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When Is Bad News Really Bad News?

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ABSTRACT

We examine whether the price response to bad and good earnings shocks changes as the relative level of the market changes. The study is based on a complete sample of annual earnings announcements during the period 1988 to 1998. The relative level of the market is based on the difference between the current market P/E and the average market P/E over the prior 12 months. We find that the stock price response to negative earnings surprises increases as the relative level of the market rises. Furthermore, the difference between bad news and good news earnings response coefficients rises with the market.

ONE OF THE LONGEST RUNNING empirical debates in finance regards the relative pricing of "value" and "glamour" stocks. Beginning with early work by Basu (1983) and Stattman (1980), evidence has accumulated that excess returns on value stocks—that is, the issues of companies for which the ratio of earnings, cash flow, or book value per share is large relative to stock price—are greater than returns on glamour stocks, for which these ratios are small. On one side, Fama and French (1992, 1993, 1995, 1996) argue that the observed differential between the returns on value and glamour stocks represents a risk premium. The alternative view, articulated by Lakonishok, Shleifer, and Vishny (1994), is that the market fails to efficiently price value and glamour stocks.

Extending Lakonishok et al. (1994), recent work in behavioral finance by, for example, Barberis, Shleifer, and Vishny (BSV, 1998) and Daniel, Hirshleifer, and Subrahmanyam (1998) argues that the value/glamour effect is the result of investor psychology. In particular, the model in BSV allows for investor underreaction (in the intermediate term) to single shocks and investor overreaction (in the longer term) to a series of shocks. This model also implies an asymmetry in the returns to value and glamour stocks following a news shock. Following a string of positive shocks observed in, say, glamour stocks, the investor in this model expects another positive shock, that is, he expects the earnings to trend. If good news is announced, the market re-

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sponse is relatively small since the positive shock was anticipated. A negative shock, on the other hand, generates a large negative return, since it is more of a surprise.

The primary empirical tests of the competing explanations for the value/glamour differential have been conducted on earnings announcements (La Porta, (1996), Dechow and Sloan (1997)). La Porta et al. (1997) and Bernard, Thomas, and Wahlen (1997) find that earnings announcement returns explain almost half of the return differential between value and glamour stocks. More recently, Skinner and Sloan (1999) use a sample of earnings announcements and find that when pre-announcement effects are included, the differential reaction to earnings announcements *completely* explains the differential returns to value and glamour stocks. In addition, Skinner and Sloan also find evidence consistent with the BSV hypothesis. In particular, they find that the response to news is asymmetric for value and glamour stocks; the market reacts more strongly to bad news for both types of firms, but the reaction to bad news for glamour stocks over the subsequent 20 quarters is much larger.

In the BSV model, the source of uncertainty is the model of earnings for a particular firm and, hence, is firm specific. Thus, this model can be used to explain a cross-sectional puzzle: Why do value stocks (or, more generally, stocks which have underperformed in the past) appear to outperform glamour stocks (or stocks which have outperformed in the past) over time? More recently, however, there is some anecdotal evidence that market-wide "glamour" effects also are possible. For example, in the November 12, 1996, *Wall Street Journal*, Deborah Lohse (1996, p. C1) speculates that the asymmetrical response of stock prices to good and bad news is related to the level of the market:

> Analysts say that stocks that surprise analysts with better-than-expected earnings are often rewarded with a ho-hum increase if any. However, the market is punishing stocks even more than usual for earnings disappointments. . . . Part of the problem is fear of the valuation levels that many stocks have reached. *With the market at these levels, if stocks are slightly down (in terms of unexpected earnings), they get severely punished*. [emphasis added]

The statement in this article suggests that there are systematic shifts in investor sentiment that are common across stocks; specifically, during good times, investor confidence rises and investors extrapolate good news for companies generally. However, firms providing specific information that the extrapolation of good news is not applicable to them are severely punished. During bad times, the reverse reaction occurs.

The notion that the market responds more strongly to bad news in good times does not necessarily require the assumption of irrationality or overreaction on the part of investors that underlies much of the value/glamour literature. For example, regime-switching models, such as those developed by David (1997) and Veronesi (1999), offer a rational explanation for why the aggregate market (although not necessarily individual stocks) can respond more strongly to bad news than good news in good times. In these models, investors are uncertain about the overall state of the market. Because investors cannot observe the current state of the market directly, they must infer it from past market performance. Following a long period of superior market performance, investors will become highly confident the market is in a good state. Under such circumstances, further good news has little impact on investor beliefs. However, bad news causes market prices to fall for two reasons. First, bad news causes investors to infer a lower probability that the market is in the good state. Second, as uncertainty in the state of the economy increases, risk-averse investors require a higher expected rate of return to hold stocks, and the market discount rate rises.

The uncertainty about the state of the economy causes an asymmetry in the response to good news and bad news. That is, when investors believe that the economy is in a bad state and good news arrives, the inferred probability that the market is in a good state increases; thus, the positive impact on prices is offset by the rising discount rate generated by increased investor uncertainty.

The regime-switching models discussed above are designed to describe aggregate market phenomena, not firm-specific responses. For instance, the discount rate effect in the regime-shifting models operates through the market risk premium. Therefore, it will be affected by market-wide information. Individual firm announcements may or may not provide market-wide information, although, in principle, such models could be extended to permit firm-specific reactions.[1]

In this paper, we examine whether the strength of firm-specific responses to new information is affected by the aggregate level of the market. The interaction between aggregate conditions and firm-specific response to news has received comparatively little attention in the literature. Theoretical models have tended to focus on stock price behavior in response to either aggregate or firm-specific information. Thus, the behavioral model suggests that investors' expectations are path dependent, but only past firm-specific information is employed in the formation of expectations, while aggregate market conditions are ignored. Conversely, in the regime-shifting models, investors' expectations are dependent on market-wide information; the response to firm-specific news is not considered. However, if the premise of the *Wall Street Journal* quote is descriptively valid, then more sophisticated models must be developed because the level of the market affects the reaction of individual firms. Consequently, the current research is not a test of current competing theories but an examination of whether or not both types of theory need to be extended.

[1] Ribeiro and Veronesi (2001) generalize the Veronesi (1999) model for the market portfolio to the case of individual firms.

In the foregoing discussion, the notion of what is meant by good times or a high market is not precisely defined. Clearly, as a result of both inflation and real economic growth, the level of the market must be measured relative to some benchmark such as earnings or dividends. In this paper, forecasts of future earnings are used to benchmark the level of prices.

Selecting a benchmark does not resolve all the ambiguity related to the level of the market. When the theoretical models or writers in the popular financial press refer to good times does that mean good times in the absolute sense or relative to recent experience? More specifically, should the absolute price-to-forecasted earnings (hereafter P/E) ratio be used to define the level of the market or should the level be defined by the relation between the current P/E ratio and the ratio observed in the recent past? In this paper, the latter, relative, definition of the level of the market is employed. In particular, the level of the market is measured by comparing the value-weighted average market P/E ratio at the end of each firm-announcement month to the monthly average market P/E observed during the previous 12 months. Thus, a high P/E month is one in which the market's valuation of future forecasted earnings, expressed through price, is higher than the valuation of future earnings has been during the last year.

The "relative" definition of the level of the market is chosen over the "absolute" definition for several reasons. To begin, it is more consistent with the behavioral models that are based on the extrapolation of recent shocks.[2] More importantly, because returns were so strongly positive over our entire sample period, it is likely that average forecast errors embedded in prices are also positive. Thus, investors were repeatedly surprised on the upside. Consistent with this, an ARIMA analysis of the price/earnings ratio during the sample period of our study reveals that the P/E ratio is closely approximated by a random walk with positive drift. See Figure 1, which plots the monthly market P/E, the sample average P/E, and 12-month moving average (rolling) P/E ratio.[3] Thus, changes in the P/E ratio *during the sample period* are permanent. Alternatively stated, forming portfolios on the basis of the absolute level of P/E becomes equivalent, in this sample period, to forming them on the basis of time. The low P/E portfolios would consist almost exclusively of observations from the late 1980s and early 1990s, while the high P/E portfolio would include almost exclusively observations from 1997 and 1998. Because our sample period does not include a significant period of price decline, the use of a relative rather than absolute measure of market performance is required to generate dis-

[2] For example, in Barberis et al. (1998, p. 318), model 2, they note that earnings shocks "are more likely to be followed by another shock of the same sign." Thus, they conclude, that in this setting, investors "extrapolate past performance too far into the future."

