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Earnings and price momentum ☆

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Abstract

This paper examines whether earnings momentum and price momentum are related. Both in time-series as well as in cross-sectional asset pricing tests, we find that price momentum is captured by the systematic component of earnings momentum. The predictive power of past returns is subsumed by a zero-investment portfolio that is long on stocks with high earnings surprises and short on stocks with low earnings surprises. Further, returns to the earnings-based zero-investment portfolio are significantly related to future macroeconomic activities, including growth in GDP, industrial production, consumption, labor income, inflation, and T-bill returns.

JEL classification: G11; G12; G14; M40

Keywords: Post-earnings announcement drift; Price momentum; Earnings momentum; Market efficiency; Macroeconomy

1. Introduction

In a seminal paper, Fama (1998) once again makes the case for the efficient markets hypothesis. Notwithstanding the recent interest in behavioral finance—an interest that is

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driven by data that are inconsistent with the standard frictionless asset pricing models—Fama argues that the null should still be one of market efficiency. However, Fama concedes that two robust and persistent anomalies pose challenges to the efficient markets paradigm. These two anomalies are (i) the post-earnings announcement drift, or earnings momentum, first documented by Ball and Brown (1968) and (ii) the short-run return continuation, or price momentum, documented by Jegadeesh and Titman (1993). Earnings momentum refers to the fact that firms reporting unexpectedly high earnings subsequently outperform firms reporting unexpectedly low earnings. The superior performance lasts for about nine months after the earnings announcement. Price momentum refers to the strategy that buys past winners and sells past losers, which earns abnormal returns for a period of up to one year after the execution of the strategy.

In this paper, we study whether earnings momentum and price momentum are related. Our analysis extends Chan et al. (1996), who also investigate whether the predictability of future returns based on past returns is subsumed by individual stock earnings surprises in cross-sectional tests. If price momentum is related to macroeconomic variables, as shown by Chordia and Shivakumar (2002), Ahn et al. (2003) and Avramov and Chordia (2005), then firm-specific characteristics, such as earnings surprises, will be insufficient to capture price momentum. We seek a relation between price momentum and the systematic component of earnings momentum.

Based on the most recent earnings surprise, measured as standardized unexpected earnings (*SUE*) following standard practice in post-earnings announcement drift literature, we sort firms into decile portfolios and then examine whether a zero-investment portfolio that is long the highest earnings surprise portfolio and short the lowest earnings surprise portfolio captures the price momentum phenomenon. Both in time-series and cross-sectional asset pricing tests, we find that the earnings-based zero-investment portfolio (denoted *PMN* for positive minus negative) captures the payoffs to price momentum strategies. For instance, the price momentum effect (as measured by the portfolio *WML*, which is long past winners and short past losers), at about 76 basis points per month, is reduced to essentially zero in time-series tests after controlling for the exposure of firms to *PMN*. Since *PMN* is a diversified portfolio, it is unlikely to reflect any firm-specific information. Thus, the above results are consistent with price momentum being primarily related to the systematic component of earnings momentum.

To better understand the ability of *PMN* to explain price momentum, we analyze the properties of *PMN*. During our sample period from January 1972 through December 1999, the payoffs to the *PMN* portfolio average a significant 90 basis points per month. These payoffs are not subsumed by the Fama and French (1993) factors or the momentum factor of Carhart (1997). Thus, while the earnings momentum anomaly subsumes the price momentum anomaly, it is not itself subsumed by the price momentum anomaly. The correlation between *PMN* and the price momentum based portfolio, *WML*, is 0.66. Also, *WML* is more volatile than *PMN*. These results suggest that price momentum is a noisy proxy for earnings momentum. This is consistent with the results in Hong et al. (2003) who examine earnings and price momentum in 11 international equity markets and find that price momentum exists only in those countries in which earnings momentum is profitable.

Using a variety of measures to capture future macroeconomic conditions, we show that the return on *PMN* forecasts future business conditions. In particular, we find that the return on *PMN* is correlated with future GDP growth, industrial production growth, consumption growth, labor income growth, inflation, and T-bill returns. These correlations persist even after controlling for the Fama–French factors. These results suggest that the *PMN* portfolio may be viewed as a risk factor that earns a risk premium.[1] However, *PMN* is negatively related to the business cycle as measured by GDP growth. Portfolios that vary countercyclically with the business cycle should not earn a positive risk premium. Thus, while *PMN* is related to the business cycle, it is unlikely to proxy for a risk factor. Overall, these results suggest that earnings momentum (or post-earnings-announcement-drift) contains a systematic component related to the macroeconomy, but that this component is unlikely to represent a (macroeconomic) risk factor.[2]

Chordia and Shivakumar (2005) offer a potential explanation for the systematic component in earnings momentum by showing that earnings momentum, or the post-earnings announcement drift, results from an inflation illusion. The inflation illusion hypothesis, first proposed by Modigliani and Cohn (1979), suggests that while bond market investors correctly anticipate the impact of inflation on discount rates, stock market investors fail to incorporate inflation when forecasting the rate of future earnings growth.[3] Thus, when inflation rises, investors do not adjust the future earnings growth rate, even though they fully adjust the discount rates. A direct implication of this hypothesis is that if earnings growth varies across stocks in response to inflation, then an inflation illusion would induce misvaluation in the cross-section. Chordia and Shivakumar (2005) show that the effect of inflation on earnings growth increases monotonically across *SUE*-sorted portfolios. Due to inflation illusion, stocks with earnings growth positively related to inflation are undervalued, whereas those with earnings growth negatively related to inflation are overvalued. The subsequent correction of this under- and overvaluation drives the post-earnings-announcement-drift.

