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The Long-run Performance Following Dividend Initiations and Resumptions: Underreaction or Product of Chance?

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ABSTRACT

We examine the long-term stock performance following dividend initiations and resumptions from 1927 to 1998. We show that postannouncement abnormal returns are significantly positive for equally weighted calendar time portfolios, but become insignificant when the portfolios are value weighted. Moreover, the equally weighted results are not robust across subsamples. We also document postannouncement reductions in the risk factor loadings of underlying stocks. Cross-sectionally, these reductions are negatively related to the contemporaneous price drifts, suggesting the price drifts may be a sample-specific result of chance. Our results underscore the importance of testing for changes in risk loadings in future long-term event studies.

An increasingly important issue in the study of financial markets is to understand the causes of the apparent abnormal equity price drifts or anomalies that are shown to follow various types of corporate events. In this paper, we address this issue within the context of dividend policy. We examine the long-term stock performance following the initiation and resumption of cash dividends during the period from 1927 to 1998. Although Michaely, Thaler, and Womack (1995) document a positive price drift for firms that initiate dividends during the 1964 to 1988 period, we can replicate their findings only for *equally weighted* portfolios and only for dividend events that occurred during the post-1964 period. This lack of intertemporal and methodological robustness suggests that the results of Michaely et al. may simply be due to chance.

We find no evidence of any abnormal price drift for the period prior to 1964. In addition, when portfolios are *value weighted* by market capitalization, the price drift of the post-1964 period disappears. The discrepancy between equal and value weighting is due to the absence of any long-term abnormal returns for the largest size decile. When this decile is excluded from the value-weighted analysis, the abnormal returns are again signifi-

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cantly positive. However, at the end of 1998, stocks in the largest decile represented about 88 percent of the total market capitalization of U.S. stocks trading on the NYSE, AMEX, and Nasdaq. This means that the observed price drift—to the extent that it represents an anomaly—is likely to be of limited significance from a macroeconomic perspective.

Still, from the perspective of a money manager, the price drift raises the question of whether a profitable trading rule could be implemented for the remaining lowest nine deciles. While accounting for only 12 percent of the U.S. market capitalization, stocks in these nine deciles were nevertheless worth an aggregate of $1.9 trillion at the end of 1998. Could markets be, to some degree, inefficient for 9 out of 10 stocks in the U.S. economy and allow them to be persistently subject to *predictable* mispricing? A rational interpretation of the price drift may reside in the significant reduction in the loadings of the three Fama and French (1993) risk factors we document during the postannouncement period.

In random industry samples, Fama and French (1997) show that declines in risk loadings are usually associated with positive abnormal returns, perhaps due to a combination of unexpectedly higher cash flows and lower discount rates. If our sample is overpopulated with firms that by chance become unexpectedly less risky or more profitable during the postannouncement period, we will expect rational investors to be surprised by the subsequent resolution of these uncertainties. Under this interpretation, the average firm in our sample would experience a positive postannouncement price drift, but this drift would occur *independently* from the dividend event. Moreover, the drift would be sample specific and would not generalize to the entire universe of firms that initiate or resume dividends. This interpretation appears quite plausible because the observed price drift is shown to lack robustness across firm sizes and time periods.

Our findings contribute to the ongoing debate surrounding the validity of the Efficient Markets Hypothesis. Along with the large number of studies reporting price drift anomalies following various corporate announcements, a behavioral stream of research has developed, seeking to attribute the observed anomalies to irrational investors who suffer from cognitive biases.[1] Theoretical models of behavioral finance were proposed by Daniel, Hirshleifer, and Subrahmanyam (1998), Odean (1998), Barberis, Shleifer, and Vishny (1998), and Hong and Stein (1999), among others. While competing with each other in terms of modeling approach, these papers nevertheless conclude that the reported anomalies provide clear evidence *against* the Efficient Markets Hypothesis.

[1] In addition to Michaely et al. (1995), researchers have reported significant price drifts following the announcement of share repurchases (Ikenberry, Lakonishok, and Vermaelen (1995)), stock splits (Ikenberry, Rankine, and Stice (1996)), new exchange listings (Dharan and Ikenberry (1995)), initial public offerings (Ritter (1991)), seasoned equity offerings (Loughran and Ritter (1995)), mergers (Agrawal, Jaffe, and Mandelker (1992)), and convertible debt issuance (Spiess and Affleck-Graves (1999)).

In response to this behavioral interpretation, Fama (1998) contends that the observed price drifts could result merely from chance, biased methodologies, or misspecified asset-pricing models. Fama's critique motivates testing the robustness of previously reported anomalies using different methodologies, sample periods, and asset pricing models. Mitchell and Stafford (2000) review the anomalies that were previously reported for corporate acquisitions, share repurchases, and seasoned equity offerings (SEOs). Using procedures that account for documented biases in the three-factor model of Fama and French (1993), they find no reliable evidence of any long-term anomaly for the three types of events they reexamine.

Similarly, Brav, Geczy, and Gompers (2000) and Eckbo, Masulis, and Norli (2000) are able to resolve the anomaly previously associated with SEOs. They argue that these firms become less risky following seasoned equity issuance, due to a corresponding decrease in leverage. As a result, SEO firms would command a lower expected return when compared to control firms or portfolios matched on preevent characteristics, which could explain the apparent underperformance in prior studies. In the same article, Brav et al. also resolve the initial public offering (IPO) anomaly by benchmarking these firms against control firms matched on size and book-to-market characteristics. Eckbo and Norli (2000) propose a different resolution of the IPO anomaly, after incorporating both leverage and liquidity as risk factors.

Our paper joins this stream of research by providing a careful scrutiny of the dividend anomaly reported by Michaely et al. (1995). Our main contribution is twofold. First, we show that the dividend price drift is confined to the period from 1964 to 1998 and does not extend to stocks belonging to the largest size decile. Second, we show that the observed price drifts are *negatively* related to contemporaneous declines in the Fama and French risk loadings of underlying stocks. This last result is in contrast to the SEO evidence in Brav et al. (2000) and Eckbo et al. (2000), where price drifts are *positively* correlated with subsequent risk changes. A negative relation between price drifts and contemporaneous changes in risk loadings—such as the one found in our study—suggests that these drifts could represent adjustments to risk changes or cash flow surprises that occur independently of the dividend event, and are *unforeseen* on the announcement date.

Our approach is different in several aspects from the one adopted by Michaely et al. (1995). First, we follow Mitchell and Stafford (2000) and measure abnormal returns using primarily equal- *and* value-weighted calendar time portfolios, as opposed to the (equally weighted) buy-and-hold methodology. Second, we examine—for the first time—the long-term stock performance following dividend *resumptions*, an event likely to have economic significance similar to dividend initiations. Third, we extend the sample to include firms that trade on Nasdaq, as well as firms with dividend announcements occurring before *and* after their original 1964 to 1988 study period.

The rest of the paper is organized as follows. Section I describes the data sources and sample selection. Section II outlines the research methodology. Section III examines the postannouncement abnormal performance. Section IV shows that postannouncement price drifts are negatively related to contemporaneous changes in the riskiness of underlying stocks. Section V concludes with a discussion of this paper's implications for future research.

I. Data and Sample Selection

Our data are obtained from the 1999 Center for Research in Security Prices (CRSP) database, except for the three factors of the Fama and French model, which were generously provided by Kenneth French. We first identify a sample of firms that either initiate or resume the payment of cash dividends from January 1927 to December 1998. We define *dividend initiation* as the first cash dividend in the history of a firm.[2] We define *dividend resumption* as the first cash dividend paid by a firm following a hiatus in payments ranging from 33 to 180 months. In order for an event to qualify as a dividend initiation or resumption, the underlying stock must meet the following additional criteria:

1. It must be traded on the NYSE, AMEX, or Nasdaq and must be an ordinary common stock of a U.S.-based corporation (CRSP share codes 10 and 11).
2. Its relevant cash dividends must be classified by CRSP as having either monthly, quarterly, semiannual, annual, or unspecified payment frequency. (Almost all "unspecified" frequencies are actually quarterly.)
3. Beginning with 1964, it must have been listed on the CRSP tape for 24 months before the dividend announcement. We impose this constraint in order to be consistent with the Michaely et al. (1995) selection criteria for their 1964 to 1988 period. For announcements before 1964, we require 12 months of listing on CRSP.

Our sample contains 1,645 initiations and 1,241 resumptions, for a total size of 2,886 dividend events. Some of our tests require that we divide our sample into two subperiods, using January 1964 as a cutoff date. The number of events during the earlier (1927 to 1963) and later (1964 to 1998) subperiods are 600 and 2,286, respectively. The high number of events in the later subperiod is due to the emergence of Nasdaq and to the continued growth in the number of stocks traded on the NYSE.