[3] After differencing the monthly market P/E series, a test of the residuals to lag 12 does not reject the null hypothesis of white noise (p-value = 0.929).

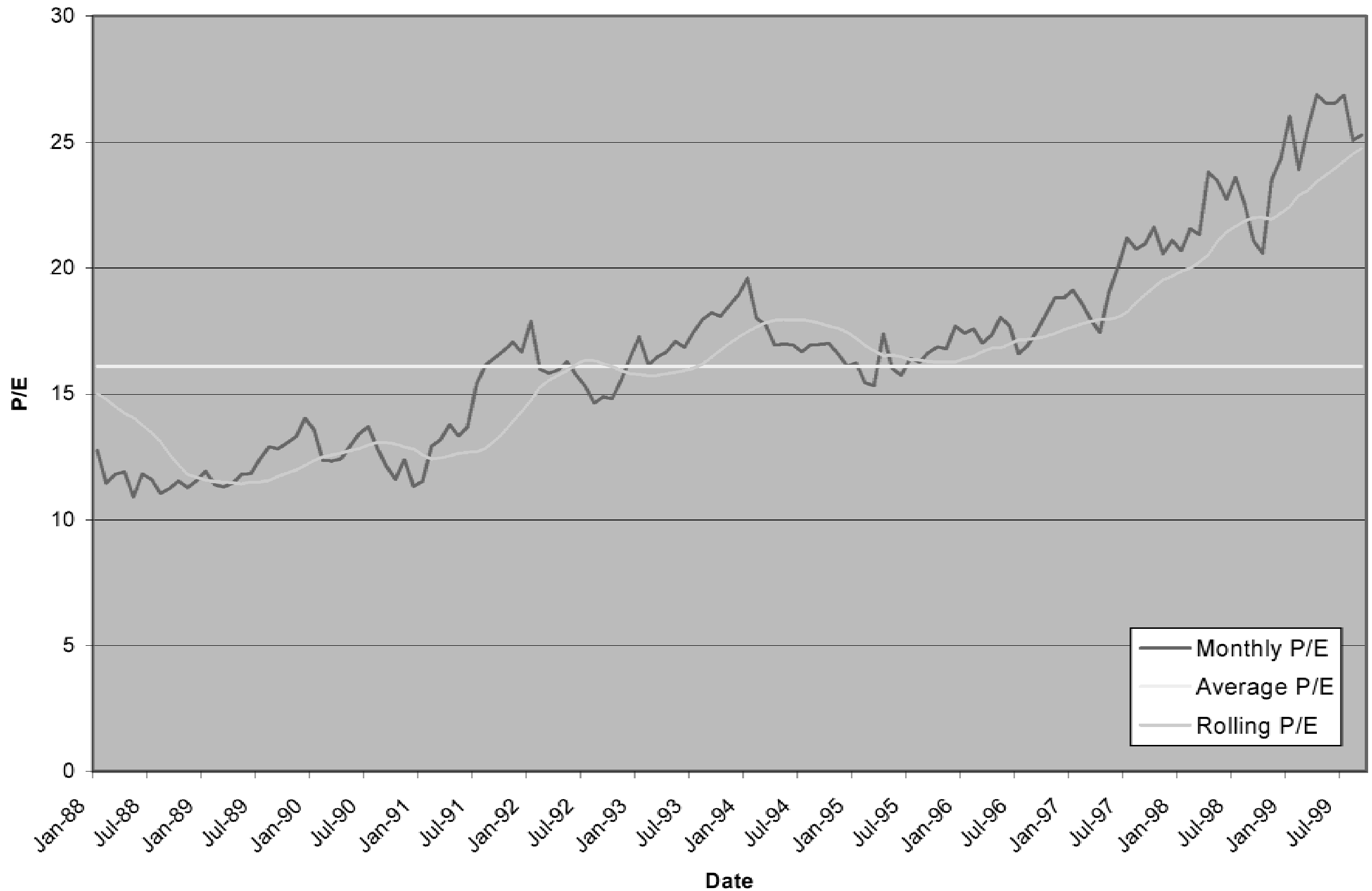

Figure 1. Monthly market P/E.

persion in market level that is not simply time.[4] As a practical matter, this intuition is confirmed by our empirical findings. If the empirical analysis is repeated using the absolute P/E instead of the relative P/E, the results (not reported) are largely insignificant and the high P/E portfolios are comprised almost exclusively of those in the latter years of our sample.

In the case of the regime-shifting models, the investor's uncertainty about the market state affects the discount rate applied to future cash flows. As uncertainty about the market state changes, discount rates vary through time. Our relative price-to-forecasted earnings measure captures changes in the discount rates applied to future earnings; changes in this ratio should reflect variations in discount rates over time.[5]

To explore whether the reaction of stock prices to earnings shocks is affected by the level of the market, we examine whether the earnings response coefficients for good and bad earnings shocks change as the relative level of the market changes. The study is based on a sample of 24,097 announcements of firms' annual earnings during the period 1988 to 1998. The level of the market is defined by the variable, *DIFFPE*, which is the difference between the market P/E ratio during the announcement month and the average market P/E during the preceding 12 months; note that the denominator is measured using next year's forecasted earnings. If the premise in the *Wall Street Journal* is valid, then stock prices should respond more strongly to negative news as relative market valuations increase (or as *DIFFPE* rises). The findings generally support this hypothesis. In particular, the stock price response to negative earnings surprises is increasing as *DIFFPE* rises. The results for good news are more ambiguous, although findings are weakly consistent with the notion that the stock price response to positive earnings surprises is decreasing as *DIFFPE* rises. Moreover, the difference between bad news and good news response coefficients is increasing across the *DIFFPE* portfolios. Findings based on subsamples of Nasdaq and NYSE data are broadly supportive of the general results. For both classes of firms, stock prices respond most strongly to bad news in good times. However, there are some differences between the two subsamples that are discussed later in the paper.

[4] We also considered forming portfolios using two other measures based on the recent past—earnings growth and stock return over the past year. However, each has the same two drawbacks. First, earnings growth and stock returns only capture separately changes in the denominator and numerator of market P/E, which reflects the market's valuation of projected future earnings. Second, perhaps as a result of this problem, neither does a good job empirically in distinguishing the pricing effects we identify based on changes in market P/E.

[5] Changes in our measure should also reflect changes in anticipated growth rates in cash flows. This can readily be seen by referring to the familiar stylized representation of value (price) given by a Gordon dividend growth model: $P = D/(r - g)$, where P is stock price, D is first period dividend, r is the risk-adjusted discount rate, and g is an assumed constant growth in dividends. If the dividends, D, are a constant proportion, K^*, of permanent economic earnings, the price/earnings ratio is given by $P/E^* = K^*/(r - g)$.

The behavioral models, as currently developed, suggest that the pricing effect we evaluate may be more pronounced for glamour than for value stocks. To see whether there is an interaction between our time series analysis and the cross-sectional findings cited above, we partition our sample into quintiles based on each firm's ratio of price to earnings in the month preceding announcement of earnings. We then examine whether changes in the asymmetry between good and bad news over time is different for glamour and value stocks. The findings fail to indicate any significant difference in the behavior of value and glamour stocks in our sample.

The paper proceeds as follows. Section I describes the research design and explains our hypotheses. Section II describes the data and sample. Section III provides the paper's findings, and Section IV summarizes and concludes the study.