This paper contributes to the ongoing debate on the sources of profits to price momentum. Several studies suggest that the momentum profits are driven by cognitive biases on the part of investors (e.g., Daniel et al., 1998; Barberis et al., 1998). In contrast, Chordia and Shivakumar (2002), Ahn et al. (2003), and Avramov and Chordia (2005) argue that the price momentum payoffs are related to the business cycle. Korajcyzk and Sadka (2004) argue that the momentum phenomenon persists due to frictions in the price adjustment process that are caused by transactions costs. The finding that price momentum is subsumed by a common factor related to the macroeconomy is significant since it does not rely on capital market frictions to explain the price momentum effect.

This paper narrows the search for an explanation of the price momentum and earnings momentum anomalies by documenting that the price momentum anomaly is a manifestation of the earnings momentum anomaly. That is, the two anomalies that Fama (1998) cites as being above suspicion may, in fact, correspond to the same anomaly, namely, the earnings momentum, or the post-earnings announcement drift, anomaly. Moreover, our results indicate that the price momentum-based factor, *WML*, in the Carhart (1997) four-factor model is merely a noisy proxy for the earnings momentum-based factor, *PMN*. This implies that *PMN*, rather than *WML*, is the more appropriate

[1]Cochrane (2000) suggests that "the central and unfinished task of absolute asset pricing is to understand and measure the sources of aggregate or macroeconomic risk that drives asset prices." Based on a survey of 392 CFOs, Graham and Harvey (2001) find that next to market risk, macroeconomic risks such as business cycle risk and inflation risks are the most important risk factors that firms consider in computing their costs of capital.

[2]Daniel et al. (2001) derive an asset pricing model that includes nonrisk factors.

[3]Basu et al. (2005) show that financial analysts also suffer from inflation illusion.

factor to use in asset pricing tests. Of course, both *PMN* and *WML* are empirically motivated and neither may represent a state variable in the Merton (1973) sense.

The rest of this paper is organized as follows. Section 2 discusses the two momentum strategies, and Section 3 discusses the formation of portfolios that are based on these momentum strategies. Section 4 presents time-series and cross-sectional asset pricing tests. Section 5 presents the properties of *PMN* and Section 6 analyses the link between *PMN* and the macroeconomy. Section 7 concludes.

2. Momentum strategies

As we mention above, the two anomalies that we focus on in this paper are the price momentum and earnings momentum anomalies. The profitability of price momentum strategies, first documented by Jegadeesh and Titman (1993), has been particularly intriguing, as, among all the anomalies examined by Fama and French (1996), it is the only anomaly that is unexplained by the Fama and French (1993) three-factor model. Jegadeesh and Titman (2001) argue that their initial results are not due to data mining as profits to momentum strategies of about 1% per month obtained in the 1990s. Furthermore, Rouwenhorst (1998) confirms the robustness of this strategy using data from stock markets in addition to that of the U.S., for which the profitability of this strategy was initially identified. Specifically, Rouwenhorst (1998) finds significantly positive momentum payoffs in 12 non-U.S. countries he examines in his study.

Earnings momentum, or the post-earnings announcement drift, is first documented by Ball and Brown (1968). Foster et al. (1984), Bernard and Thomas (1989), and others confirm the robustness of Ball and Brown's (1968) findings using more recent data, and Foster et al. (1984) document an annualized payoff of 25% from earnings momentum strategies. Hew et al. (1996) and Booth et al. (1996) extend the post-earnings announcement drift evidence to non-U.S. data. Note that the post-earnings announcement drift is also a robust anomaly that has defied rational explanations. The phenomenon has been attributed to a delayed price response to information. Since stock prices are likely to be driven by earnings, we test whether the price momentum and the post-earnings announcement drift phenomena are related.

3. Zero-investment portfolios

To study the impact of earnings momentum on price momentum, we first create earnings portfolios that capture the post-earnings announcement drift phenomenon. For each month, we sort all NYSE-AMEX firms on the monthly Center for Research in Security Prices (CRSP) files with data on COMPUSTAT into deciles based on their *SUE* from the most recent earnings announcement.[4] We sort firms into deciles based on current-quarter earnings less earnings four quarters ago. For cross-sectional comparison, we standardize this change in earnings by the standard deviation of the earnings changes in the prior eight quarters. We standardize earnings changes by the standard deviation rather than by the stock price, market capitalization, total assets, or sales as these variables may themselves

[4]We exclude ADRs, REITs, Americus Trust Components, units, and closed-end funds from all our analyses. Data on earnings announcements are available for most Nasdaq stocks as of 1984. Including Nasdaq stocks in our analyses has no qualitative impact on our results.

proxy for size or expected returns, which could bias our results towards capturing cross-sectional differences in the expected returns associated with these variables. Moreover, our methodology is consistent with prior studies in accounting that investigate the post-earnings announcement drift phenomenon (see, e.g., Bernard and Thomas, 1989).[5] We implement this sort for each month using the same methodology as Chan et al. (1996). Thus, in each portfolio formation month, we sort firms using only the most recent earnings announced by the firms. To avoid using stale earnings, we require the most recent earnings announcements to be released no earlier than four months before the end of the formation month.