II. Research Methodology

Although previous studies of long-term performance generally favor the use of buy-and-hold abnormal returns (BHARs), we measure postevent abnormal performance using two *calendar time* methodologies: the Fama and French (1993) three-factor model and the mean monthly calendar time abnormal returns (CTARs). The BHARs and the closely related cumulative

[2] A firm is defined as a corporate entity having a unique PERMNO in the 1999 CRSP database.

abnormal returns (CARs) are particularly vulnerable to the problem of cross-sectional dependence among event firms, as documented by Mitchell and Stafford (2000). Cross-sectional dependence is of paramount concern in this study since our sample of over 2,800 dividend events is certainly affected by calendar time clustering and substantial overlapping of the postannouncement horizons.[3] Mitchell and Stafford explain that with the calendar time methodology, "monthly returns are less susceptible to the [misspecified] model problem" and "cross-correlations of event-firm abnormal returns are automatically accounted for in the portfolio variance" (p. 288).

The BHAR methodology—while certainly appealing to the investment community and arguably more representative of the overall investment experience—does not adequately control for cross-sectional correlation among individual firms in nonrandom samples and is, therefore, likely to yield t-statistics that are overstated (Mitchell and Stafford (2000)). Test statistics calculated from BHARs are also misspecified due to the severe skewness of the distribution of BHARs in most studies. The skewness problem is usually corrected by a bootstrapping method; however, bootstrapping does not address the issue of cross-sectional correlation.

Mitchell and Stafford (2000) also state that BHARs exacerbate the misspecified model problem. Since the true asset pricing model is not known, any potentially spurious "abnormal" return occurring at the beginning of the postevent period would be compounded by the BHAR measure, leading to significant biases over longer horizons.

A. Fama and French Three-factor Regression

A.1. Traditional Fama–French Model

We initially construct calendar time portfolios and calculate abnormal performance using the three-factor regression model developed by Fama and French (1993). For each calendar month, we calculate the monthly return to both equally weighted and value-weighted portfolios of firms that have been subject to dividend events during the $[c - h, c - 1]$ prior period, where c is the calendar month and h is the investment horizon of interest.[4]

[3] For example, 195 dividend events occurred during 1975, for an average of approximately 16 events per month.

[4] 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 since several previous studies that report significant abnormal returns 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 is also considered more representative of the wealth generating aspects of any presumed anomaly.

We then estimate the following three-factor regression model:

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

where $R_{p,t}$ represents the calendar time portfolio of dividend event firms, and $R_{f,t}$ is the return of the one-month Treasury Bill. The three independent variables are the excess return on the market portfolio ($R_{m,t} - R_{f,t}$), the difference between the returns of value-weighted portfolios of small and big firm stocks (SMB_t), and the difference in returns of value-weighted portfolios of high and low book-to-market stocks (HML_t). The intercept α_p is interpreted as the mean monthly abnormal return of the calendar time portfolio. Because the portfolio in any given month may contain from 1 to over 100 firms, we estimate this regression using both ordinary (OLS) and weighted least squares (WLS) procedures.

A.2. Adjusted Fama–French Model

Fama and French (1993) as well as Mitchell and Stafford (2000) show that the traditional three-factor model described in equation (1) does not completely explain the cross section of stock returns. For example, when equation (1) is estimated in randomly chosen samples of firms with both low book-to-market ratios and small sizes, the null hypothesis of zero abnormal performance is overrejected. In addition, Fama and French (1996) document a momentum bias for the three-factor model of equation (1).

Following the spirit of Mitchell and Stafford, we control for the potential size and momentum biases by estimating an "adjusted" Fama and French model. We build a hedge (zero-investment) calendar time portfolio consisting of long positions on event firms and short positions on control firms that are matched on both size and preevent momentum. We construct the match using the following algorithm: We first select all control firm candidates whose market value of equity as of the month prior to the event month is between 60 percent and 140 percent of the market value of equity of the corresponding sample firm. From this set, we identify the firm whose one-year preevent performance is closest to that of the sample firm. We define "preevent performance" as the holding period return over the $[m - 12, m - 1]$ period (where m is the month of dividend announcement).

We then regress the returns of this hedge portfolio on the three Fama and French factors:

$$R_{event,t} - R_{control,t} = \alpha_p + \beta_p(R_{m,t} - R_{f,t}) + s_p SMB_t + h_p HML_t + e_{p,t}. \qquad (2)$$

The "adjusted" Fama–French intercept obtained in this manner (α_p) represents a measure of long-term abnormal performance that specifically corrects for the size and momentum biases that are inherent in the traditional three-factor model. We report both the traditional and adjusted intercepts in the paper.[5]

A.3. Firm-specific Fama–French Regressions

The monthly loadings on the Fama–French factors are likely to vary with time, as individual firms enter and exit the calendar time portfolio (Mitchell and Stafford (2000)). The same argument can also be made in terms of entry and exit of various industries. To account for possible biases resulting from this time variance, we estimate individual Fama and French regressions for each event firm and compute the equally weighted cross-sectional mean of all firm-specific intercepts. We then estimate individual Fama–French regressions for each control firm and use their cross-sectional mean intercept to adjust the mean intercept obtained with the event sample.

B. Mean Monthly Calendar Time Abnormal Returns (CTARs)

We assess the robustness of our results by examining the mean monthly calendar time abnormal return (CTAR), calculated as

$$\text{mean monthly CTAR} = \sum_{t=1}^{n} \frac{CTAR_t}{n}, \quad t \in \{1, n\}, \tag{3}$$

where $CTAR_t = R_{p,t} - E(R_{p,t})$, t is the calendar month, and n is the total number of months in the time series under examination. In this representation, $R_{p,t}$ is the return of the calendar time portfolio comprising dividend event firms, constructed as in the previous Fama–French procedure. $E(R_{p,t})$ is the expected return of the event portfolio, computed as the raw return of a control portfolio, dynamically matched during each calendar month, according to size and momentum.[6] As with the Fama–French regression,

[5] Our adjusted procedure differs from that shown in equation (6) of Mitchell and Stafford (2000). The adjusted procedure employed in this paper produces an intercept identical to theirs; however, in our implementation, the resulting t-statistic of the intercept must properly account for the covariance between the event firm and control firm calendar time returns. We thank the referee for suggesting this variation of the adjusted Fama–French procedure.

[6] Because event firms are likely to experience changes in size and momentum throughout the postevent horizon, we identify, for each firm, a *different* match for every month of the postevent horizon. We construct the match according to the following algorithm: For each month we select all control firm candidates whose market value of equity as of the prior month is between 60 percent and 140 percent of the market value of equity of the corresponding sample firm. From this set, we identify the firm whose one-year prior performance is closest to that of the sample firm. We define one-year prior performance as the holding period return over the 12-month horizon ending with the prior month. This *dynamic* CTAR matching selects a new control firm for each month of the postevent horizon. Mitchell and Stafford (2000) use a similar dynamic matching procedure. (The previous control firms in subsection A of Section II are selected *only* on characteristics based on the month *before the event date* and consequently, each event firm is matched to only *one* control firm for the duration of the postevent horizon.)

the monthly cross-sectional average return of the calendar time portfolios, $CTAR_t$, is calculated using both equally weighted and value-weighted methodologies.

The monthly values of $CTAR_t$ are averaged intertemporally over the sample period of interest, as shown in equation (3), producing the mean monthly CTAR measure reported in the paper. The *t*-statistic of the mean monthly CTAR is calculated using the intertemporal standard deviation of the monthly $CTAR_t$, following the procedure detailed in Lyon, Barber, and Tsai (1999).

III. Postannouncement Abnormal Returns

A common characteristic of firms included in most long-term event studies is a period of abnormally positive or negative performance preceding the event. Fama and French (1996) document that prior momentum induces biases in the regression intercept of the three-factor model. Lyon et al. (1999) show that firms with high preevent momentum yield positively biased *t*-statistics in random samples over one-year horizons and negatively biased *t*-statistics over the three- and five-year horizons (p. 186). A more recent study by Lee and Swaminathan (2000) documents that, when combined with past trading volume, momentum extends to a three-year postselection horizon. Chan, Jegadeesh, and Lakonishok (1996) find that firms having above- or below-normal one-year stock return momentum continue to experience the momentum over a six-month postselection horizon. They further observe that this momentum effect is partially related to the well-documented post-earnings-announcement drift. Taken together, these studies strongly suggest that failing to control for preevent momentum in the measurement of long-term abnormal returns may lead to misleading inferences about the magnitude and causes of the reported anomalies.

Before we proceed with our main tests, we examine the preevent abnormal performance during the 12-month period preceding the dividend announcement, using the intercept from the Fama and French three-factor model (equal and value weighted, OLS and WLS). Our results demonstrate that dividend announcements are preceded by abnormally *high* stock returns (not shown). For example, the mean monthly Fama–French regression intercepts range from 0.77 percent to 1.41 percent per month, or 9.28 percent to 16.91 percent per year, depending on methodology. These results, which are comparable to those of Michaely et al. (1995), clearly highlight the importance of controlling for preevent momentum in this current study.