I. Research Design

The discussion in the preceding section suggests examining two hypotheses. The first is that the market responds asymmetrically to unexpected good and bad earnings news in good and bad states. Bernard et al. (1997), La Porta et al. (1997), and Skinner and Sloan (1999) document substantial differences in the responses to good and bad earnings shocks. However, these studies examine exclusively mean returns for (longer-run) periods subsequent to announcement of earnings. In contrast, we focus on the elasticity of stock prices to earnings surprises at announcement, that is, earnings response coefficients. Specifically, we estimate regressions that allow for a differential response to positive and negative earnings in the overall sample. In this regard, we create two indicator variables. The first, *UP*, is set equal to one if the shock is positive and zero otherwise; the second, *DOWN*, is equal to one if the shock is negative and zero otherwise. We then estimate two regressions of the form

$$RET_{it} = a_0 + a_1 UE_{it} + a_2 SIZE_{it} + \epsilon_{it}, \tag{1}$$

and

$$RET_{it} = b_0 + b_1 UEUP_{it} + b_2 UEDOWN_{it} + b_3 SIZE_{it} + \epsilon_{it}. \tag{2}$$

In equations (1) and (2), RET_{it} is the excess return on firm i during the earnings announcement period as defined below, and UE_{it} is the unexpected earnings for firm i at time t computed using the IBES data to measure expectations. In equation (2), *UEUP* (*UEDOWN*) is the product of unexpected earnings, *UE*, and the indicator variable, *UP* (*DOWN*); this permits the slope on *UE* to be different, conditional on the sign of the earnings surprise. The variable *SIZE* is the natural log of equity market value (in thousands) in the period prior to the firm's earnings announcement. The size proxy is included to control for risk differences not already reflected in excess return (Fama and French (1992, 1993)) and for potential scale differences (Barth and Kallapur (1996)). If the market response to good and bad earnings innovations is asymmetric, then equation (2) should have significantly more explanatory power than equation (1) and the hypothesis that $b_1 = b_2$ should be rejected.

The second, and more innovative, hypothesis is that the degree of asymmetry depends on the level of the market. To investigate this hypothesis, we use the ratio of price to forecasted future earnings to measure the relative level of the market. More specifically, using *DIFFPE*, the sample of earnings announcements is divided into quintiles. The first quintile contains the earnings announcements that occurred when *DIFFPE* is the lowest. The other quintiles contain earnings announcements that occurred when *DIFFPE* is progressively larger. If the announcement asymmetry depends on the level of the market, then the difference between b_1 and b_2 should change as equation (2) is estimated for quintiles defined by progressively greater levels of *DIFFPE*.

Finally, it is possible that the time-series hypotheses examined here interact with the cross-sectional value/glamour effects documented in the literature reviewed in the previous section. More specifically, the asymmetry between the response to good and bad news may vary both over time and across stocks, with the relative sensitivity to bad news being more pronounced for glamour stocks in good times and less pronounced for value stocks in bad times. To test for this interaction, we sort individual companies into quintiles based on each firm's P/E ratio in the month preceding its earnings announcement, where the ratio is based on next year's forecasted earnings. The companies in the lowest quintile are the "value" stocks and the companies in the highest quintile are the "glamour" stocks. Using these definitions, we test whether the asymmetry varies cross-sectionally as well as over time.

II. Data

Our sample period of individual firms' earnings announcements extends from 1988 through 1998. Consensus earnings forecasts, realized earnings, and earnings report dates are collected from IBES. For each earnings announcement, we define an event window which extends from day −20 through 0 relative to the earnings report date. This window is divided into a pre-announcement period, extending from day −20 through day −6, and an announcement period, extending from day −5 through day 0.[6] These earnings data are subsequently matched with price, shares, and returns data from

[6] The study was repeated using a shorter window, day −1 to day 0. The two-day window ameliorates the problem of residual correlation in the regressions by reducing the number of overlapping observations. However, the two-day window may fail to account for pre-announcement leakage of information. The results using the shorter window are similar to those reported. The study also was replicated using a subsample including only nonoverlapping earnings announcement dates. Although this significantly reduces the sample size, it does not fundamentally alter the results.

the Center for Research in Security Prices (CRSP). For each firm and each report date, raw returns are summed across the event window; excess returns are calculated as the sum of the firm's raw returns during the event window less the sum of the CRSP value-weighted market return over the same period.[7] A firm must have a price available on the earnings report date to be included in the sample. To analyze the price impact of an earnings shock in the announcement period, unexpected earnings, *UE*, (or earnings shocks) are calculated as

$$UE = (\text{actual earnings} - \text{consensus forecast earnings})/\text{price}(-6), \tag{3}$$

where price(−6) is the share price six days before the earnings announcement (or one day prior to the announcement window).[8]

To reduce the impact of outliers, we delete observations for which the earnings shock is greater than (less than) 0.5 (−0.5). To minimize the effect of market frictions (see, e.g., Ball, Kothari, and Shanken (1993)), observations with price(−6) (price (−20)) in the announcement (pre-announcement) window of less than $5.00 are deleted. To remove the impact of stock splits or stock dividends in the event window, we delete observations where the number of shares outstanding 20 days prior to the earnings announcement differs from shares outstanding on the announcement date. Consistent with prior research showing earnings response coefficients are essentially zero for firms reporting negative earnings (e.g., Hayn (1995) and Lipe, Bryant, and Widener (1998)), observations where firms report negative earnings are deleted. Finally, observations for which the ratio of actual earnings to market capitalization on day zero exceeds one are also deleted.

To categorize whether the earnings announcement occurred in a "high" or "low" valuation state, a monthly time series of market price/earnings ratios are estimated using the IBES data. To calculate the market P/E for a particular month t, we first collect the consensus earnings forecast for the next fiscal year made in month t for each firm, as well as the observed price in month t for that firm. We construct a value-weighted average of the IBES forecast earnings-to-price ratio across firms, then take the reciprocal of this number as the market price/earnings ratio. Thus, we calculate

$$P/E(mkt)_t = 1\Big/\left[\sum_{i=\{1,Nt\}} w_{it}(\mathrm{E}_t[EPS_{i\tau}]/P_{it})\right], \tag{4}$$

[7] As a specification check, we also examine excess compounded returns, which minimize bid-ask bias; these excess returns are calculated as the compounded raw return in the announcement (pre-announcement) period, less the compounded market return over the same interval. As discussed in Section IV.C below, findings are essentially identical to those reported.

[8] As an additional specification check, we examine the period extending from day −20 to day −6 to test for pre-announcement leakage, particularly for bad news (Skinner and Sloan, 1999). For these specifications, earnings shocks are standardized by the price on day −20.

where w_{it} is the value of firm i relative to the total market value of firms available in the sample for month t, P_{it} is the share price of firm i in month t, and $\mathrm{E}_t(EPS_{i\tau})$ is the consensus analysts' forecast in month t for annual earnings reported in month τ. Firms are deleted from the average if they do not have price, forecasted earnings, or shares outstanding numbers available; only earnings forecasts less than one year old are considered when constructing this average (i.e., τ must be no more than 12 months distant from t).[9]

Expected future earnings are used in the P/E calculation, because the resulting ratio better measures the market's valuation of future expected cash flows than does a ratio constructed with current earnings. Furthermore, it insulates the measure from transitory shocks to earnings that can cause fluctuations in P/E ratios based on trailing earnings that are difficult to interpret.[10]

After the time series of market price/earnings ratios is constructed, *DIFFPE*, the difference between each month's market price/earnings ratio and the average of the market's monthly price/earnings ratio over the previous 12 months' period is calculated. Earnings announcements are grouped into one of five portfolios based on the value of this difference as of the announcement month. The mean *DIFFPE* for the five portfolios are −1.45, 0.01, 0.62, 1.09 and 1.99. Note that although the sample period of earnings announcements extends from 1988 through 1998, because *DIFFPE* is based on the difference between the announcement month's market P/E and the average market P/E over the prior 12 months, the $P/E(mkt)_t$ series extends back an additional year to 1987.