We form decile portfolios, which we also refer to as *SUE* portfolios, by equally weighting all firms in the decile rankings. The positions are held for the six-month holding period, t through $t+5$. Following Jegadeesh and Titman (1993), we form decile portfolios that avoid test statistics based on overlapping returns. Note that with a six-month holding period, each month's return is a combination of the past six ranking strategies, with only the weights of 1/6 of the securities changing each month, and the rest being carried over from the previous month.

Panel A, Table 1 presents the returns on the *SUE* portfolios. Over the entire sample period from January 1972 to December 1999, the monthly holding-period returns increase monotonically from 0.79% for the lowest *SUE* portfolio, P_1, to 1.68% for the highest *SUE* portfolio, P_{10}. The difference in returns between the highest and the lowest *SUE* portfolios, $P_{10}-P_1$, is a statistically and economically significant 0.9% per month with over 75% of the months having $P_{10}-P_1>0$. These results are consistent with both Foster et al. (1984) and Bernard and Thomas (1989). For instance, based on an event study for the 1974 to 1986 period, Bernard and Thomas (1989) report a significant payoff of 4.2% on a portfolio that is long P_{10} and short P_1 in the 50 event days subsequent to an earnings announcement.

We also conduct subperiod analysis for the periods January 1972–December 1979, January 1980–December 1989, and January 1990–December 1999. In each of the subperiods the difference in the monthly holding-period returns between the highest and the lowest *SUE* portfolio is economically and statistically significant, and we are unable to reject the null that the $P_{10}-P_1$ returns are the same across subperiods. In other words, the results are robust over the entire sample as well as across each of the subperiods. We use the portfolio $P_{10}-P_1$ to study the impact of the post-earnings announcement drift phenomenon on stock returns. We refer to this portfolio as the portfolio *PMN* to signify that the difference between extreme *SUE* portfolios represents positive minus negative earnings changes.[6]

Using the same approach as above, we also form ten price momentum portfolios based on past returns. Thus, for each month t, we rank all NYSE-AMEX stocks with returns for months $t-6$ through $t-1$ into deciles based on their formation period, $t-6$ through $t-1$, returns. The momentum portfolios are formed by equally weighting all firms in the decile rankings. The positions are held for the following six-month period, t through $t+5$.[7] Once

[5]We repeat the analyses allowing for a drift in earnings, following Bernard and Thomas (1989); the results remain qualitatively unchanged with this modification.

[6]We replicate the main results of the paper after defining *PMN* as $P_{10}+P_9+P_8+P_7+P_6-P_5-P_4-P_3-P_2-P_1$.

[7]When we skip a month between the formation and holding periods, the holding period is from $t+1$ through $t+6$.

Table 1

Monthly returns on *SUE* and momentum portfolios For each month, firms are sorted into deciles based on their standardized change in earnings from the most recent earnings announcement (*SUE* portfolios) or on their returns in the past six-months (momentum portfolios). In each month, *SUE* portfolios are computed using all earnings announcements made in the prior four months. Earnings changes are computed using a seasonal random walk model. That is, the standardized unexpected earnings (*SUE*) for month $t = (E_{iq} - E_{iq-4})/\sigma_{iq}$, where E_{iq} is the most recently announced earnings and σ_{iq} is the standard deviation of $(E_{iq} - E_{iq-4})$ over the prior eight quarters. The momentum portfolios are sorted based on the returns in the prior six-month period. The portfolios are held for the following six-month period. The table reports the returns to these portfolios as well as the payoffs from a strategy of being long the highest portfolio (P_{10}) and short the lowest portfolio (P_1). The table also reports the *p*-value from *F*-test for test of equality of payoffs across subperiods. Panel A reports results for *SUE* portfolios, while Panel B reports the results for momentum portfolios.