A. Adjusted Fama–French Analysis

We begin our main analysis by examining the long-term abnormal performance using the adjusted Fama–French procedure previously described. Table I reports the results for the one-, three-, and five-year postannouncement horizons, for the combined sample of initiations and resumptions during

Table I

Postannouncement Abnormal Returns Using Fama–French Calendar Time Portfolio Regressions

Abnormal returns are estimated for the one-, three-, and five-year postannouncement horizons for the combined sample of dividend initiations and resumptions during the period 1927 to 1998. Equal- and value-weighted (monthly rebalanced) calendar time portfolio returns are calculated each month from all firms experiencing a dividend initiation or resumption in the previous 12, 36, or 60 calendar months. The monthly excess returns to the calendar time portfolios, $R_{p,t} - R_{f,t}$, are regressed on the Fama and French (1993) three-factor model in order to calculate the *unadjusted* intercept:

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

$R_{f,t}$ is the return of the one-month T-bills. $(R_{m,t} - R_{f,t})$ is the excess return of the CRSP market portfolio (equal- and value-weighted index for equal- and value-weighted OLS regressions, respectively). SMB_t is the difference in returns between portfolios of small and big stocks. HML_t is the difference in returns between portfolios of high and low book-to-market ratio stocks. The regression intercept provides an estimate of monthly abnormal performance. For the *adjusted* Fama and French regression, a control firm calendar time portfolio (comprised of firms matched on size and preevent momentum) is subtracted from the event firm portfolio, and this difference is regressed on the three Fama and French factors:

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

Ordinary and weighted least squares (OLS and WLS) time series regressions are estimated. Monthly returns in the WLS model are weighted by the square root of the number of firms contained in the month. The t-statistics of the intercepts are shown in brackets under each parameter. The WLS model t-statistics are calculated using the White (1980) method.

		Horizon = 1 year (No. Obs. = 858 months)			Horizon = 3 years (No. Obs. = 858 months)			Horizon = 5 years (No. Obs. = 858 months)		
		Intercept of the Fama–French Regression		R-square of the Adjusted Fama–French Regression	Intercept of the Fama–French Regression		R-square of the Adjusted Fama–French Regression	Intercept of the Fama–French Regression		R-square of the Adjusted Fama–French Regression
Calendar Portfolio Weighting	Model Estimated	Unadjusted	Adjusted		Unadjusted	Adjusted		Unadjusted	Adjusted	
Equally weighted	OLS	0.00517	0.00365	0.030	0.00225	0.00252	0.006	0.00128	0.00127	0.082
		[3.27]***	[2.08]**		[2.78]***	[2.78]***		[2.03]**	[1.84]*	
	WLS	0.00565	0.00530	0.037	0.00253	0.00315	0.029	0.00174	0.00200	0.035
		[6.63]***	[5.20]***		[4.00]***	[4.81]***		[3.11]***	[3.75]***	
Value weighted	OLS	0.00294	0.00090	0.012	0.00128	0.00135	0.099	−0.00001	0.00010	0.158
		[1.54]	[0.41]		[1.02]	[1.03]		[−0.06]	[0.06]	
	WLS	0.00185	0.00139	0.063	0.00142	0.00101	0.135	0.00101	0.00099	0.142
		[1.14]	[0.75]		[1.31]	[0.89]		[1.09]	[1.00]	

***, **, * Significantly different from zero at the 1, 5, and 10 percent levels, respectively (two tail test).

1927 to 1998.[7] All of the adjusted intercepts for the *equally weighted* regressions (OLS and WLS) are positive and typically significant at either the 1 percent or 5 percent levels. For the three-year horizon, the adjusted intercepts suggest abnormal performance of 0.25 percent (t = 2.78) and 0.32 percent (t = 4.81) per month for the OLS and WLS models, respectively. Compounding these over 36 months implies three-year abnormal performance of 9.48 percent and 11.99 percent. Similarly, for the one-year horizon, the equally weighted adjusted intercepts suggest annual abnormal performance of 4.47 percent and 6.55 percent for the OLS and WLS models. In an economic sense, this abnormal performance appears large when compared to the realized equity market risk premium ($Rm - Rf$) of approximately 8.5 percent over the 1927 to 1998 period. The equally weighted results thus suggest that the *average firm* experiences significantly positive abnormal returns during the period following dividend announcements.

To account for the possible time variance in the loadings of the Fama and French model, we also estimate *firm-specific* Fama–French regressions and measure abnormal performance as the cross-sectional average of individual event intercepts (equally weighted), less the cross-sectional average of individual control intercepts. These results (not shown) tend to be of even greater magnitude than the equally weighted results shown in Table I and are always significant at better than the one percent level.

In contrast to the equally weighted results, the *value-weighted* adjusted intercepts—while always positive—do not achieve statistical significance for any horizon, whether estimated with OLS or WLS regressions. The largest adjusted intercept is obtained for the one-year WLS model, 0.14 percent (t = 0.75) per month, suggesting an economically weak abnormal performance of only 1.68 percent when compounded over the one year.

Loughran and Ritter (2000) argue that if selective management events (such as dividend announcements, IPOs, and SEOs) are motivated by industry- or economy-wide misvaluation, such events would be bunched in calendar time, and any resulting abnormal returns would likely be concentrated or bunched into the months containing large numbers of event firms. Our WLS regression technique should thus reveal any "event bunching" effect that may occur with "selective" management events. The similarity between the OLS and WLS results in Table I (and subsequent tables) provides no consistent support for this event-bunching hypothesis.

B. Mean Monthly Calendar Time Abnormal Returns Analysis

In Table II, we report the results of the mean monthly CTARs that are "dynamically matched" during each calendar month, based on size and prior one-year momentum. The equally weighted CTARs for the one-, three-, and

[7] Because the Fama–French factors begin coverage with July 1927, Table I reports the results of calendar time portfolios over the period from July 1927 to December 1998.

Table II

Postannouncement Mean Monthly Calendar Time Abnormal Returns

Abnormal returns are estimated for the one-, three-, and five-year postannouncement horizons, for the combined sample of dividend initiations and resumptions during the period 1927 to 1998. Equal- and value-weighted (monthly rebalanced) calendar time portfolio returns are calculated each month from all firms experiencing a dividend event in the previous 12, 36, and 60 calendar months. For every month containing an event firm in the portfolio, we match each firm to one control firm based on size and one-year momentum, as of the *prior month*. The Calendar Time Abnormal Return (CTAR) is calculated as the mean monthly difference between the returns of event-firm portfolios and the returns of control-firm portfolios. *T*-statistics are shown in brackets and are calculated from the intertemporal variance of monthly CTARs.

	Mean Monthly Calendar Time Abnormal Returns (CTAR) *Dynamically* Matched on Size and Momentum	
Postannouncement Horizon	Equally Weighted Calendar Portfolios	Value-Weighted Calendar Portfolios
1 year	0.00581	0.00369
(no. obs. = 863 months)	[2.67]***	[1.62]
3 year	0.00307	0.00232
(no. obs. = 863 months)	[2.84]***	[1.66]*
5 year	0.00241	0.00099
(no. obs. = 863 months)	[2.71]***	[0.90]

***, * Significantly different from zero at the 1 and 10 percent levels, respectively (two tail test).

five-year horizons are all positive and statistically significant at better than the one percent level. While the value-weighted results are positive in sign, only the three-year CTAR achieves significance at the 10 percent level. For the three-year horizon, the mean monthly CTARs are 0.31 percent (t = 2.84) and 0.23 percent (t = 1.66) for the equal- and value-weighted portfolios, respectively. The magnitudes of these CTARs are larger than the corresponding adjusted regressions of Table I; however, they control only for size and past one-year performance—they do not control for the book-to-market and market portfolio characteristics.

C. Long-term BHARs and CARs

For completeness, we also examine postevent BHARs and CARs following the methodology outlined in Barber and Lyon (1997). We estimate the long-term 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

Table III

Postannouncement Cumulative and Buy-and-Hold Abnormal Returns

Shown are the mean cumulative abnormal returns (CAR) and the mean buy-and-hold abnormal returns (BHAR) for each indicated postannouncement horizon. The CARs are calculated by subtracting the sum of monthly returns of a matching control benchmark from the sum of monthly returns of the corresponding dividend initiating or resuming firm. BHARs are calculated by subtracting the buy-and-hold return of the control benchmark from the buy-and-hold return of the corresponding event firm. The benchmark used for each event firm is a control firm of similar size and 12-month preevent momentum. Both CARs and BHARs are equally weighted. *T*-statistics are shown in brackets under each parameter. In the case of BHARs, *t*-statistics are skewness adjusted, and significance levels are computed using the bootstrapping procedure outlined in Lyon, Barber, and Tsai (1999). For calculating the value-weighted CAR and BHAR, we standardize each firm's market capitalization (calculated for the event month) by the level of the CRSP value weighted index.