III. Findings

A. Descriptive Statistics

Table I, Panel A, presents sample descriptive statistics for earnings surprises, UE_{it}, and earnings response coefficients and adjusted R^2 values from a regression of excess returns, RET_{it}, and UE_{it} in equation (1). Consistent with prior research indicating analysts are optimistically biased (O'Brien (1988) and Kang, O'Brien, and Sivaramakrishna (1994)), mean *UE* is negative. However, when we examine average earnings shocks by year, we find that the largest negative shocks occur, on average, in the early part of our

[9] Consensus analysts forecasts in each month t are constructed by IBES on the Thursday following the third Friday of each month (i.e., it is always between the 14th and 20th). For the price appearing in the denominator of equation (4), we use the IBES reported price for each firm i in month t. By construction, IBES selects a price that is essentially contemporaneous with the consensus forecast date. Thus, each firm's earnings–price ratio is calculated using the most recent consensus forecast available and for each firm, a matching price.

[10] Since prices represent claims to future cash flows, forecasted earnings are likely a better proxy for future cash flows than past earnings or cash flow measures. In particular, earnings forecasts presumably are based on a richer information set than past earnings series.

Table I

Unexpected Earnings Summary Statistics and Regression Statistics

The sample period is annual earnings for 1988 to 1998. *UE* is calculated as (actual earnings − consensus forecast earnings)/price(−6), where price(−6) is the share price six days before the earnings announcement (day 0); *RET* is the six day excess return, from day −5 to day 0, calculated as the sum of a firm's raw returns, less the sum of the CRSP value-weighted market return over the same two-day period.

Variable		Mean	Std. Dev.	*N*
		Panel A: Full Sample		
UE		−0.00060	0.02093	24,097
RET		0.00590	0.07061	24,097
		$RET = a_0 + a_1 UE + a_2 SIZE + e$		
a_0	0.02716			
	(7.76)			
a_1	0.27847			
	(12.76)			
a_2	−0.00169			
	(−6.08)			
Adj. R^2	0.0080			
		Panel B: Sample Partitioned by Sign of Unexpected Earnings		
$UE > 0$				
UE		0.0065	0.0200	12,448
RET		0.0154	0.0678	12,448
$UE = 0$				
RET		0.0058	0.0706	1,687
$UE < 0$				
UE		−0.0096	0.0205	9,962
RET		−0.0060	0.0723	9,962
		$RET = b_0 + b_1 UEUP + b_2 UEDOWN + b_3 SIZE + E$		
b_0	0.0286			
	(7.94)			
b_1	0.241			
	(7.74)			
b_2	0.321			
	(9.60)			
b_3	−0.002			
	(−6.29)			
Adj. R^2	0.0080			

sample period. For example, three out of the seven years between 1992 and 1998 have positive average shocks, with the 1997 mean *UE* significantly positive and averaging 0.05 percent of the stock price. The sample average earnings response coefficient of 0.278 is significant and also is comparable to that found in prior research (Brown et al. (1987)).[11] Finally, mean excess return, *RET*, is approximately zero, which is expected, given that mean *UE* is economically small.

Table I, Panel B, presents findings analogous to those in Panel A, but broken down by sign of earnings surprise. Mean positive unexpected earnings, *UEUP*, and negative unexpected earnings, *UEDOWN*, are of similar magnitude, 0.0065 and −0.0096. Untabulated findings indicate no major trends in their magnitude differences, although the number of earnings shocks in high (low) market valuation states is generally rising (declining) during the sample period.

The earnings response coefficients corresponding to positive and negative earnings shocks, b_1 and b_2 from equation (2), are 0.241 and 0.321, and they are marginally significantly different (p-value = 0.09), indicating a weak asymmetry in the market's response to good and bad news. Consistent with this finding, the adjusted R^2 of 0.80 percent is virtually identical to that in equation (1) in which b_1 and b_2 are constrained to be equal. This finding is only weakly consistent with Skinner and Sloan (1999), who find an asymmetry in mean security returns for positive and negative earnings surprises, unconditional on the magnitude of the earnings surprise. Finally, mean excess return, *RET*, is 0.0154, 0.0058, and −0.0060 for positive, zero, and negative unexpected earnings subsamples, indicating that the mean return difference for negative and zero *UE* firms, 1.18 percent, is larger than that for the positive and zero *UE* firms, 0.96 percent.

Before turning to findings from estimations of equations (1) and (2) for the separate market P/E portfolios, it is useful to extend the earlier analysis on the time series characteristics of the market during our sample period. Table II provides a description of the level of the market over time as measured by our forward-looking weighted average price/earnings multiple. Panel A presents statistics for market P/E as well as *DIFFPE*, the difference between market valuations prevailing during firm-announcement months and the average market P/E during the prior 12 months. Mean annual P/E is just the average for the 12 calendar months. Although mean P/E is generally rising during the 1990s, it clearly is not monotonic, falling from an average of 17.80 in 1994 to 15.97 in 1995. Mean market P/E values rise to their highest levels in 1997 and 1998, reflecting the booming stock prices during this period.

Mean annual *DIFFPE* is computed as the mean *DIFFPE* for the sample firm announcements during a given year, *not* the simple monthly mean difference in market P/E and average market PE for the prior 12 months. In contrast to P/E, mean *DIFFPE* exhibits no apparent secular trend. This is because *DIFFPE* is a relative measure reflecting the difference between the level of the market when earnings are announced and the recent (12-month) historical level of the market. However, as with P/E, the maximum mean annual *DIFFPE* also appears in 1998, when the market is at historically high levels.

[11] Throughout the paper, we use a five percent criterion for assessing statistical significance.

Table II

Summary Statistics for Price-Earnings Ratios

Market *PE* statistics are based on end-of-month values. The variable *DIFFPE* is the difference between each month's *PE* and the average of the market's monthly *PE* over the previous 12 months' period. *DIFFPE* statistics are computed using *DIFFPE*s prevailing during each firm-announcement month.

Panel A: Market Price-to-Forecasted Earnings, *PE*, and *DIFFPE*, Summary Statistics

	Mean	Std. Dev.	Max.	Min.	*N*
			Market PE		
1988	11.599	0.464	12.740	10.913	1021
1989	11.961	0.767	14.037	11.295	1462
1990	12.680	0.605	13.700	11.327	1548
1991	13.800	1.744	17.044	11.527	1592
1992	16.020	0.863	17.883	14.631	1806
1993	16.992	0.889	18.950	16.142	2017
1994	17.805	0.956	19.595	16.097	2472
1995	15.967	0.618	17.695	15.332	2894
1996	17.560	0.499	18.808	16.592	2960
1997	19.193	1.114	21.614	17.437	3162
1998	21.762	0.976	24.342	20.585	3163
All years					
Mean	16.842				
Std. dev.	3.064				
Max.	24.342				
Min.	10.913				
			DIFFPE		
1988	−2.450	0.966	−0.129	−3.333	1021
1989	0.377	0.624	1.890	−0.205	1462
1990	0.075	0.699	1.223	−1.483	1548
1991	1.074	1.373	3.315	−1.058	1592
1992	0.481	1.295	3.149	−1.697	1806
1993	0.991	0.524	1.776	0.420	2017
1994	0.105	1.058	2.123	−1.347	2472
1995	−0.802	0.819	1.431	−1.482	2894
1996	0.843	0.466	1.417	−0.551	2960
1997	1.099	0.754	2.937	−0.519	3162
1998	1.288	0.797	3.282	−1.422	3163
All years					
Mean	0.445				
Std. dev.	1.237				
Max.	3.315				
Min.	−3.333				

Panel B: Distribution of Firm-Announcements by Year and *DIFFPE* Portfolio

			DIFFPE		
	1	2	3	4	5
1988	919	102	0	0	0
1989	0	804	246	363	49
1990	125	1,014	104	305	0
1991	217	0	835	93	447
1992	318	411	836	0	241
1993	0	0	988	307	722
1994	870	1,202	0	0	400
1995	2,162	247	320	86	79
1996	93	546	177	2,144	0
1997	0	511	1,102	400	1,149
1998	156	0	165	1,111	1,731

Table III

Descriptive Statistics for Unexpected Earnings

Unexpected Earnings, *UE*, sample summary statistics by *DIFFPE* Portfolio. *DIFFPE* is the difference between each month's *PE* and the average of the market's monthly *PE* over the previous 12 months' period. *UE* = (actual earnings − consensus forecast earnings)/price(−6), where price(−6) is the share price six days before the earnings announcement (day 0).