	Low	2	3	4	5	6	7	8	9	High	*PMN* = high–low
Panel A: *SUE portfolios*											
Jan 1972–Dec 1999											
Mean (%)	0.79	0.97	1.06	1.21	1.34	1.45	1.51	1.60	1.60	1.68	0.90
t-stat	2.39	3.10	3.32	3.75	4.23	4.66	4.90	5.25	5.32	5.69	7.47
% > 0	56.55	58.04	59.52	60.12	61.31	61.31	63.69	63.10	61.90	63.10	75.60
Jan 1972–Dec 1979											
Mean (%)	0.66	0.97	1.11	1.25	1.38	1.39	1.51	1.52	1.63	1.63	0.96
t-stat	0.83	1.26	1.42	1.57	1.79	1.85	2.05	2.14	2.36	2.45	2.95
% > 0	46.88	50.00	52.08	50.00	53.13	53.13	54.17	55.21	56.25	56.25	77.08
Jan 1980–Dec 1989											
Mean (%)	0.90	1.06	1.12	1.37	1.52	1.67	1.69	1.84	1.85	1.95	1.05
t-stat	1.79	2.16	2.25	2.74	3.07	3.36	3.39	3.67	3.69	3.94	7.11
% > 0	59.17	60.00	59.17	63.33	63.33	61.67	65.83	64.17	62.50	65.00	79.17
Jan 1990–Dec 1999											
Mean (%)	0.77	0.89	0.95	1.02	1.14	1.29	1.32	1.43	1.32	1.46	0.69
t-stat	1.75	2.24	2.33	2.51	2.74	3.23	3.44	3.57	3.36	3.62	4.54
% > 0	61.67	62.50	65.83	65.00	65.83	67.50	69.17	68.33	65.83	66.67	70.83
F-test (*p*-value)	0.78	0.79	0.79	0.81	0.76	0.61	0.63	0.52	0.68	0.52	0.66
Panel B: *momentum portfolios*											
Jan 1972–Dec 1999											
Mean (%)	0.84	1.08	1.24	1.25	1.32	1.35	1.33	1.36	1.40	1.60	0.76
t-stat	1.75	3.05	3.87	4.15	4.58	4.72	4.71	4.73	4.72	4.77	2.48
% > 0	50.89	55.36	59.52	61.90	62.50	63.69	63.69	63.39	62.80	63.99	65.48
Jan 1972–Dec 1979											
Mean (%)	1.25	1.24	1.23	1.20	1.13	1.14	1.22	1.24	1.28	1.41	0.16
t-stat	1.17	1.42	1.53	1.61	1.60	1.66	1.82	1.85	1.90	1.95	0.25
% > 0	46.88	46.88	52.08	56.25	52.08	53.13	54.17	54.17	55.21	59.38	65.63
Jan 1980–Dec 1989											
Mean (%)	0.30	1.09	1.42	1.46	1.63	1.71	1.61	1.71	1.74	1.74	1.44
t-stat	0.47	2.14	2.98	3.15	3.50	3.66	3.40	3.51	3.37	2.96	4.28
% > 0	50.83	57.50	60.83	63.33	65.83	65.83	65.00	65.00	65.00	64.17	67.50
Jan 1990–Dec 1999											
Mean (%)	1.05	0.94	1.08	1.07	1.17	1.15	1.14	1.10	1.17	1.62	0.57
t-stat	1.28	1.92	2.59	2.87	3.35	3.34	3.31	3.13	3.15	3.55	0.97
% > 0	54.17	60.00	64.17	65.00	67.50	70.00	70.00	69.17	66.67	67.50	63.33
F-test (*p*-value)	0.70	0.95	0.91	0.86	0.73	0.64	0.76	0.64	0.70	0.93	0.23

again, we follow Jegadeesh and Titman (1993) and compute portfolio returns that avoid overlapping returns. Table 1, Panel B presents the results.

Over the entire sample period from January 1972 through December 1999, the monthly holding-period returns increase from 0.84% for the lowest past-return portfolio, P_1, to 1.60% for the highest past-return portfolio, P_{10}. The difference in returns between the highest and the lowest past-return portfolios, $P_{10}-P_1$, is a statistically and economically significant 0.76% per month with over 65% of the months having $P_{10}-P_1>0$.[8] This result is consistent with Grundy and Martin (2001) who document a payoff of 0.86% per month over the sample period 1962 to 1995 and Chordia and Shivakumar (2002) who report a payoff of 0.73% per month over the sample period 1963–1994. We refer to the zero-investment portfolio, $P_{10}-P_1$, as *WML* to denote winners minus losers.

Subperiod analysis shows that the payoff to buying winners and selling losers is an insignificant 0.16% per month over January 1972 through December 1979,[9] 1.44% per month over January 1980 through December 1989, and 0.57% per month over January 1990 through December 1999. The lack of a significant payoff to *WML* in the 1990s is driven primarily by a negative payoff to the strategy in the early 1990s, as this period covers an economic recession, during which time momentum strategies are known to earn negative payoffs (Chordia and Shivakumar, 2002). When payoffs are measured over 1993 to 1999, the average payoffs increase to a statistically significant 1.2% per month (t-statistic = 2.88). In each subperiod, over 60% of the months realize a positive payoff. While there is wide variation in the average monthly returns to *WML* across the subperiods, we are unable to reject the null that the payoffs are the same across the periods.

4. Asset pricing tests

To study the relation between earnings- and price-based momentum strategies, we examine whether the systematic component of one strategy fully subsumes payoffs to the other. The primary motivation for our focus on the systematic component are the findings of Chordia and Shivakumar (2002), Ahn et al. (2003), and Avramov and Chordia (2005) that price momentum is related to macroeconomic variables and that it is unrelated to firm-specific news. We implement our tests by extending the Fama–French three-factor model to include either earnings- or price momentum-based zero-investment portfolios, i.e., *PMN* or *WML*, and then examining the ability of this model to explain payoffs to the other momentum strategy. Since the zero-investment portfolios are well diversified, their returns reflect only systematic information. Fama and French (1996) show that their three-factor model captures all CAPM-related anomalies except for momentum. The issue then is whether including *PMN* or *WML* along with the Fama–French factors can overcome this limitation.

The following subsection presents asset pricing tests in a time-series context, while Section 4.2 discusses cross-sectional asset pricing tests. In Section 4.3, we reevaluate the relation between the two momentum strategies using proxies for firm-specific news rather than using diversified hedge portfolios and relate our findings to those of Chan et al. (1996).