Postannouncement Horizon	Average Cross-sectional Abnormal Returns			
	Equally Weighted		Value Weighted (Market Value Standardized with the Level of the CRSP Value-weighted Index)	
	CAR	BHAR	CAR	BHAR
1 year	0.05721	0.07892	0.01125	0.01076
(no. obs. = 2,652)	[4.78]***	[5.03]***	[1.22]	[0.96]
3 year	0.10427	0.21728	0.09558	0.12491
(no. obs. = 2,565)	[5.24]***	[5.04]***	[5.98]***	[4.50]***
5 year	0.09596	0.24884	0.12376	0.23467
(no. obs. = 2,395)	[3.75]***	[2.56]**	[5.68]***	[4.14]***

***, ** Significantly different from zero at the 1 and 5 percent levels, respectively (two tail test).

and one-year momentum.[8] We compute the cross-sectional averages (equal and value weighted) of these return differences and report the results in Table III.[9] All of the equally weighted CARs and BHARs are positive and statistically significant at either the one percent or five percent level. Although the BHAR test statistics are adjusted for skewness using the bootstrapping procedure described in Lyon et al. (1999), they are not corrected for cross-sectional correlation. For the three-year horizon, we report a mean

[8] We construct the match using the same algorithm employed in the adjusted Fama–French procedure. We first select all control firm candidates whose market value of equity as of the month prior to the event month is between 60 percent and 140 percent of the market value of equity of the corresponding sample firm. From this set, we identify the firm whose one-year preevent performance is closest to that of the sample firm. We define "preevent performance" as the holding period return over the $[m - 12, m - 1]$ period (where m is the month of dividend announcement).

[9] For calculating the value-weighted results, we standardize each firm's market capitalization (calculated for the event month) by the level of the CRSP value weighted index.

equally weighted CAR and BHAR of 10.43 percent ($t = 5.24$) and 21.73 percent ($t = 5.04$), respectively. In comparison, Michaely et al. (1995) report market-adjusted BHARs of 7.5 percent and 24.8 percent for the one- and three-year postevent horizons.

The *value-weighted* mean CARs and BHARs are insignificant for the one-year horizon, but achieve significance for the three- and five-year horizons. These results conflict with the value-weighted adjusted Fama–French intercepts reported in Table I, where the strongest results are reported for the one-year horizon.

The results in Table III reiterate Mitchell and Stafford's (2000) argument that over long horizons, BHARs tend to magnify spurious abnormal performances induced by potentially misspecified asset pricing models. For example, the three-year BHAR in Table III (21.73 percent) is more than twice the size of the abnormal return that is implied by the adjusted Fama–French regression for the same horizon (9.07 percent for the equally weighted OLS regressions). In addition, both the equal- and value-weighted five-year BHARs are larger in magnitude than the corresponding three-year BHARs, which *could* suggest the presence of positive abnormal performance during the fourth and fifth year following the dividend announcement. The next subsection, however, shows that such inference would be misleading.

These results are also consistent with Mitchell and Stafford's (2000) argument that cross-sectional dependence leads to overinflated test statistics. For example, although the three-year equally weighted CAR (10.43 percent, $t = 5.24$) is comparable in magnitude to the abnormal performance implied from the adjusted OLS intercept in Table I (0.25 percent per month, 9.07 percent over three years, $t = 2.78$), the t-statistic of the CAR measure is almost twice the size of its adjusted Fama–French counterpart. A similar difference in magnitude is observed with the t-statistics of the BHAR measure. This upward bias most likely results from failing to account for cross-sectional dependence among event firms when computing t-statistics with the CAR and BHAR methodologies.

D. Duration of Long-term Abnormal Returns

Although the equally weighted results in Tables I, II, and III reveal evidence of significantly positive abnormal returns for the one-, three-, and five-year horizons, they do not allow us to infer the exact duration of the abnormal performance. To obtain a better estimate of this duration, we repeat most of our tests for two intermediate postevent horizons: months 13 to 36 and months 37 to 60.

Table IV presents the results. For postevent months 13 to 36, the *equally weighted* adjusted Fama–French intercepts and CTARs indicate positive abnormal performance that is significant at better than the five percent level. However, the unadjusted Fama–French intercepts are insignificant, suggesting that the higher t-statistics of the adjusted intercepts result from a negative abnormal performance of the control firm portfolio. The magnitude of

Table IV

Duration of Long-term Abnormal Returns

Fama–French and mean monthly calendar time abnormal returns are estimated for the 1927 to 1998 period in order to examine abnormal performance during postannouncement years 2 and 3 and during postannouncement years 4 and 5. All models are estimated for the combined sample of dividend initiations and resumptions. Equal- and value-weighted calendar time portfolio returns are calculated each month from firms experiencing dividend announcements during the previous 13 to 36 months, and, alternatively, during the previous 37 to 60 months. For the *unadjusted* Fama and French (1993) three-factor model, the monthly excess returns of the calendar time portfolios, $R_{p,t} - R_{f,t}$ ($R_{f,t}$ is the return of the one-month T-bill) are regressed on (1) $(R_{m,t} - R_{f,t})$, the excess return of the CRSP market portfolio (equal- and value-weighted index for equal- and value-weighted regressions, respectively); (2) SMB_t, the difference in returns between portfolios of small and big stocks; and (3) HML_t, the difference in returns between portfolios of high and low book-to-market ratio stocks. For the *adjusted* Fama and French regression, a control firm calendar time portfolio (comprised of firms matched on size and preevent momentum) is subtracted from the event firm portfolio, and this difference is regressed on the three Fama and French factors. The regression intercepts provide an estimate of monthly abnormal performance. Estimated are both ordinary and weighted least squares (OLS and WLS) regressions. The WLS model weights each month with the square root of the number of firms for that month. To obtain the *dynamically matched* calendar time abnormal returns (CTAR) during each month containing an event firm, we match every event firm with a control firm of similar size and momentum, computed as of the *prior* calendar month. The mean monthly CTAR is calculated as the mean monthly difference in returns between the calendar time portfolios of event and control firms. T-statistics are shown in brackets below each entry. The WLS model t-statistics are calculated using the White (1980) method. The CTAR t-statistics are calculated from the intertemporal variance of monthly CTARs.

			Horizon = Postevent Months 13 to 36 (No. Obs. = 852 Months)				Horizon = Postevent Months 37 to 60 (No. Obs. = 828 months)			
				Intercept of the Fama–French Regression				Intercept of the Fama–French Regression		
Estimation Method	Calendar Portfolio Weighting	Model Estimated	Mean Monthly CTAR	Unadjusted	Adjusted	R-square of the Adjusted Fama–French Regression	Mean Monthly CTAR	Unadjusted	Adjusted	R-square of the Adjusted Fama–French Regression
Fama–French regressions	Equally weighted	OLS		0.00128	0.00210	0.030		0.00107	−0.00102	0.140
				[1.60]	[2.33]**			[0.65]	[−1.29]	
		WLS		0.00050	0.00150	0.021		0.00037	−0.00089	0.058
				[0.78]	[2.38]**			[0.67]	[−1.63]	
	Value weighted	OLS		0.00083	0.00142	0.104		−0.00097	−0.00396	0.225
				[0.71]	[1.13]			[−0.53]	[−3.61]***	
		WLS		0.00073	0.00115	0.107		−0.00086	−0.00091	0.075
				[0.69]	[0.96]			[−0.99]	[−0.88]	
Mean monthly CTAR (dynamically matched)	Equally weighted		0.00298				0.00152			
			[2.46]**				[0.87]			
	Value weighted		0.00216				0.00226			
			[1.73]*				[1.12]			

***, **, * Significantly different from zero at the 1, 5, and 10 percent levels, respectively (two tail test).

the equally weighted monthly abnormal performance is also smaller than that reported in Table I for the first year, suggesting that the highest post-announcement abnormal returns occur in the first 12 months following the dividend announcement.

An examination of the corresponding value-weighted results reveals further mixed evidence for months 13 to 36: The adjusted Fama–French intercepts are insignificant, while the CTAR is positive and significant at the 10 percent level.

For postevent months 37 to 60, the adjusted value-weighted Fama–French intercepts are actually *negative*, although not always significant. By contrast, the CTARs for that same horizon are positive, albeit insignificant.

We thus find (at best) weak evidence of abnormal performance during years two and three, and no reliable evidence of any significant abnormal performance during years four and five following dividend announcements. Because any postannouncement anomaly appears confined to, at most, the three years following the announcement, and because later in this paper we show that our event firms experience a significant decrease in risk during this same three-year horizon, we focus the remaining analysis on the three-year postannouncement horizon.

E. Analysis by Market Capitalization Size Deciles

Our results thus far demonstrate that the long-term abnormal returns following the initiation and resumption of dividends are significantly positive *only* when the calendar time portfolios are equally weighted. In contrast, the value-weighted results are weak and statistically insignificant, suggesting that the price drift is confined to the smaller firms in our sample. Perhaps if a sufficient number of the largest firms are omitted from the study, the remaining firms will produce statistically significant value-weighted results. We explore this possibility by splitting the total sample into two calendar time portfolios: one comprised of firms contained in CRSP size deciles 1 to 9 and the other comprised of decile 10 firms.[10] Decile 1 contains the smallest firms. We report the value-weighted adjusted Fama and French intercepts in Table V for the three-year postannouncement horizon. For the firms listed in the first nine deciles, both the adjusted and unadjusted intercepts are positive and significant at traditional levels. The adjusted intercepts imply abnormal performance of 0.26 percent (t = 2.10) and 0.22 percent (t = 2.22) per month for the OLS and WLS models, respectively.

The value-weighted intercepts for decile 10 are of low negative magnitude and never attain statistical significance. It would therefore appear that although firms in the largest decile are not subject to any long-term price drift

[10] We use the size decile rankings from CRSP. These deciles are assigned annually, based on the entire universe of CRSP listed NYSE, AMEX, and Nasdaq firms.

Table V

Size Decile Analysis for the Value-weighted Portfolios

Value-weighted Fama–French intercepts are estimated for the 1927 to 1998 period in order to examine the *separate* long-term reactions to firms contained in CRSP size deciles 1 to 9 and size decile 10. CRSP deciles are computed based on the entire universe of CRSP listed firms. Value-weighted calendar time portfolio returns are calculated each month from firms experiencing dividend announcements during the previous 36 months. For the *unadjusted* Fama and French (1993) three-factor model, the monthly excess returns of the calendar time portfolios, $R_{p,t} - R_{f,t}$ ($R_{f,t}$ is the return of the one-month T-bill) are regressed on (1) ($R_{m,t} - R_{f,t}$), the excess return of the CRSP market portfolio (equal- and value-weighted index for equal- and value-weighted regressions, respectively); (2) SMB_t, the difference in returns between portfolios of small and big stocks; and (3) HML_t, the difference in returns between portfolios of high and low book-to-market ratio stocks. For the *adjusted* Fama and French regression, a control firm calendar time portfolio (comprised of firms matched on size and preevent momentum) is subtracted from the event firm portfolio, and this difference is regressed on the three Fama and French factors. The regression intercepts provide an estimate of monthly abnormal performance. Estimated are both ordinary and weighted least squares (OLS and WLS) regressions. The WLS model weights each month with the square root of the number of firms for that month. The t-statistics are shown in brackets under each parameter. The WLS model t-statistics are calculated using the White (1980) method.

	Size Deciles 1 through 9			Size Decile 10		
	Intercept of the Fama–French Regression		R-square of the Adjusted Fama–French Regression	Intercept of the Fama–French Regression		R-square of the Adjusted Fama–French Regression
Model Estimated	Unadjusted	Adjusted		Unadjusted	Adjusted	
OLS	0.00221	0.00262	0.058	−0.00188	−0.00208	0.060
	[1.88]*	[2.10]**		[−0.66]	[−0.65]	
WLS	0.00180	0.00215	0.119	−0.00012	0.00020	0.084
	[1.96]*	[2.22]**		[−0.06]	[0.08]	

**, * Significantly different from zero at the 5 and 10 percent levels, respectively (two tail test).

anomaly, the remaining 90 percent of firms experience significantly positive postannouncement abnormal returns. This likely explains the reason why equally weighted results in Table I are statistically significant while the corresponding value-weighted results are not.

At first glance, the results of Table V imply that an investment strategy involving postdividend purchases of equities in size deciles 1 to 9 could produce statistically and economically significant abnormal returns. Although stocks in these nine smallest deciles comprise only 12 percent of the total U.S. market capitalization, they were nevertheless worth an aggregate of \$1.9 trillion at the end of 1998. This means that a trading strategy that identifies even a small annual rate of "true" abnormal performance could potentially result in very large amounts of arbitrage profits. Whether or not such profits are possible depends largely on whether the price drift results from irrational mispricing, as opposed to being a mere sample-specific coincidence. We thus continue to scrutinize the price drift by examining its robustness across various subperiods of our initial sample. If a putative anomaly is not robust to such scrutiny, it likely remains a sample-specific result and possibly only a curious artifact of chance.

F. Out-of-sample Extensions of the Michaely et al. (1995) Study

We examine various extensions to the original sample employed by Michaely et al. (1995) and report the results in Table VI. We begin by focusing on dividend initiations that were announced during their 1964 to 1988 sample period. Unlike Michaely et al., who use BHARs, we measure abnormal performance with the adjusted Fama–French model. For this particular subsample, the equal- and value-weighted adjusted OLS and WLS intercepts are positive and statistically significant. For example, the adjusted intercepts are 0.35 percent (t = 2.41) and 0.51 percent (t = 2.50) per month for the equal- and value-weighted OLS regressions, respectively. However, the significant value-weighted results may be due to the unusually low incidence of initiations among large firms from 1964 to 1988, when only 1.60 percent of the initiations occur within CRSP size decile 10.

We next extend Michaely et al.'s (2000) original 1964 to 1988 sample along two out-of-sample dimensions. First, we examine dividend events announced during the 1927 to 1963 and 1989 to 1998 periods that are immediately adjacent to the original period. Second, we examine a related dividend event, dividend resumptions.

Our extension to dividend initiations announced during the period 1927 to 1963 reports no significant abnormal performance for either equal- or value-weighted adjusted intercepts, and the adjusted intercept is actually negative in the case of the OLS value-weighted regression. During our other subperiod (1989 to 1998), the sample of dividend initiations reports significantly positive equally weighted adjusted intercepts of 0.56 percent (t = 2.94) and 0.58 percent (t = 3.59) per month for the OLS and WLS regressions, respectively. However, the corresponding value-weighted results—while seemingly large in magnitude—are statistically insignificant, perhaps due to the high percentage (6.90 percent) of firms that belong to the largest CRSP size decile. Such firms are shown to produce insignificant intercepts in the size decile analysis of Table V.

Results for the 1995 to 1998 subperiod, subsequent to the publication of Michaely et al. (2000), are also reported in Table VI. Similar to the 1989 to 1998 period, the equally weighted results are large in magnitude and statistically significant, while the value-weighted results, although seemingly high in magnitude, do not attain significance. Once again, this discrepancy is likely due to the 9.17 percent of initiating firms that belonged to CRSP's largest size decile during that period.

Table VI

Separate Analysis of Initiations and Resumptions for Various Time Periods

Equally and value-weighted *adjusted* Fama–French intercepts are estimated separately for initiations, resumptions, and all firms in order to examine the *separate* long-term reactions for announcements occurring in the 1927 to 1963, 1964 to 1999 (and subperiods of the 1964 to 1999 period), and 1927 to 1999 periods. Equally and value-weighted calendar time portfolio returns are calculated each month from firms experiencing dividend announcements during the previous 36 months. For the *adjusted* Fama and French regression, a control firm calendar time portfolio (comprised of firms matched on size and preevent momentum) is subtracted from the event firm portfolio, and this difference is regressed on the three Fama and French (1993) factors: (1) $(R_{m,t} - R_{f,t})$, the excess return of the CRSP market portfolio (equal- and value-weighted index for equal- and value-weighted regressions, respectively); (2) SMB_t, the difference in returns between portfolios of small and big stocks; and (3) HML_t, the difference in returns between portfolios of high and low book-to-market ratio stocks. The regression intercepts provide an estimate of monthly abnormal performance. Both ordinary and weighted least squares (OLS and WLS) regressions are estimated. The WLS model weights each month with the square root of the number of firms for that month. The t-statistics are shown in brackets under each parameter. The WLS model t-statistics are calculated using the White (1980) method.

Type of Dividend Event	Model Estimated	Time Period Estimated 1927–1963	1964–1988	1989–1998	1964–1998	1995–1998	1927–1998
Panel A: Equally Weighted Calendar Time Portfolios (Three Year Postannouncement Horizon)							
Initiations	OLS	0.00127	0.00348	0.00557	0.00377	0.00723	0.00213
		[0.34]	[2.41]**	[2.94]***	[3.26]***	[1.97]*	[1.07]
	WLS	0.00109	0.00349	0.00580	0.00414	0.00834	0.00348
		[0.39]	[3.29]***	[3.59]***	[4.67]***	[3.00]***	[4.11]***
Resumptions	OLS	0.00301	0.00154	0.00359	0.00240	0.00382	0.00289
		[1.12]	[0.90]	[1.60]	[1.76]*	[0.92]	[1.92]*
	WLS	0.00208	0.00123	0.00457	0.00267	0.00625	0.00270
		[1.73]*	[0.84]	[1.96]*	[2.20]**	[1.33]	[3.09]***
Initiations and resumptions	OLS	0.00188	0.00286	0.00489	0.00337	0.00572	0.00252
		[1.20]	[2.67]***	[3.11]***	[3.83]***	[1.84]*	[2.78]***
	WLS	0.00173	0.00273	0.00536	0.00363	0.00733	0.00315
		[1.51]	[3.09]***	[3.69]***	[4.77]***	[2.26]**	[4.81]***

Panel B: Value-weighted Calendar Time Portfolios (Three Year Postannouncement Horizon)							
Initiations	OLS	−0.00066	0.00512	0.00387	0.00439	0.00509	0.00132
		[−0.18]	[2.50]**	[0.95]	[2.38]**	[0.64]	[0.63]
	WLS	0.00127	0.00520	0.00320	0.00393	0.00551	0.00345
		[0.31]	[3.38]***	[0.88]	[2.44]**	[0.80]	[2.25]**
Percentage of firms belonging to CRSP size decile 10		1.98%	1.60%	6.90%	3.04%	9.17%	2.98%
Resumptions	OLS	0.00670	−0.00398	0.00115	−0.00293	−0.00307	0.00168
		[2.26]**	[−1.47]	[0.37]	[−1.38]	[−0.65]	[0.93]
	WLS	0.00233	−0.00353	0.00182	−0.00203	−0.00098	−0.00069
		[1.23]	[−1.40]	[0.64]	[−1.07]	[−0.20]	[−0.51]
Percentage of firms belonging to CRSP size decile 10		2.80%	4.33%	8.94%	5.94%	8.82%	4.67%
Initiations and resumptions	OLS	0.00266	0.00039	0.00388	0.00109	0.00407	0.00135
		[1.25]	[0.23]	[1.25]	[0.73]	[0.64]	[1.03]
	WLS	0.00179	0.00153	0.00346	0.00171	0.00455	0.00101
		[1.01]	[1.01]	[1.30]	[1.29]	[0.80]	[0.89]
Percentage of firms belonging to CRSP size decile 10		2.66%	2.42%	7.68%	3.98%	9.01%	3.71%

***, **, * Significantly different from zero at the 1, 5, and 10 percent levels, respectively (two tail test).

In Table VI, we also extend the analysis to dividend resumptions, our second out-of-sample dimension. Resumptions are associated with positive equally weighted adjusted intercepts during the 1964 to 1998 period, achieving adjusted intercepts of 0.24 percent ($t = 1.76$) and 0.27 percent ($t = 2.20$) per month for the OLS and WLS models, respectively. However, most of this effect appears to be concentrated in the 1989 to 1998 period. Resumptions are also associated with relatively high equally weighted adjusted intercepts during the period 1927 to 1963, but only the WLS intercept attains marginal significance at the 10 percent level. For the entire 1927 to 1998 period, the equally weighted adjusted intercepts are positive (nearly 0.30 percent per month) and statistically significant. However, in the case of value-weighted results, resumptions are only significant for the 1927 to 1963 OLS regression, and this result is not robust to the WLS specification.

Finally, we examine the intertemporal robustness of the combined portfolio containing dividend initiations and resumptions. Our results show that the equally weighted adjusted intercepts during the period 1964 to 1998 and subsets of that period are all positive and statistically significant. However, none of the corresponding value-weighted intercepts are statistically significant. In addition, all of the equally and value-weighted combined sample results for the 1927 to 1963 period are statistically insignificant. We thus find no convincing evidence of any long-term abnormal returns for the period prior to 1964. After that date, the abnormal returns are confined to firms in the smallest nine size deciles.

IV. Postannouncement Changes in Risk Factor Loadings

We now attempt to rationalize the observed price drift within the framework of the Efficient Markets Hypothesis. We observe significant decreases in the loadings of the three Fama–French factors during the postannouncement period and show that they are cross-sectionally related to the concurrent price drifts: Firms with the largest reductions in risk factor loadings are also the ones with the biggest increases in stock prices. We identify two plausible rational explanations for this phenomenon, which are neither mutually exclusive nor collectively exhaustive.

First, a decrease in risk loadings might represent an unpredicted decrease in the corporate cost of equity that is unrelated to dividends. If the market does not fully anticipate this decrease on the day of the dividend announcement, stock prices will increase during the postannouncement period, as investors gradually discover the lower discount rate and incorporate it into stock prices.

Second, a decrease in the loadings of the SMB and HML factors is generally associated with a period of unexpectedly stronger cash flows (Fama and French (1997)). If our sample is overpopulated with firms that are *by chance* performing better than expected during the postannouncement period, such firms will simultaneously experience both positive abnormal returns and concurrent reductions in the SMB and HML loadings.

We next explore these two possibilities.

A. Abnormal Returns and Unexpected Reductions in the Corporate Cost of Equity

A.1. Aggregate Changes in the Cost of Equity

We begin with a test of postannouncement changes in the three Fama–French risk loadings. For each sample firm, we estimate the following Fama and French regression for the $[m - 36, m + 36]$ period, where m is the month of the dividend announcement:[11]

$$\begin{aligned} R_{i,t} - R_{f,t} &= \alpha_i + D_t \alpha_{\Delta i} + \beta_i (R_{m,t} - R_{f,t}) + s_i SMB_t + h_i HML_t \\ &\quad + \beta_{\Delta i} D_t \,(R_{m,t} - R_{f,t}) + s_{\Delta i} D_t SMB_t + h_{\Delta i} D_t HML_t + e_{i,t}. \end{aligned} \tag{4}$$

In this representation, D_t is a dummy variable equal to one if the month is in the postevent period (zero otherwise).[12] We compute the cross-sectional averages of all firm-specific regression coefficients and report the results in Table VII. Our focus is on coefficients $\beta_{\Delta i}$, $s_{\Delta i}$, and $h_{\Delta i}$, which provide estimates of postevent changes in firm-specific risk factor loadings. The cross-sectional averages of these three coefficients are all negative and significant at better than the one percent level.[13] This suggests that the risk factor loadings are significantly lower during the $[m + 1, m + 36]$ period when compared to $[m - 36, m - 1]$.

If the Fama–French model is the correct asset pricing model, we can use the three changes in risk loadings to obtain a proxy for the change in the required cost of equity. To do so, we multiply the cross-sectional average of each change in risk loadings ($\beta_{\Delta i}$, $s_{\Delta i}$, and $h_{\Delta i}$), by the mean monthly realization of the corresponding risk factor during the period 1927 to 1998 (mean-$[Rm - Rf]$, mean$[SMB]$, mean$[HML]$). We interpret the sum of the three products as the average change in the required cost of equity between the pre- and postevent periods.

For clarity of the exposition, we henceforth use the terms "risk change," "risk increase," and "risk decrease" to denote postannouncement changes in the required cost of equity. As shown in the bottom section of Table VII, we find that the average firm in our sample experiences a large risk decrease, equal to approximately −0.17 percent per month ($t = -6.64$) over the three-

[11] The announcement month is itself excluded from the regression model.

[12] Using a similar method, Grullon, Michaely, and Swaminathan (2001) document a *reduction* in equity systematic risk following large increases in *existing* quarterly dividends. However, they do not examine changes in systematic risk in relation with dividend initiations or resumptions. Unlike these authors, we focus our attention on firms that either *initiate* or *resume* regular dividends.

[13] For example, the mean beta of the event firms drops from 1.12 to 1.02 after the dividend announcement ($t = -4.31$).

Table VII

Postevent Risk Changes

We examine the postannouncement changes in equity risk loadings by estimating the following *firm-specific* regressions, based on the Fama and French (1993) three-factor model:

$$R_{i,t} - R_{f,t} = \alpha_i + D_t\alpha_{\Delta i} + \beta_i(R_{m,t} - R_{f,t}) + s_i SMB_t + h_i HML_t$$
$$+ \beta_{\Delta i} D_t(R_{m,t} - R_{f,t}) + s_{\Delta i} D_t SMB_t + h_{\Delta i} D_t HML_t + e_{i,t}.$$

Each regression is performed on a 72-month window: 36 months before and 36 months after the announcement (the month of announcement is excluded). The dummy variable D_t is equal to one (zero if otherwise) if the calendar month is in the postevent period, and thus the three coefficients $\beta_{\Delta i}$, $s_{\Delta i}$, and $h_{\Delta i}$ represent estimates of the postevent changes in the loadings of the three Fama–French risk factors. In the top section of the table, we report the preevent cross-sectional averages of the three Fama–French risk coefficients (β_i, s_i, h_i), as well as the averages of the postevent changes in each of these three coefficients ($\beta_{\Delta i}$, $s_{\Delta i}$, $h_{\Delta i}$). The middle section shows the intertemporal mean monthly average of the actual Fama–French risk factors: ($R_{m,t} - R_{f,t}$), SMB_t, and HML_t. Each of the three risk change coefficients ($\beta_{\Delta i}$, $s_{\Delta i}$, $h_{\Delta i}$) is multiplied by the respective mean of the corresponding Fama–French risk factor ($R_{m,t} - R_{f,t}$, SMB_t, HML_t), and the summation of the three products is reported in the bottom section as the average monthly change in the required rate of return on equity during the 36-month postevent horizon.

	Parameter Value	*t*-statistic
Cross-sectional averages of the risk-change regression coefficients		
Average preevent coefficients		
β_i	1.11556	[65.63]***
s_i	0.25920	[8.20]***
h_i	0.03784	[1.42]
Average change in risk coefficients after the event		
$\beta_{\Delta i}$	−0.09581	[−4.31]***
$s_{\Delta i}$	−0.15033	[−3.71]***
$h_{\Delta i}$	−0.10869	[−3.06]***
Number of firm-specific cross-sectional regressions estimated	2,593	
Intertemporal mean monthly realizations of the Fama–French factors		
mean[$Rm - Rf$], equally weighted	0.01003	[3.86]***
mean[SMB]	0.00184	[1.71]*
mean[HML]	0.00379	[3.10]***
Average monthly change in the required return on equity around dividend initiations and resumptions	−0.00165	[−6.64]***

***, * Significantly different from zero at the 1 and 10 percent levels, respectively (two tail test).

year postannouncement period. We attempt to detect any additional risk changes that may have occurred during the fourth and fifth years, but find that these changes are small in magnitude and statistically insignificant (not shown). We also examine postannouncement risk changes for dividends announced during the 1927 to 1963 and 1964 to 1998 periods (not shown) and find that such changes are smaller during the pre-1964 period: The monthly risk changes are −0.13 percent and −0.18 percent per month for the 1927 to 1963 and 1964 to 1998 periods, respectively. The pre-1964 figure is consistent with the results in Table VI, which indicate lower abnormal stock performance during that earlier period.

While the event-time, firm-specific regression approach of equation (4) provides evidence of significant postannouncement risk decreases, the test statistics assume cross-sectional independence among the event firms. To control for possible cross-correlation, we explore an alternative test for risk changes, based on calendar time portfolios.[14] We estimate the following hedge regression:

$$R_{postevent,t} - R_{preevent,t} = \alpha_p + \beta_p(R_{m,t} - R_{f,t}) + s_p SMB_t + h_p HML_t + e_{p,t}, \tag{5}$$

where $R_{preevent,t}$ and $R_{postevent,t}$ are the returns of pre- and postevent calendar time portfolios, comprised of firms with dividend events occurring during the 36 months before and after the dividend announcement. The coefficients β_p, s_p, and h_p provide estimates of the differences in the Fama–French risk loadings between the post- and preannouncement periods, while accounting for the cross-sectional correlation between event firms.[15] When each of these coefficients is multiplied by the mean monthly realization of the corresponding risk factor, the sum of the three products provides an estimate of the postannouncement change in the cost of equity. We assess the statistical significance of this risk change by testing the following linear restriction of equation (5):

$$\beta_p * \text{mean}(R_{m,t} - R_{f,t}) + s_p * \text{mean}(SMB_t) + h_p * (\text{mean}HML_t) = 0, \tag{6}$$

where mean$(R_{m,t} - R_{f,t})$, mean(SMB_t), and (meanHML_t) are the mean monthly realizations of the risk factors, estimated over the period from 1927 to 1998.

Equation (5), along with linear restriction (6), is estimated for the entire 1927 to 1998 period and for the 1927 to 1963 and 1964 to 1998 subperiods. The results for the 1927 to 1998 period suggest monthly risk changes of −0.04 percent ($t = -1.08$) and −0.10 percent ($t = -3.20$), for the OLS and WLS specifications, respectively, and thus provide mixed evidence on the

[14] We thank the referee for suggesting this calendar time approach for testing risk changes.

[15] Both ordinary and weighted least squares regressions are estimated. The WLS regression weights each month by the geometric average of the number of firms in the preevent and postevent portfolios.

significance of risk changes across the full period. This mixed evidence, however, is just a consequence of a major intertemporal difference that exists between risk changes occurring before and after January of 1964. When equations (5) and (6) are estimated separately for the two subperiods, we observe a strong, negative risk change during the period 1964 to 1998, but no significant risk change prior to 1964. The OLS and WLS regressions suggest risk changes of −0.14 percent ($t = -4.96$) and −0.18 percent ($t = -7.93$) per month, respectively, for the 1964 to 1998 subperiod and risk changes of only −0.02 percent ($t = -0.55$) and −0.05 percent ($t = -1.20$) per month, respectively, for the earlier 1927 to 1963 subperiod. These results are once again consistent with those reported in Table VI, in that both risk changes and abnormal returns are shown to be confined to the time interval from 1964 to 1998.

A.2. Cross-sectional Relation Between Risk Changes and Abnormal Returns

The risk change results, in conjunction with the results of Table I (and others), thus far suggest that at the *aggregate* level, equity risk seems to decrease, while at the same time stock prices seem to increase during the three-year postannouncement period. In addition, we find no evidence of either risk changes or abnormal returns during the fourth and fifth years following the announcement. One possible explanation is that our sample is overpopulated with firms whose subsequent risk reductions are stronger than expected. In such case, we would expect to find a negative cross-sectional relation between *firm-specific* abnormal returns and contemporaneous risk changes. We next examine this cross-sectional relation.

We estimate *firm-specific* risk changes by multiplying the vector of firm-specific changes in the risk factor loadings estimated in equation (4) ($\beta_{\Delta i}$, $s_{\Delta i}$, $h_{\Delta i}$), by the vector of mean monthly realizations of the corresponding risk factors (mean[$Rm - Rf$], mean[SMB], mean[HML]). We group our firms into six risk change categories, based on the value of their individual risk changes. For each category, we measure the three-year postannouncement abnormal returns using the adjusted Fama and French methodology with both equal- and value-weighted portfolios. Table VIII presents the results. In the first four columns of the table, we categorize firms by risk change quartiles. Firms in the lower quartile have the largest decrease in risk. Firms in the upper quartile actually experience a risk increase. In the fifth and sixth columns, we divide our full sample according to the *sign* of the firm-specific risk change.

Our results provide strong evidence of a negative cross-sectional relation between postannouncement risk changes and contemporaneous abnormal returns. Indeed, firms in the lower risk change quartile (largest decrease in risk) have substantially larger adjusted intercepts when compared to those in the upper quartile (risk increase). Moreover, as we move across from the highest to the lowest quartile, the adjusted intercepts appear to be monotonically increasing. A similar observation holds for the comparison of adjusted intercepts associated with risk decreases and increases (in the fifth and sixth columns of the table).[16]

Perhaps some level of risk change should be anticipated on the announcement date. Healy and Palepu (1997) argue that "managers [could] initiate dividend payments when they forecast that their firms' earnings will be more stable relative to past earnings. In this case, investors will view dividend initiations as a signal of a decrease in the riskiness of the initiating firms" (p. 30). Their argument suggests that investors should expect the cost of equity to be somewhat lower during the postannouncement period. In this case, stock price responses during the event window should be negatively related to subsequently observed risk changes. In a cross-sectional analysis, we regress firm-specific risk changes on the *event-window* cumulative abnormal returns (CAR), for the post-1964 period (not shown).[17] The coefficient of CAR is -0.0106 ($t = -4.20$), suggesting that some of the subsequent risk changes may be anticipated on the day of the dividend announcement. However, the postannouncement results presented in Tables VII and VIII appear to indicate that not all of the postannouncement risk changes are immediately priced on day zero. One rational explanation is that our sample may, by chance, be overpopulated with firms that experience stronger than normal postannouncement risk reductions. The fact that neither the price drifts nor risk changes appear to reliably extend to the pre-1964 period is certainly consistent with this possibility.

B. Abnormal Returns and Unexpectedly Higher Cash Flows

There is another rational interpretation for the results in Tables VII and VIII. Fama and French (1997) show that the loadings of the SMB and HML factors are time varying and that the intertemporal variability in these loadings is partly related to concurrent changes in the size of the underlying firm. In their Table IV, Fama and French document a negative relation between firm size and each of the SMB and HML risk loadings.[18] This relation is further supported by evidence in their Appendix B, in which the slopes of the HML and SMB factors are shown to be negatively related to the inter-

[16] While the *monotonic* relation between risk changes and abnormal returns conforms to our main conjecture, we find it puzzling that the adjusted intercepts are not *significantly* negative in the case of postevent risk *increases*. For example, adjusted intercepts in the upper risk change quartile (largest risk increases) are either close to zero or moderately negative, and generally lack statistical significance. We would expect them to be *significantly* negative. This suggests that the cross-sectional variation in long-term abnormal returns is *only partially* explained by changes in postevent equity risk premia.

[17] The *event-window* abnormal returns are measured using the market-adjusted methodology proposed by Brown and Warner (1985). The event window is the three-day $[-1, +1]$ period centered on the event date. The CRSP equally weighted market index is used as the benchmark.

[18] In fact, Fama and French in their Table IV examine the relation between the HML loadings and the inverse of firm size. This relation is positive, indicating that the relation between HML loadings and firm size is negative.

Table VIII

Relation Between Postannouncement Abnormal Returns and Contemporaneous Risk Changes

The one- and three-year postannouncement abnormal returns are examined in relation to contemporaneous postannouncement changes in the riskiness of underlying stocks. For each event firm, we calculate the mean monthly risk change for the three-year postannouncement horizon, using the Fama–French model estimated over the $[m - 36, m + 36]$ period (where m is the announcement month). Event firms are then sorted into quartiles, the lowest (highest) quartile containing firms with largest risk decreases (increases). Alternatively, event firms are also sorted into risk increase/decrease groupings, based on the *sign* of the postannouncement risk change. Equal- and value-weighted calendar time portfolio returns are calculated each month from firms experiencing a dividend announcement in the previous 12 or 36 months, for each risk change quartile and risk increase/decrease category. For the *unadjusted* Fama and French (1993) three-factor model, the monthly excess returns of the calendar time portfolios, $R_{p,t} - R_{f,t}$, ($R_{f,t}$ is the return of the one-month T-bill) are regressed on (1) ($R_{m,t} - R_{f,t}$), the excess return of the CRSP market portfolio (equal- and value-weighted index for equal- and value-weighted regressions, respectively); (2) SMB_t, the difference in returns between portfolios of small and big stocks; and (3) HML_t, the difference in returns between portfolios of high and low book-to-market ratio stocks. For the *adjusted* Fama and French regression, a control firm calendar time portfolio (comprised of firms matched on size and preevent momentum) is subtracted from the event firm portfolio, and this difference is regressed on the three Fama and French factors. The regression intercepts provide an estimate of monthly abnormal performance. Estimated are both ordinary and weighted least squares (OLS and WLS) regressions. The WLS model weights each month with the square root of the number of firms for that month. The t-statistics are shown in brackets under each parameter. The WLS model t-statistics are calculated using the White (1980) method.

Postevent Horizon	Calendar Portfolio Weighting	Estimation Method	Risk Premium Chage Quartiles				All Risk Increases	All Risk Decreases
			Upper	Upper-middle	Lower-middle	Lower		
1 year	Equally	OLS	−0.00285	0.00053	0.00354	0.00688	−0.00043	0.00507
(no. obs. = 857 months)	weighted		[−0.64]	[0.17]	[1.27]	[2.03]**	[−0.17]	[2.56]***
		WLS	0.00111	0.00102	0.00385	0.01255	0.00124	0.00712
			[0.49]	[0.59]	[2.08]**	[5.67]***	[0.78]	[4.82]***
	Value	OLS	−0.00578	−0.00056	0.00257	0.00579	−0.00137	0.00309
	weighted		[−1.22]	[−0.16]	[0.80]	[1.57]	[−0.51]	[1.16]
		WLS	−0.00497	−0.00147	0.00302	0.01031	−0.00353	0.00466
			[−1.58]*	[−0.54]	[1.10]	[3.43]***	[−1.49]	[2.11]**
3 years	Equally	OLS	−0.00226	0.00043	0.00428	0.00627	−0.00058	0.00440
(no. obs. = 857 months)	weighted		[−0.83]	[0.22]	[2.63]***	[2.95]***	[−0.35]	[3.77]***
		WLS	−0.00140	0.00090	0.00340	0.00766	−0.00002	0.00449
			[−1.00]	[0.85]	[2.63]***	[5.06]***	[−0.02]	[4.40]***
	Value	OLS	−0.00311	−0.00165	0.00308	0.00240	−0.00151	0.00238
	weighted		[−1.25]	[−0.80]	[1.30]	[0.95]	[−0.83]	[1.43]
		WLS	−0.00368	−0.00167	0.00182	0.00364	−0.00301	0.00197
			[−1.84]*	[−0.93]	[0.97]	[1.92]*	[−1.88]*	[1.37]

***, **, * Significantly different from zero at the 1, 5, and 10 percent levels, respectively (two tail test) cept of the three-factor model. Their results imply that for any firm or industry, a period of decreasing risk loadings is likely to *coincide* with a period of unexpected positive abnormal performance: "The surprise onset of good times likely implies a decline in an industry's HML [and SMB] loading[s] and positive abnormal returns" (Fama and French (1997), p. 182). This abnormal performance occurs *in addition to* any price adjustment caused by concurrent declines in the cost of equity.

Under this alternative interpretation, our particular sample is overpopulated with firms that by chance become unexpectedly more profitable during the postannouncement period. Once again, the absence of any reliable price drift or risk change prior to 1964 provides clear support for this possibility.

V. Summary and Discussion

We show that the positive postdividend price drift documented by Michaely et al. (1995) is confined to the period from 1964 to 1998 and does not include stocks belonging to CRSP's largest size decile. During 1964 to 1998, stocks that initiate or resume dividends experience a significantly positive price drift only when the Fama–French calendar time portfolios are equally weighted. When the portfolios are value-weighted, the price drift becomes generally insignificant. This difference is due to firms in the largest size decile, for which no evidence of price drift is observed. Prior to 1964, neither the equal- or value-weighted price drifts are reliably different from zero.

We also document significant postannouncement reductions in the three Fama–French risk loadings and show that they are negatively related to the contemporaneous abnormal returns. This suggests that postannouncement stock prices are gradually responding to a combination of lower required rates of returns and favorable news about firm profitability. But if firms in our sample become less risky or more successful during the postannouncement period, why are these changes not being fully priced on day zero?

At first glance, one might be tempted to conclude that our results support the behavioral paradigm, in which investors suffer from cognitive biases that prevent Bayesian updating. Under this explanation, the postdividend changes in risk and profitability are predictable on day zero, but overconfident investors are slow to update their prior beliefs, leading to positive price drifts during the postannouncement period.

There is, however, another interpretation that is fully consistent with rationality: chance. If firms in our sample become unexpectedly less risky or more profitable, stock prices will increase after the dividend announcement, reflecting investors' rational reaction to the discovery of unexpected information.

Although our goal in this paper is not to discriminate between these two interpretations, we favor the rational explanation on grounds that the price drift is not robust across time periods and methodologies. Fama (1998) observes that "one way to test whether [an] anomaly is real or the sample-specific result of chance is to examine a different sample period" (p. 300). We show that both the positive price drifts and risk loading reductions are confined to the period from 1964 to 1998, suggesting that the post-1964 sample might by chance be overpopulated with firms that become unexpectedly less risky or more profitable after the dividend announcement.

Another argument in favor of the rational interpretation is that the positive price drift is confined to firms belonging to the nine smallest size deciles. Firms in the tenth decile have no abnormal returns. According to Fama (1998), this asymmetry is difficult to reconcile with the extant behavioral-based explanations, because "cognitive psychology, the basis of behavioral finance, does not seem to provide a basis for the common presumption that small stocks are more likely to be mispriced" (p. 296).

On a broader scale, our research contributes to the ongoing debate surrounding the validity of the Efficient Markets Hypothesis. Our findings provide empirical support for many of the arguments presented by Fama (1998) in his seminal critique of long-term event studies. We show that the dividend price drift is rather fragile, in that it seems to disappear when examined with value-weighted portfolios, and lacks robustness during the pre-1964 period. We also show that the magnitude of the price drift is significantly lower when the abnormal performance is measured with the adjusted Fama–French intercepts proposed by Mitchell and Stafford (2000), as opposed to the buy-and-hold methodology. More importantly, we argue that *chance* is a tenable rational explanation for the postdividend price drift identified by Michaely et al. (1995), and in doing so, we strengthen Fama's argument that the Efficient Markets Hypothesis cannot be refuted solely on the basis of the long-term anomalies that have thus far been documented in the extant literature.

Our paper also makes an important methodological contribution: We caution future researchers of long-term anomalies to be aware of situations in which the presumed price drift is in the opposite direction of any concurrent postannouncement changes in the Fama–French risk factor loadings. Such phenomena generally occur when firms in a particular sample are coincidentally experiencing changes in risk or profitability, independently of the corporate event being studied. If the adjusted Fama–French intercepts are positive during a period of decreasing risk factor loadings, the particular event sample may be overpopulated with firms that are by chance becoming safer or more profitable. Alternatively, a negative intercept that coincides with a contemporaneous increase in risk factor loadings is a good indication that sample firms may become unexpectedly riskier or less profitable. In either case, the *chance* explanation is quite plausible and must be carefully considered as a tenable alternative to the behavioral interpretation.

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