DIFFPE	Mean	Std. Dev.	Min.	Max.	*N*
		Panel A: Overall Sample			
1	−0.0012	0.0212	−0.343	0.400	4,860
2	−0.0010	0.0245	−0.349	0.356	4,837
3	−0.0008	0.0211	−0.290	0.487	4,773
4	0.0002	0.0176	−0.173	0.467	4,809
5	0.0001	0.0197	−0.381	0.480	4,818
		Panel B: Sample Partitioned by Sign of Unexpected Earnings			
$UE > 0$					
1	0.0062	0.0185	0.0000	0.4	2,480
2	0.0082	0.0235	0.0000	0.356	2,351
3	0.0067	0.0198	0.0000	0.487	2,407
4	0.0058	0.0175	0.0000	0.467	2,573
5	0.0058	0.0200	0.0000	0.48	2,637
$UE < 0$					
1	−0.0103	0.0223	−0.343	−0.0000	2,051
2	−0.0110	0.0231	−0.349	−0.0000	2,203
3	−0.0098	0.0209	−0.290	−0.0000	2,006
4	−0.0083	0.0160	−0.173	−0.0000	1,878
5	−0.0081	0.0182	−0.381	−0.0000	1,824

The annual standard deviation in *DIFFPE* ranges from 0.47 in 1996 to 1.37 in 1991, indicating there is within-year variation in *DIFFPE*. Stated another way, firm-announcements from particular years will not necessarily all be placed within one specific *DIFFPE* group, of which there are five. Table II, Panel B, lists the number of firms within each *DIFFPE* portfolio for each sample year. High *DIFFPE* firm-announcements (*DIFFPE* = 4 and 5) are bunched disproportionately in 1998, when the market is at its historically high level. However, in the prior year, 1997, when the market PE is second only to 1998, there are nearly as many middle-range *DIFFPE* firm-announcements as there are high ones. With the exception of the two lowest *DIFFPE* categories in 1993 and the highest *DIFFPE* categories in 1988, firm-announcements span the *DIFFPE* categories in each sample year.

Table III, Panel A, presents unexpected earnings statistics for the five *DIFFPE* portfolios for the full sample. Panel B presents analogous statistics, but broken down by the sign of unexpected earnings. Panel A reveals a monotonic increase in earnings shocks across the five portfolios, ranging from a mean of −0.0012 in *DIFFPE* portfolio 1 to 0.0001 for portfolio 5. Panel B reveals no apparent pattern in the magnitude of positive and negative shocks across the five portfolios

Table IV

Full Sample Findings: Average Coefficients, Associated *t*-statistics, and Adjusted R^2s for Return Regressions for Annual Earnings Announcements, with a Sample Period Covering Annual Earnings for 1988 to 1998.

The variable *UE* is calculated as (actual earnings − consensus forecast earnings)/price(−6), where price(−6) is the share price six days before the earnings announcement (day 0); *RET* is the six day excess return, from day −5 to day 0, calculated as the sum of a firm's raw returns, less the sum of the CRSP value-weighted market return over the same two-day period. The variable *SIZE* represents the natural log of equity market value in the period prior to the firm's earnings announcement, generally day −6.

Panel A: $RET_{it} = a_0 + a_1 UE_{it} + a_2 SIZE_{it} + \epsilon_{it}$

DIFFPE	*UE*	*SIZE*	Adj. R^2	NOBS
1	0.245	−0.002	0.74	4,857
	(5.36)	(−3.34)		
2	0.183	0.000	0.39	4,828
	(4.54)	(0.14)		
3	0.377	−0.003	1.35	4,766
	(7.35)	(−3.88)		
4	0.236	−0.001	0.41	4,802
	(4.15)	(−2.15)		
5	0.383	−0.003	1.33	4,813
	(7.02)	(−4.10)		

Panel B: $RET_{it} = b_0 + b_1 UEUP_{it} + b_2 UEDOWN_{it} + b_3 SIZE_{it} + \epsilon_{it}$

DIFFPE	*UEUP*	*UEDOWN*	*SIZE*	Adj. R^2	NOBS
1	0.244	0.245	−0.002	0.72	4,857
	(3.40)	(3.86)	(−3.30)		
2	0.133	0.237	−0.000	0.39	4,828
	(2.27)	(3.89)	(−0.07)		
3	0.387	0.366	−0.003	1.33	4,766
	(5.15)	(4.80)	(−3.37)		
4	0.168	0.335	−0.001	0.42	4,802
	(2.17)	(3.56)	(−2.38)		
5	0.285	0.576	−0.003	1.42	4,813
	(4.13)	(5.79)	(−4.48)		

B. Regression Results for DIFFPE Portfolios

We now turn to the primary hypothesis of our study, that the difference between earnings response coefficients, b_1 and b_2 in equation (2), should increase with the valuation level of the market. Table IV, Panels A and B, presents the regression summary statistics corresponding to equations (1) and (2) for each of the *DIFFPE* portfolios.[12] The findings for equation (1) in Panel A, in which b_1 and b_2 are constrained to be equal, indicate no apparent trend in the *UE* response coefficients. The highest earnings response coefficient estimates obtain for middle and highest *DIFFPE* portfolios. All response coefficients are significantly positive.

The evidence in Table IV, Panel B, reveals a striking difference in the pattern of response coefficients for positive and negative earnings surprises. For ease of comparison, the *UEUP* and *UEDOWN* coefficients across the five *DIFFPE* portfolios are plotted in Figure 2, and their differences are plotted in Figure 3. Looking first at the negative earnings surprises, the coefficients on *UEDOWN*, b_2, are all significantly positive. More importantly, the response coefficients exhibit a generally increasing pattern across the five *DIFFPE* groupings. A regression of the *UEDOWN* coefficient on *DIFFPE* indicates the slope is significantly positive (p-value = 0.026 under a one-sided alternative).[13] This finding suggests that the market reacts more strongly to bad news as market levels rise and is consistent with predictions of both the regime-shifting models and behavioral models.

In contrast to the negative earnings surprises, the *UEUP* response coefficients, b_1, exhibit no apparent trend. As depicted in Figure 2, the pattern of b_1 values exhibits a W-shape, beginning at 0.24 for the lowest *DIFFPE* portfolio, rising to a peak of 0.39 for the *DIFFPE* = 3 portfolio, and declining to 0.29 for the highest *DIFFPE* portfolio. A regression of the *UEUP* coefficient on *DIFFPE* produces a slope that is insignificantly different from zero. In contrast with the results for bad news, this finding indicates that the stocks respond similarly to good earnings shocks at all relative market levels.

One unique feature of the regime-shifting models relates to the expected pattern in the *difference* in market responses to bad news and good news. In up markets (i.e., when *DIFFPE* = 5), bad news conveys both negative future cash flow implications as well as an increase in uncertainty and an increase in the discount rate; good news conveys little information regarding either cash flow or uncertainty. In down markets (i.e., when *DIFFPE* = 1), good news conveys positive future cash flow implications. However, the positive reaction is dampened by an increase in uncertainty and the discount rate. Bad news conveys little information regarding either cash flow or uncertainty. Although these are

[12] The market value information is calculated at day −6; if this value is missing, we use the day −7 value, then day −8 and so on. If the value is still missing at day −10, we use the value at day −1.

[13] An alternative method for assessing whether the coefficient on earnings shocks is higher in up markets relative to down markets is to examine only the coefficient differences in two most extreme *DIFFPE* portfolios. For example, assuming coefficient from different *DIFFPE* portfolios are uncorrelated, the t-statistic for the difference in *UEDOWN* coefficients for *DIFFPE* = 5 and *DIFFPE* = 1 is 2.80 ($[0.576 - 0.245]/\mathrm{sqrt}(0.0635^2 + 0.0995^2)$). An advantage of this approach is that it incorporates coefficient dispersion in the test of significance. A disadvantage is that it ignores the trend in coefficients for intermediate *DIFFPE* portfolios. Nonetheless, both approaches yield qualitatively similar inferences for all regressions reported in the paper.

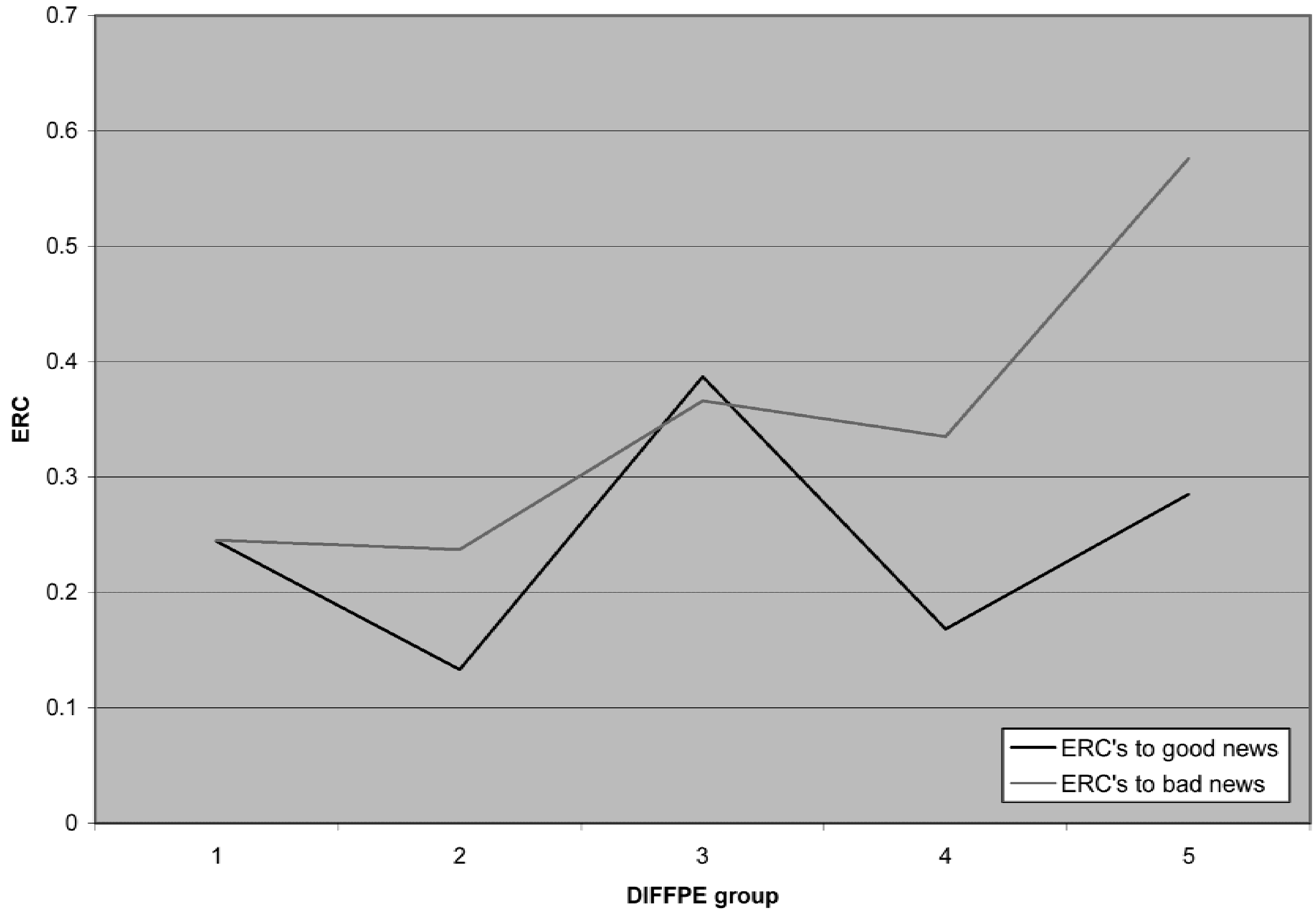

Figure 2. Earnings response coefficients by market level.

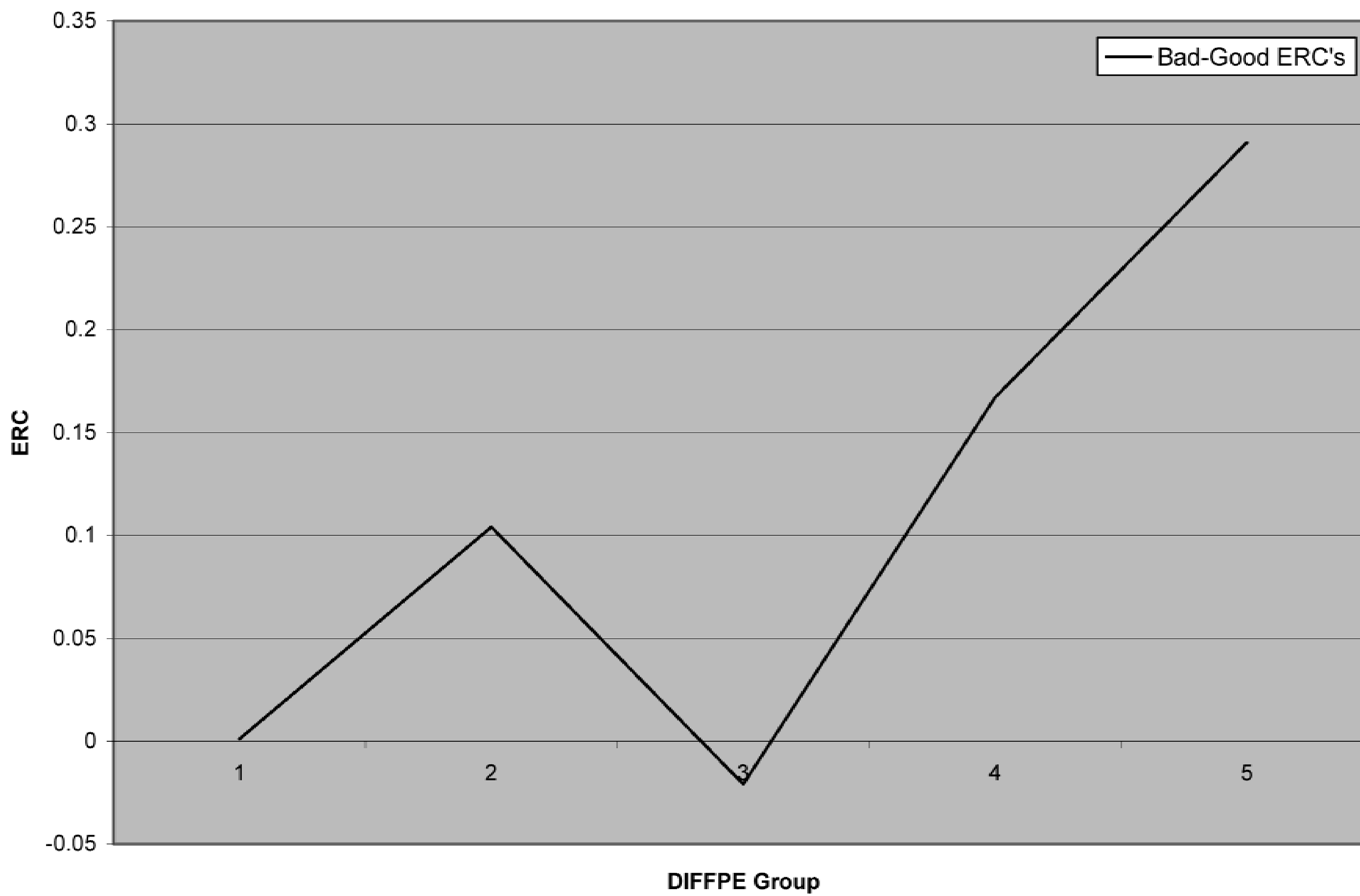

Figure 3. Difference in bad and good news ERCs by market level.

aggregate market predictions, it is worth determining whether analogous results hold for individual firms. That is, does the difference between bad news and good news response coefficients in up markets exceed the difference between the good news and bad news response coefficients in down markets? Alternatively stated, is the difference between bad news and good news response coefficients increasing across the *DIFFPE* portfolios?

The plot of the difference between the bad news and good news response coefficients is depicted in Figure 3. The figure indicates that findings generally are consistent with the suggestions of the regime shifting applied to the individual firm data. The difference in *UEDOWN* and *UEUP* coefficients rises from 0.001 in the *DIFFPE* = 1 portfolio to 0.291 in the *DIFFPE* = 5 portfolio. A regression of the coefficient differences on *DIFFPE* indicates the slope is significantly positive (p-value = 0.053 under a one-sided alternative). With the exception of the drop from *DIFFPE* = 2 to *DIFFPE* = 3, the difference increases across all portfolios.

It should be noted that the results are also consistent with a straightforward extrapolation of the behavioral models to take account of "market-wide investor confidence." That is, if investors' extrapolation of good news is more pronounced as the aggregate market increases in value, then we would expect higher market prices to be associated with a larger response to bad news. Unlike the regime-shifting models, however, the behavioral models seem to predict a symmetric impact when stock prices fall. This prediction is not consistent with our findings for individual firms.

C. Value–Glamour

One strand of the value/glamour literature reviewed in the first section suggests that greater average returns on value as compared to glamour stocks returns is attributable to the tendency of analysts and/or investors to extrapolate irrationally historical corporate performance. As a result, glamour stocks with historically high growth rates tend to become overpriced. When bad news finally reveals to investors the error of their ways, the prices of glamour stocks fall sharply.

If the irrational extrapolation theory operates because it captures the impact of swing in investor confidence generally, then it should be the case that as the market rises, the asymmetry in the response to good and bad news becomes relatively more pronounced for glamour stocks. To assess whether this is true, we examine how the asymmetrical response to good and bad news varies across stocks as well as over time. Specifically, we sort individual companies into quintiles based on their P/E ratios in the month preceding each firm announcement, where earnings are again based on next year's forecasted earnings. The companies in the lowest quintile are the "value" stocks and the companies in the highest quintile are the "glamour" stocks.[14]

[14] As La Porta et al. (1997) report, the impact of the glamour/value distinction is relatively invariant to variables used to define it, be they E/P, Book/Market, or some similar measure.

We then estimate equations (1) and (2) for each of the five *DIFFPE* groupings, permitting the coefficients of *UEUP* and *UEDOWN* to vary for value stocks (quintile 1), glamour stocks (quintile 5), and average stocks (quintiles 2 through 4).

The results, which are not reported, fail to reveal any significant distinction between the behavior of value and glamour stocks. That is, the findings reported in Table IV, Panel B, and depicted in Figures 2 and 3 hold approximately equally for value and glamour stocks.

In light of the findings of Bernard et al. (1997), La Porta et al. (1997), and Skinner and Sloan (1999), our results may appear to be somewhat surprising. However, the studies are not directly comparable. Not only are the sample periods different, but we examine the response coefficients to unexpected good and bad news in the event period, whereas the previous studies examined mean returns for (longer) periods following earnings announcements.

D. Nasdaq versus NYSE

Although size is included as a control for risk in equations (1) and (2), it is possible that response effects differ for Nasdaq and NYSE firms because of risk differences or other characteristics for which we have not controlled. To explore this possibility, we reestimate equations (1) and (2) separately for Nasdaq and NYSE firms.

Table V, Panels A and B, presents regression summary statistics for equations (1) and (2) for Nasdaq firms; Table VI, Panels A and B, presents regression summary statistics for equations (1) and (2) for NYSE firms. Inspection of findings in both tables indicates that the findings for the overall sample broadly generalize to both sets of firms—stock prices respond most strongly to bad news in good times. Nonetheless, there are some differences between the two sets of companies. Turning first to the results for the Nasdaq firms presented in Table V, Panel B, both the *UEUP* and *UEDOWN* coefficients show evidence of a W shape. This pattern is particularly evident for the *UEUP* coefficients that show no trend whatsoever. This is also confirmed by the regression of the *UEUP* coefficient on *DIFFPE*, which produces a slope that is insignificantly different from zero. In the case of the *UEDOWN* coefficients, the W is much less pronounced and the basic hypothesis holds. The largest coefficient, 0.84, is for portfolio 5, and with the exception of portfolio 3, the coefficients increase monotonically. However, because of the spike up in the portfolio 3 *UEDOWN* coefficient, the regression of the *UEDOWN* coefficient on *DIFFPE* indicates the slope is less significantly positive (p-value = 0.077 under a one-sided alternative) than it is for the full sample in Table IV, Panel B. Finally, the regression of the difference between the *UEDOWN* and *UEUP* coefficients on *DIFFPE* indicates the difference is increasing in relative market levels (p-value = 0.055 under a one-sided alternative), which is consistent with predictions of the regime-shifting models.

In contrast, the results for the NYSE firms reported in Panel B of Table V appear much more similar to those reported for the overall sample in Table IV,

Table V

Findings for Nasdaq Firms: Average Coefficients, Associated *t*-statistics, and Adjusted R^2s for Return Regressions for Annual Earnings Announcements, with a Sample Period Covering Annual Earnings for 1988 to 1998.

The variable *UE* is calculated as (actual earnings − consensus forecast earnings)/price(−6), where price(−6) is the share price six days before the earnings announcement (day 0); *RET* is the six day excess return, from day −5 to day 0, calculated as the sum of a firm's raw returns, less the sum of the CRSP value-weighted market return over the same two-day period. The variable *SIZE* represents the natural log of equity market value in the period prior to the firm's earnings announcement, generally day −6.

Panel A: $RET_{it} = a_0 + a_1 UE_{it} + a_2 SIZE_{it} + \epsilon_{it}$

DIFFPE	*UE*	*SIZE*	Adj. R^2	NOBS
1	0.323	−0.002	0.80	2,378
	(4.36)	(−1.67)		
2	0.191	0.003	0.42	2,436
	(2.87)	(1.84)		
3	0.610	−0.003	1.96	2,361
	(6.75)	(−2.22)		
4	0.225	0.000	0.22	2,327
	(2.67)	(0.20)		
5	0.488	−0.001	0.90	2,392
	(4.75)	(−1.08)		

Panel B: $RET_{it} = b_0 + b_1 UEUP_{it} + b_2 UEDOWN_{it} + b_3 SIZE_{it} + \epsilon_{it}$

DIFFPE	*UEUP*	*UEDOWN*	*SIZE*	Adj. R^2	NOBS
1	0.439	0.266	−0.002	0.80	2,378
	(3.20)	(2.84)	(−1.48)		
2	0.100	0.287	0.002	0.45	2,436
	(1.04)	(2.87)	(1.58)		
3	0.637	0.585	−0.003	1.92	2,361
	(4.57)	(4.48)	(−2.11)		
4	0.183	0.308	0.000	0.20	2,327
	(1.71)	(1.98)	(0.06)		
5	0.325	0.844	−0.002	1.04	2,392
	(2.53)	(4.26)	(−1.56)		

Panel B. The *UEUP* coefficients show no distinct variation across the *DIFFPE* portfolios; a regression of the *UEUP* coefficient on *DIFFPE* produces a slope that is insignificantly different from zero. In comparison, there is a notable uptrend in the *UEDOWN* coefficients and no evidence of a W pattern. The regression of the *UEDOWN* coefficient on *DIFFPE* indicates the slope is significantly positive. However, in contrast to Nasdaq firms, the regression of the difference between the *UEDOWN* and *UEUP* coefficients on *DIFFPE* indicates the difference is not significantly increasing in relative market levels.

Table VI

Findings for NYSE Firms: Average Coefficients, Associated *t*-statistics, and Adjusted R^2s for Return Regressions for Annual Earnings Announcements, with a Sample Period Covering Annual Earnings for 1988 to 1998.

The variable *UE* is calculated as (actual earnings − consensus forecast earnings)/price(−6), where price(−6) is the share price six days before the earnings announcement (day 0); *RET* is the six day excess return, from day −5 to day 0, calculated as the sum of a firm's raw returns, less the sum of the CRSP value-weighted market return over the same two-day period. The variable *SIZE* represents the natural log of equity market value in the period prior to the firm's earnings announcement, generally day −6.

Panel A: $RET_{it} = a_0 + a_1 UE_{it} + a_2 SIZE_{it} + \epsilon_{it}$

DIFFPE	*UE*	*SIZE*	Adj. R^2	NOBS
1	0.175	−0.002	0.54	2,427
	(3.16)	(−2.42)		
2	0.167	−0.000	0.50	2,446
	(3.75)	(−0.61)		
3	0.204	−0.002	0.75	2,406
	(3.66)	(−2.75)		
4	0.254	−0.001	0.50	2,477
	(3.37)	(−1.91)		
5	0.315	−0.003	1.76	2,424
	(5.71)	(−3.49)		

Panel B: $RET_{it} = b_0 + b_1 UEUP_{it} + b_2 UEDOWN_{it} + b_3 SIZE_{it} + \epsilon_{it}$

DIFFPE	*UEUP*	*UEDOWN*	*SIZE*	Adj. R^2	NOBS
1	0.141	0.221	−0.002	0.52	2,427
	(1.84)	(2.46)	(−2.51)		
2	0.177	0.157	−0.000	0.46	2,446
	(2.68)	(2.36)	(−0.55)		
3	0.229	0.175	−0.002	0.71	2,406
	(2.89)	(2.04)	(−2.58)		
4	0.155	0.337	−0.001	0.51	2,477
	(1.30)	(3.12)	(−2.12)		
5	0.262	0.413	−0.003	1.77	2,424
	(3.70)	(4.21)	(−3.68)		

It is also worth noting that the Nasdaq coefficients are almost twice the magnitude of the NYSE coefficients. This added sensitivity of Nasdaq firms is difficult to understand. While it is true that both earnings surprises and stock price movements are larger on average for Nasdaq firms, this does not explain why the response coefficient is larger for a given percentage earnings shock. Furthermore, during our sample, including 1997 and 1998, the aggregate stock returns of Nasdaq and NYSE stocks are not fundamentally different. For example, in the period 1997 through 1998, the difference in aggregate market returns was only approximately 4 percent (69 percent vs. 65 percent). Thus, the pricing difference for the two samples appears not to be a result of differences in stock price runups. Developing a more complete understanding for the pricing differences in the two markets is a challenge for future research.

E. Other Specification Checks

Although we control for firm size in our models by including size as a separate regressor, it is possible that firm size plays an interactive role in the relation between returns and earnings surprises.[15] Consequently, we considered other specifications of the relations between returns, earnings shocks, and size. First, we estimated models in which the size variable was interacted with unexpected earning by adding $UE * SIZE$ to equation (1) and $UEUP * SIZE$ and $UEDOWN * SIZE$ to equation (2). However, the coefficient on the interaction term was rarely significant and had a confusing impact on the *UE* variables because of the high collinearity between *UE* and $UE * SIZE$. For these reasons, the size interactions terms are not included in the reported results. Second, we separated our sample into three size-based portfolios in which size is measured relative to other sample firms at the beginning of each year. We reestimated equations (1) and (2) separately on these size-ranked portfolios. Untabulated findings indicate that inferences for the full sample are reasonably robust for the medium and large market capitalization portfolios.[16] However, for the small market capitalization portfolio, earnings response coefficients are relatively flat across *DIFFPE* portfolios, and the relation between *UEUP* and *UEDOWN* coefficients has no discernible pattern. One possible explanation for the disparity in findings for the smallest firms is that the substantially larger volatility of earnings surprises for small firms may indicate lower information content for these firms, making it difficult to detect a relation between these surprises and excess returns. In addition, the excess returns for the small firms also exhibit greater volatility. Relatedly, the untabulated findings indicate that the earnings response coefficients for small market capitalization firms are estimated with less precision than for medium and large firms.

Skinner and Sloan (1999) find evidence of significant pre-announcement effects in their sample of earnings announcements. To test whether such effects are present in our sample, we examine the period extending from day -20 to day -6. Consistent with Skinner and Sloan, untabulated findings provide evidence of pre-announcement market responses, particularly for

[15] We thank the referee for making this observation.

[16] A regression of the differences in *UEDOWN* and *UEUP* coefficient on *DIFFPE* indicates the slope is positive for medium and large portfolios (p-values = 0.075 and 0.11) under a one-sided alternative.

negative earnings shocks. However, adding the pre-announcement earnings response coefficients to those from the announcement period does not alter the tenor of our reported findings.

Finally, as an additional specification check, we reestimated equations (1) and (2) for each of the *DIFFPE* portfolios compounding returns rather than cumulating them (Bernard and Thomas (1989), Conrad and Kaul (1993)). Not surprisingly, given the relatively short window we examine (6 days in the announcement period and 15 days in the pre-announcement period), the results are virtually identical to those reported in Table IV.

IV. Summary and Concluding Remarks

A practitioner hypothesis, quoted in the *Wall Street Journal*, is that the stock prices of individual firms become relatively more sensitive to bad news than good news as the market rises. This hypothesis is related to two strands of literature. The first strand, based on research in behavioral psychology, suggests that investors inappropriately extrapolate past performance. As a result, a bad news "shock" has a particularly telling impact after a period of good news, because it has the effect of correcting overoptimistic projections. However, that literature focuses on individual firms and does not examine the impact of market-wide shifts in investor confidence. The second strand, based on extended regime shifting models, also predicts that the market will respond more strongly to bad news than good news when stock prices are high. However, those models focus exclusively on market-wide effects. Consequently, if the practitioner hypothesis is correct, then both the behavioral and regime-shifting models need to be extended.

To test the hypothesis that stock prices for individual companies respond more strongly to bad news than good news when stock prices are high, we examine the reaction of stock prices to 24,097 announcements of firms' annual earnings during the period 1988–1998. Firm-announcements are placed into one of five portfolios based on a relative market level of the market measure, the difference between the current market P/E ratio and the average P/E during the preceding 12 months. We then examine the reaction of stock prices to both positive and negative earnings surprises for each portfolio separately.

The findings generally support the hypothesis that stock prices respond most strongly to bad news in good times. In particular, the stock price response to negative earnings surprises is increasing as the market level rises. The results for good news, while less strong, are consistent with the notion that the stock price response to positive earnings surprises is decreasing as the market level rises. However, consistent with the implications of regime-shifting models, the difference between bad news and good news response coefficients is increasing as the market level rises. Findings based on subsamples of Nasdaq and NYSE are broadly consistent. For both subsamples, stock prices respond most strongly to bad news in good times. However, the Nasdaq firms respond both more strongly and more erratically. Instead of rising monotonically across *DIFFPE* portfolios, the response coefficients for both good and bad news earnings surprises exhibit a W-shaped pattern. Findings based on subsamples of small, medium, and large firms indicate the greater stock price response to bad news in good times exists solely for large and medium firms, and there is an overall weakening of the results when firms are partitioned into these size-based portfolios.

In conclusion, the findings clearly suggest several areas for future research. Most importantly, both the behavioral models and the regime-shifting model could be extended to explain more fully how the response of individual firms to earnings announcement may depend on the level of the market. In particular, such extensions could provide more precise guidance for determining how best to measure relative market valuation levels. Another challenge is to understand why Nasdaq firms respond more strongly than NYSE companies, but in a fashion less consistent with the underlying hypothesis. The fact that there is no apparent distinction between value and glamour stocks also invites further research.

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