[8]With a one-month gap between the formation and holding periods the average monthly payoff is 1.11% (t-statistic = 3.87).

[9]Jegadeesh and Titman (1993) also find insignificant momentum payoffs in the 1970s.

4.1. Time-series tests

We initially replicate the result from Fama and French (1996) asset pricing tests for the price momentum anomaly. We regress the returns of each momentum portfolio on the three-factor Fama–French model and, using the Gibbons et al. (1989) (GRS) statistic, test the null hypothesis that estimated intercepts are equal to zero across all portfolios.[10] The results from this replication are presented in Table 2, Panel A. Consistent with Fama and French (1996), we find that the intercepts increase monotonically from –0.86% per month for the loser portfolio to 0.26% per month for the winner portfolio. Thus, even after controlling for Fama–French factors, a strategy of buying winners and selling losers generates a payoff of 1.12% per month. In fact, compared to Panel B of Table 1, the risk-adjusted payoff to buying winners and selling losers is even higher, suggesting that the loser portfolio is riskier but earns a lower return than the winner portfolio. The GRS test statistic is highly significant (p-value<0.001), and therefore rejects the null hypothesis that the Fama–French three-factor model is well specified for momentum portfolios.

We extend the Fama–French model by including the earnings-based zero-investment portfolio, *PMN*, as an additional factor. Under this model, expected excess return on a portfolio is explained by the sensitivity of its return to the three Fama–French factors and the earnings-based factor, *PMN*.

$$E(R_i) - R_F = b_i * [E(R_M) - R_F] + s_i * E(SMB) + h_i * E(HML) + p_i * E(PMN), \tag{1}$$

where $E(R_M) - R_F$, $E(SMB)$, $E(HML)$, and $E(PMN)$ are expected premia and the factor loadings are the slopes in the time-series regression

$$R_i - R_F = \alpha_i + b_i * (R_M - R_F) + s_i * SMB + h_i * HML + p_i * PMN + e_i. \tag{2}$$

From Panel B of Table 2, we observe that the coefficient on *PMN* is highly significant for most of the portfolios and increases monotonically from -1.09 for the loser portfolio to 0.37 for the winner portfolio. This indicates that a firm's exposure to *PMN* systematically varies across the momentum portfolios. More importantly, the estimated intercepts decrease from 0.29 for the loser portfolio to -0.12 for the winner portfolio, suggesting that the price momentum strategy generates a negative payoff, if any, after controlling for the portfolios' exposures to *PMN*. The Fama and French (1993) three-factor model augmented by *PMN* captures the impact of past returns on future returns. Given the robustness of price momentum, this is a significant finding. Indeed, this is strong evidence that the short-term return continuations of Jegadeesh and Titman (1993) are primarily attributable to cross-sectional variation in the exposure of momentum portfolios to the earnings-based factor, *PMN*.

The above findings raise the question of whether a factor based on price momentum subsumes the payoffs to the *SUE* portfolios. To evaluate this question, we first use the Fama–French three-factor model in Panel C and in Panel D we test the ability of the Carhart (1997) four-factor model to explain returns across *SUE* portfolios. The results, reported in Panel C of Table 2, show that the estimated intercepts increase monotonically from a low of -0.70% to a high of 0.35% per month, suggesting that the Fama and

[10]We thank Kenneth French for making the Fama–French factors available on his website http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/.

Table 2

Time-series regressions This table reports the regression estimates from time-series regressions of excess portfolio returns on the Fama–French three-factor model, on a four-factor model that extends the Fama–French model by including *PMN* as a factor, or on the four-factor Carhart (1997) model. In Panels A and B the portfolios are sorted into deciles based on past six-month returns; in Panels C and D the portfolios are sorted based on the most recent standardized unexpected earnings. This table presents the Gibbons et al. (1989) test statistics and the associated *p*-values.

	Low	2	3	4	5	6	7	8	9	High
Panel A: momentum portfolios—three-factor model										
INTERCEPT	−0.86	−0.44	−0.23	−0.19	−0.08	−0.05	−0.05	−0.02	0.04	0.26
MKT	1.24	1.10	1.07	1.05	1.03	1.04	1.03	1.04	1.05	1.08
SMB	1.63	1.10	0.91	0.78	0.72	0.67	0.64	0.64	0.67	0.86
HML	0.77	0.60	0.55	0.52	0.47	0.45	0.42	0.37	0.31	0.16
t (intercept)	−3.45	−3.47	−2.24	−2.35	−1.20	−0.86	−0.89	−0.24	0.58	2.44
t (*MKT*)	20.16	34.91	42.87	52.28	62.68	68.46	71.29	66.21	56.26	40.56
t (*SMB*)	18.62	24.52	25.71	27.46	30.95	30.89	31.10	28.51	25.05	22.60
t (*HML*)	7.92	12.10	13.97	16.49	18.04	18.89	18.19	14.93	10.56	3.91
Adj-R^2 (%)	74.9	87.8	90.8	93.2	95.0	95.7	96.0	95.4	93.9	90.4
GRS test statistic: 17.56										
GRS (*p*-value): 0.000										
Panel B: momentum portfolios—four-factor model with PMN										
INTERCEPT	0.29	0.20	0.24	0.12	0.09	0.01	−0.07	−0.14	−0.20	−0.12
MKT	1.18	1.07	1.05	1.03	1.02	1.04	1.03	1.05	1.07	1.10
SMB	1.32	0.92	0.79	0.70	0.68	0.65	0.64	0.67	0.73	0.96
HML	0.45	0.42	0.42	0.44	0.42	0.44	0.42	0.41	0.38	0.27
PMN	−1.09	−0.61	−0.44	−0.29	−0.16	−0.06	0.01	0.12	0.23	0.37
t (intercept)	1.16	1.67	2.39	1.38	1.25	0.14	−0.99	−1.98	−2.44	−1.09
t (*MKT*)	21.74	39.58	47.45	55.61	64.00	68.14	70.84	67.55	59.87	44.18
t (*SMB*)	16.00	22.59	23.50	24.82	27.92	28.10	29.01	28.42	26.98	25.30
t (*HML*)	4.93	9.33	11.36	14.01	15.72	17.07	17.15	15.58	12.71	6.51
t (*PMN*)	−9.94	−11.30	−9.96	−7.78	−5.01	−1.93	0.43	3.79	6.38	7.31
Adj-R^2 (%)	80.6	91.2	92.9	94.2	95.4	95.7	96.0	95.6	94.6	91.7
GRS test statistic: 2.62										
GRS (*p*-value): 0.04										
Panel C: SUE portfolios—three-factor model										
INTERCEPT	−0.70	−0.49	−0.40	−0.28	−0.12	0.01	0.09	0.21	0.24	0.35
MKT	1.09	1.07	1.07	1.10	1.07	1.07	1.08	1.07	1.05	1.04
SMB	0.94	0.86	0.91	0.92	0.92	0.85	0.80	0.75	0.72	0.66
HML	0.54	0.51	0.52	0.56	0.54	0.48	0.44	0.36	0.31	0.24
t (intercept)	−7.26	−5.82	−4.61	−3.41	−1.55	0.11	1.26	3.17	3.50	4.87
t (*MKT*)	46.04	52.13	50.17	54.04	55.69	59.73	63.45	66.63	62.72	58.19
t (*SMB*)	27.76	29.41	29.98	31.87	33.69	33.14	33.06	32.94	30.16	25.92
t (*HML*)	14.38	15.72	15.31	17.51	17.69	16.72	16.48	14.02	11.53	8.62
Adj-R^2 (%)	92.0	93.4	93.1	93.9	94.4	94.9	95.3	95.7	95.2	94.4
GRS test statistic: 11.23										
GRS (*p*-value): 0.000										
Panel D: SUE portfolios—four-factor Carhart model										
INTERCEPT	−0.44	−0.29	−0.21	−0.11	0.04	0.12	0.16	0.25	0.25	0.36
MKT	1.06	1.05	1.05	1.07	1.05	1.06	1.06	1.06	1.05	1.04

Table 2 (*continued*)

	Low	2	3	4	5	6	7	8	9	High
SMB	0.76	0.73	0.78	0.80	0.81	0.77	0.74	0.72	0.71	0.65
HML	0.40	0.41	0.41	0.47	0.45	0.42	0.40	0.33	0.30	0.24
WML	−0.23	−0.17	−0.17	−0.16	−0.14	−0.10	−0.07	−0.04	−0.01	−0.01
t (intercept)	−5.99	−4.16	−2.82	−1.48	0.57	1.74	2.42	3.88	3.51	4.88
t (*MKT*)	59.12	61.34	58.68	61.88	63.24	63.22	65.31	66.77	61.96	57.44
t (*SMB*)	27.68	27.85	28.41	30.02	31.93	29.96	29.68	29.44	27.34	23.45
t (*HML*)	13.66	14.61	14.11	16.47	16.50	15.12	14.77	12.65	10.81	8.11
t (*WML*)	−16.29	−12.67	−12.45	−11.45	−11.16	−7.57	−5.78	−3.35	−0.59	−0.59
Adj-R^2 (%)	95.5	95.5	95.3	95.6	95.9	95.6	95.8	95.9	95.2	94.3

GRS test statistic: 7.31
GRS (*p*-value): 0.00

French (1993) model does not capture the impact of earnings surprises on returns. In Panel D, the estimated intercepts continue to increase monotonically from –0.44% per month to a high of 0.36% per month. This suggests that, even after controlling for price momentum, a strategy of buying the highest *SUE* portfolio and selling the lowest *SUE* portfolio would earn a significant payoff of 0.80% per month. Further, the GRS test statistic of 7.31 (p-value<0.001) rejects the null hypothesis that the Carhart (1997) model can explain the returns to the *SUE* portfolios. Thus, although earnings momentum captures the price momentum effect, the converse is not true.

The above results are based on analysis of momentum strategies that hold portfolios for six months following the portfolio formation month. However, Jegadeesh and Titman (1993) show that momentum payoffs are significantly positive for as much as one year after the formation month. We test the robustness of the above results to varying the length of the holding period from three months to 12 months. We also verify the sensitivity of the results to using prior three, nine or 12-month returns, instead of prior six-month returns, to sort stocks into momentum portfolios. Further, since Griffin et al. (2003) and Cooper et al. (2004) argue that the profitability of price momentum strategies is sensitive to controls for microstructure-induced biases, we test the robustness of the results to excluding small firms and penny stocks as well as to allowing for a month's gap between the portfolio formation month and the beginning of holding period.

Table 3 presents the robustness tests. This table reports the intercepts from the regression of *WML* (*PMN*) on the Fama–French factors and *PMN* (*WML*), using *WML* and *PMN* values obtained from varying the length of the holding period or formation period and/or including controls for microstructure-induced biases.[11] Of the 64 regressions of *WML* on the Fama–French factors and the *PMN* that we estimate, the intercept is significantly positive at the 5% level in only one of the regressions. But even in this one regression, the magnitude of the intercept is only 0.28. Although not reported, the coefficients on *PMN* are always significant in these regressions, with *t*-statistics exceeding 10.0 in magnitude and the adjusted-R^2 varying between 43% and 64%. In contrast to the

[11]To conserve space, we only report the intercepts in Table 3. The full regression results are available from the authors upon request.

results from the regression of *WML*, intercepts from the regression of *PMN* on the Fama–French factors and *WML* are always in excess of 0.3 and are significantly positive in all 64 regressions. These results confirm the robustness of our earlier findings that while price momentum is almost entirely explained by earnings momentum, the converse is not true.

There is one more concern about the manner in which *PMN* and *WML* are calculated. Recall that *PMN* is the difference in returns between the extreme *SUE* portfolios, and that *SUE* is standardized by past volatility. On the other hand, while *WML* represents the difference in returns between the past returns of the extreme winners and losers, *WML* is not standardized by past volatility. It may be the case that *PMN* is able to explain *WML* due to the standardization by past volatility. Thus, we determine whether our results are robust when *WML* is formed on the basis of past standardized returns. In unreported results, when we regress the standardized *WML* on the Fama–French factors as well as *PMN*, we find that the intercepts are insignificant while the coefficient on *PMN* is highly significant and the adjusted-R^2 varies from 35% to 39%. Moreover, when we regress *PMN* on the standardized *WML* as well as the Fama–French factors, we find that the intercepts remain highly significant. This confirms that the results in Table 3 are robust to whether or not *WML* is formed by standardizing past returns by volatility.

4.2. Cross-sectional tests

We use the Brennan et al. (1998) methodology in our cross-sectional asset pricing tests. The Brennan et al. (1998) methodology examines individual security returns adjusted for their exposure to known factors. This approach not only avoids the data-snooping biases that are inherent in the portfolio-based approaches (see Lo and MacKinlay, 1990) but also avoids the error-in-variables bias created by errors in estimating factor loadings.

Assume that returns are generated by an L-factor approximate factor model:

$$\tilde{R}_{jt} = \mathrm{E}(\tilde{R}_{jt}) + \sum_{k=1}^{L} \beta_{jk}\tilde{f}_{kt} + \tilde{e}_{jt}, \tag{3}$$

where R_{jt} is the return on security j at time t, and f_{kt} is the return on the kth factor at time t. We begin by estimating for each year from 1972 to 1999 the factor loadings β_{jk} for all securities that had at least 24 return observations over the prior 60 months. Since our factor data begins in January 1972, the factor loadings in the first month of the regression period (January 1974) are estimated from 24 observations per factor, the factor loadings in the second month of the regression period are estimated from 25 observations, and so on, until the 60-month level is reached, at which point the observation interval is kept constant at 60 months. In order to allow for thin trading, we use the Dimson (1979) procedure with one lag to adjust the estimated factor loadings.

The equilibrium version of the arbitrage pricing theory (APT) in which the market portfolio is well diversified with respect to the factors can be written as

$$\mathrm{E}(\tilde{R}_{jt}) - R_{Ft} = \sum_{k=1}^{L} \lambda_{kt}\beta_{jk}, \tag{4}$$

where R_{Ft} is the return on the riskless asset and λ_{kt} is the premium for factor k.

Table 3
Robustness tests for time-series regressions This table presents the intercepts from regressions of *WML* (*PMN*) on Fama–French factors and *PMN* (*WML*). *PMN* is a zero-investment portfolio that is long stocks in the highest *SUE* decile and short stocks in the lowest *SUE* decile. *WML* is defined as returns on the winner decile minus the loser decile, where the winner and loser momentum portfolios are obtained by sorting stocks on past stock returns. The momentum portfolios are formed using either the prior three, six, nine, or 12 months, either after allowing for a one-month gap between the formation and holding periods or without allowing for such a gap. The analysis may exclude stocks with prices less than $1.00 at the end of formation month, as well as the smallest decile of stocks in the formation month. Both momentum and *SUE* portfolios are held either for three-months, six-months, nine-months, or 12-months after the portfolio formation. *t*-statistics are presented within parentheses.

Holding period	1-month gap	Exclusion criteria	Dependent variable = *WML*				Dependent variable = *PMN*			
			Formation period				Formation period			
			3 months	6 months	9 months	12 months	3 months	6 months	9 months	12 months
3 mon	NO	NONE	−1.29	−1.01	−0.81	−0.55	1.21	1.08	1.00	0.95
			(−5.18)	(−3.56)	(−2.66)	(−1.83)	(12.13)	(10.89)	(10.12)	(9.48)
3 mon	YES	NONE	−0.49	−0.53	−0.12	−0.37	1.03	0.97	0.89	0.92
			(−2.13)	(−1.92)	(−0.42)	(−1.24)	(10.00)	(9.63)	(8.55)	(9.06)
3 mon	YES	Price < $1	−0.16	−0.07	−0.11	0.07	0.92	0.81	0.87	0.75
			(−0.84)	(−0.33)	(−0.41)	(0.29)	(9.02)	(8.21)	(8.55)	(7.60)
3 mon	YES	Smallest decile	−0.47	−0.45	−0.06	−0.29	1.04	0.96	0.87	0.90
			(−2.17)	(−1.76)	(−0.21)	(−1.06)	(10.30)	(9.73)	(8.58)	(9.14)
6 mon	NO	NONE	−0.62	−0.41	−0.19	−0.40	0.89	0.80	0.74	0.77
			(−2.91)	(−1.55)	(−0.70)	(−1.39)	(9.69)	(8.70)	(7.98)	(8.44)
6 mon	YES	NONE	−0.26	−0.04	−0.02	−0.41	0.80	0.73	0.71	0.78
			(−1.28)	(−0.17)	(−0.06)	(−1.41)	(8.57)	(7.76)	(7.55)	(8.49)
6 mon	YES	Price < $1	0.02	0.30	0.02	−0.02	0.64	0.54	0.64	0.57
			(0.10)	(1.55)	(0.07)	(−0.11)	(6.95)	(5.83)	(7.02)	(6.63)

6 mon	YES	Smallest decile	−0.20	0.06	0.10	−0.29	0.77	0.69	0.65	0.72
			(−1.02)	(0.27)	(0.43)	(−1.12)	(8.32)	(7.36)	(7.09)	(8.08)
9 mon	NO	NONE	−0.39	−0.14	−0.29	−0.53	0.67	0.59	0.59	0.63
			(−2.14)	(−0.60)	(−1.15)	(−1.99)	(8.16)	(7.05)	(7.30)	(7.93)
9 mon	YES	NONE	−0.01	0.00	−0.23	−0.58	0.56	0.55	0.59	0.65
			(−0.05)	(0.01)	(−0.97)	(−2.19)	(6.67)	(6.59)	(7.18)	(8.10)
9 mon	YES	Price < $1	0.27	0.25	−0.15	−0.14	0.42	0.40	0.55	0.49
			(1.93)	(1.82)	(−0.70)	(−0.74)	(4.83)	(4.66)	(6.70)	(6.32)
9 mon	YES	Smallest decile	0.08	0.13	−0.09	−0.42	0.56	0.54	0.56	0.63
			(0.48)	(0.62)	(−0.38)	(−1.76)	(6.46)	(6.29)	(6.82)	(7.79)
12 mon	NO	NONE	−0.07	−0.14	−0.32	−0.53	0.45	0.45	0.48	0.52
			(−0.47)	(−0.68)	(−1.37)	(−2.14)	(5.92)	(6.05)	(6.48)	(7.05)
12 mon	YES	NONE	0.01	−0.16	−0.35	−0.60	0.42	0.46	0.49	0.54
			(0.07)	(−0.79)	(−1.55)	(−2.44)	(5.50)	(6.10)	(6.63)	(7.30)
12 mon	YES	Price < $1	0.28	0.19	−0.25	−0.15	0.30	0.33	0.47	0.41
			(2.20)	(1.21)	(−1.23)	(−0.83)	(3.81)	(4.38)	(6.31)	(5.75)
12 mon	YES	Smallest decile	0.09	−0.02	−0.19	−0.42	0.43	0.46	0.49	0.54
			(0.62)	(−0.11)	(−0.90)	(−1.89)	(5.38)	(5.82)	(6.32)	(7.01)

The factor-adjusted return on each of the securities, $\tilde{R}^*_{jt}$, for each month t of the following year is then calculated as

$$\tilde{R}^*_{jt} \equiv \tilde{R}_{jt} - R_{Ft} - \sum_{k=1}^{L} \beta_{jk} \tilde{F}_{kt}, \tag{5}$$

where $\tilde{F}_{kt} \equiv \lambda_{kt} + \tilde{f}_{kt}$ is the sum of the factor realization and its premium. Our adjustment procedure imposes the assumptions that the zero-beta return equals the risk-free rate, and that the APT factor premium is equal to the excess return on the factor. The factor-adjusted returns from the above equation constitute the raw material for the following equation:

$$\tilde{R}^*_{jt} = c_0 + \sum_{m=1}^{M} c_m Z_{mjt} + \tilde{e}'_{jt}, \tag{6}$$

where Z_{mjt} is the value of security characteristic m for security j in month t.

We first calculate an estimate of the vector of characteristics rewards $\hat{c}_t$ each month from a simple ordinary least squares (OLS) regression, that is,

$$\hat{c}_t = (Z'_t Z_t)^{-1} Z'_t R^*_t, \tag{7}$$

where Z_t is the vector of firm characteristics in month t andR^*_t is the vector of factor-adjusted returns. The standard Fama and Macbeth (1973) estimators are the time-series averages of these coefficients. