the daily returns of the event firm for the window of interest and subtracting corresponding return to the CRSP equally weighted market index (EWRETD index).

Event window	July 1962– Dec. 1974	Jan. 1975– Dec. 1987	Jan. 1988– Dec. 2000	July 1962– Dec. 2000
Announcement period CAR	2.766	3.781	3.107	3.357
$[t_{annc} - 1$ to $t_{annc} + 1]$	[19.01]***	[28.51]***	[19.02]***	[36.96]***
CAR from $[t_{annc} + 2$ to $t_{ex} - 2]$,	0.692	0.705	1.831	1.153
from annc to ex-period	[1.58]	[3.21]***	[5.87]***	[6.64]***
Ex-date period	0.734	2.262	1.708	1.778
CAR $[t_{ex} - 1$ to $t_{ex} + 1]$	[4.76]***	[17.01]***	[10.17]***	[19.28]***
Post ex-date period	0.896	0.830	0.107	0.552
CAR $[t_{ex} + 2$ to $t_{ex} + 10]$	[3.64]***	[4.62]***	[0.49]	[4.44]***
Median market days from	54	27	24	28
announcement to ex-date	N = 950	N = 2,370	N = 2,217	N = 5,537

*** indicate statistical significance at the 0.01 level.

split date windows. Table 6 reports the CARs for the three-day announcement- and split-date windows as well as the abnormal return for the entire interval between announcement and split dates, and for the ten-day interval after the split date.

As numerous other studies report, the average market reaction to stock-split announcements (Table 6) is significantly positive. This further confirms that stock splits are interpreted as management signals of favorable private information concerning the firm's value, as suggested by Fama, Fisher, Jensen and Roll (1969), Grinblatt, Masulis and Titman (1984) and others. However, we observe significantly positive split-date CARs across all periods. The split-date CARs are a component of the long-run positive post-announcement abnormal returns in Tables 2 and 3, because the median number of trading days between the announcement date and the subsequent split date is 28. Not surprisingly, the subperiod reporting the largest matched calendar-time intercepts in Table 2, 1975–1987, is also the one containing the largest mean split date CAR (2.262%, $t = 17.01$).

We also examine a ten-day window after the split date to determine whether positive abnormal returns persist after the ex-date. The ten-day post-split CAR is 0.896% ($t = 3.64$) and 0.830% ($t = 4.62$) for 1962–1974 and 1975–1987, respectively. However, post-split abnormal returns are no longer evident during 1988–2000.

The results in Table 6 suggest that a substantial portion, if not all, of the post-announcement abnormal returns in Table 2 and in previous studies are attributable to the short-term price adjustments that occur shortly after the announcement date and end near the split date. As an example, for 1975–1987, summing the CARs from (1) the period between the announcement and split dates, (2) the split-date window and (3) the ten-day post-split-date period, yields a total CAR of 3.797%. For comparison, the total one-year post-announcement CAR in Table 3 is 4.961% during 1975–1987. Thus, most of the one-year post-announcement CAR occurs in the relatively short period between the announcement date and a few days after the split date.

6. Correlations between market friction price delays and short-term abnormal returns

In the previous sections, we examine the timing of abnormal returns related to stock-split announcements and payments. These abnormal returns are primarily confined to the period between the announcement date and the ex-date, and there is no evidence of any multiyear drift. This price drift pattern is not consistent with a behavioral-based underreaction explanation, but it is more consistent with the possibility of delayed price response to new information due to market frictions.

In this section, we use an explicit, parsimonious measure of the severity of market frictions faced by individual splitting firms. Controlling for firm-specific market frictions, we compare the return characteristics of high-friction firms to that of low-friction firms. These test results are supportive of the hypothesis that post-announcement price drifts are related to market frictions that prevent information from being efficiently priced at the announcement.

We use the market friction measure of Hou and Moskowitz (2005). They use weekly returns for the CRSP value-weighted market index (VWRETD) and for individual firms. We calculate the needed Thursday-to-Wednesday weekly returns from compounded CRSP daily returns. As in Equation (1) of Hou and Moskowitz, we regress each firm's weekly returns on the contemporaneous market index weekly return and four lagged market weekly returns as follows:

$$R_{i,t} = \alpha_i + \beta_i R_{m,t} + \delta_{i,t-1} R_{m,t-1} + \delta_{i,t-2} R_{m,t-2} + \delta_{i,t-3} R_{m,t-3} + \delta_{i,t-4} R_{m,t-4} + \varepsilon_{i,t}. \quad (3)$$

We estimate a second, constrained regression, restricting $\delta_{i,t-1}$ through $\delta_{i,t-4}$ to be zero. We construct the market friction measure, *delay,* as in Equation (2) of Hou and Moskowitz (2005), by dividing the R^2 of the restricted model by the R^2 of the full model and subtracting the result from one.

$$\text{Delay} = 1 - [R^2(\text{restricted model})/R^2(\text{full model})]. \quad (4)$$

The larger the delay, the more return variation is captured by the lagged returns. Thus, high values of delay indicate that there is a strong delay response in return innovations. Simply, Hou and Moskowitz (2005) observe that firms with high delay values respond more slowly to new information.

Table 7 shows the CARs for each delay-decile group. Decile 1 contains the firms with the lowest delay values, and decile ten, the highest. We calculate CARs for several intervals: (1) the three-day announcement-date window, (2) the interval between the announcement date and the ex-date, (3) the three-day ex-date window and (4) a post-split period beginning two days after the announcement and ending ten days after the ex-date. The fourth interval subsumes the second and third interval and captures returns for several days following the ex-date.

Table 7, Panel A, presents the results over the full sample period. The CARs are increasing in delay, with decile ten reporting the largest announcement period CARs. The difference of means between delay deciles ten and one is statistically significant (t = 5.85). The results suggest that stocks experiencing greater ex ante price delay exhibit a greater surprise response to split announcements than do stocks that more efficiently incorporate new information. High delay stocks respond as though investors are more surprised by management's positive news.

Examining the CARs between the announcement and the payment date and those around the ex-date, we see that the most price-delayed stocks also exhibit the largest abnormal returns after the announcement. We interpret these results to mean that although the market is more surprised by the split announcement for high-delay firms, such firms still remain more price-delayed than other firms. Thus, although the split announcement helps to correct high-delay firms' market values to reflect positive news, the firms are nevertheless relatively sluggish in incorporating the information implied by the split announcement.

Panels B through D of Table 7 provide subperiod results. The results are generally consistent with the full-sample results, although during 1963–1974, when fewer splits occur, no relation between delay deciles and post-announcement returns is evident.

Table 7

Cumulative abnormal returns around stock-split announcement and ex-dates, sorted by price delay

Weekly returns (compounding daily returns from Thursday through Wednesday) are generated for the CRSP value-weighted market index (VWRETD) and for sample firms announcing stock splits. As in Equation (1) of Hou and Moskowitz (2005), for each sample firm we estimate an OLS regression of the firm's 52 weekly returns immediately before the announcement week by regressing the firm's pre-announcement weekly returns on the contemporaneous VWRETD weekly return and four lagged weekly returns of the VWRETD.

$$R_{i,t} = \alpha_i + \beta_i R_{m,t} + \delta_{i,t-1} R_{m,t-1} + \delta_{i,t-2} R_{m,t-2} + \delta_{i,t-3} R_{m,t-3} + \delta_{i,t-4} R_{m,t-4} + \varepsilon_{i,t}.$$

A second regression is estimated that restricts $\delta_{i,t-1}$ through $\delta_{i,t-4}$ to zero. The price delay measure (delay) is constructed as in Equation (2) of Hou and Moskowitz, by dividing the R^2 of the restricted model by the R^2 of the full model and subtracting the result from 1.0.

$$\text{Delay} = 1 - [R^2 \text{ (restricted model)}/R^2 \text{ (full model)}]\,.$$

The mean CARs (reported in percentage) are presented for (1) the three-day stock-split announcement date window, (2) the period or interval between the announcement and ex-dates, (3) the three-day effective split or ex-date window and (4) the period beginning two days after the announcement and ending ten days after the ex-date. The t-statistic of the difference in means between delay deciles ten and one is in brackets.

		Mean cumulative abnormal return			
Delay decile	Number of firms	annc − 1 to annc + 1	annc + 2 to ex − 2	ex − 1 to ex + 1	annc + 2 to ex + 10
Panel A: July 1963–December 2000					
1	543	2.563	1.177	0.522	1.475
2	544	1.961	1.262	0.822	2.829
3	544	2.252	1.267	1.251	3.166
4	543	3.086	1.551	0.530	2.884
5	544	2.622	2.061	−0.221	2.281
6	544	3.187	1.846	1.125	4.105
7	543	3.413	1.622	−0.021	2.445
8	544	3.759	1.571	1.050	2.982
9	544	4.860	2.136	1.899	4.376
10	543	5.535	2.655	3.643	6.571
10−1		[5.85]***	[3.74]***	[3.42]***	[4.91]***
Panel B: July 1963–December 1974					
1	89	2.977	0.735	1.136	2.737
2	90	1.605	0.807	1.456	3.524
3	90	2.018	1.134	−0.259	1.822
4	90	2.017	1.304	2.594	5.339
5	90	2.300	0.782	−0.518	−0.034
6	90	3.161	0.089	1.045	2.683
7	90	1.989	1.039	0.337	1.719
8	90	3.854	0.504	0.848	2.687
9	90	3.453	0.266	0.341	0.235
10	89	4.009	0.520	−0.233	2.150
10–1		[1.50]	[−0.66]	[−0.33]	[−0.23]

(*continued*)

Table 7 (*continued*)

Cumulative abnormal returns around stock-split announcement and ex-dates, sorted by price delay

		Mean cumulative abnormal return			
Delay decile	Number of firms	annc − 1 to annc + 1	annc + 2 to ex − 2	ex − 1 to ex + 1	annc + 2 to ex + 10
Panel C: January 1975–December 1987					
1	231	2.283	1.423	−0.498	1.325
2	232	2.787	1.350	−0.565	1.965
3	232	2.440	1.736	0.636	2.895
4	232	3.624	2.256	−0.411	3.269
5	232	3.313	2.370	0.573	3.873
6	232	3.558	2.494	0.662	4.440
7	232	4.165	2.348	−0.465	3.028
8	232	4.313	2.130	0.008	2.361
9	232	4.833	2.722	1.429	5.003
10	231	6.286	3.258	4.198	7.711
10−1		[5.13]***	[4.16]***	[2.85]***	[4.79]***
Panel D: 1988–December 2000					
1	222	2.483	1.081	1.459	1.222
2	222	1.611	1.358	2.325	3.748
3	222	1.853	0.429	2.812	3.940
4	222	2.712	1.639	−0.505	0.893
5	222	2.312	1.690	−0.506	1.734
6	222	2.804	1.921	1.201	3.913
7	222	2.972	1.317	0.647	3.017
8	222	4.150	1.770	3.123	4.472
9	222	4.952	2.482	3.493	5.362
10	222	5.201	2.332	3.371	6.088
10–1		[3.33]***	[1.37]	[1.67]*	[2.67]***

***, * indicate statistical significance at the 0.01 and 0.10 level, respectively.

In tests not presented in a table, we examine other visibility measures for the firms in each decile. For July 1963 to December 2000, the average NYSE size decile for the least delayed firms is 7.10, but the most price-delayed firms have an average NYSE size decile of 2.84. As one would expect, the least price-delayed firms are covered by many more analysts. For January 1988 to December 2000, firms in delay decile 1 are covered by an average of 15.09 analysts, but firms in delay decile ten are covered by an average of only 2.68 analysts. In the most price-delayed decile, only 38.3% of the splitting firms have any analyst coverage reported by the I/B/E/S database.[14]

[14] Hou and Moskowitz (2005) report that price delay is correlated with variables that signify a lack of attention by investors. For example, higher price delay is associated with lower analyst coverage, lower institutional ownership, measures related to costlier air travel, illiquidity, smaller size and fewer employees. They state (p. 1,017), "These results cannot be explained by microstructure, liquidity effects, market risk, or other known determinants of average returns, but appear most consistent with Merton's (1987) investor recognition hypothesis."

7. Conclusion

This paper reconciles the conflicting results of Byun and Rozeff (2003) and Ikenberry and Ramnath (2002). Our tests, covering stock splits announced between 1949 and 2000, reveal no long-term abnormal returns following the ex-date, after taking industry momentum into account as Moskowitz and Grinblatt (1999) suggest. Even without taking into account industry momentum, we find no consistent evidence of post-split positive abnormal returns, because we only see abnormal returns during the narrow window from 1975 to 1987.

Although we find no evidence of post-ex-date long-term drift, our results continue to support the idea that stock-split announcements are unambiguously positive news that is greeted by a positive stock price response. However, incorporation of this good news seems to be slowed by market frictions. Specifically, we find high correlations between the magnitude of split returns (including the delayed return) and the Hou and Moskowitz (2005) market friction measure.

In summary, the long-run post-announcement drift that several previous studies report, which is imprecisely referred to as "post-split" drift, actually occurs during a much shorter window, and the short-duration return drift that does occur seems to be related to market frictions rather than behavioral bias.

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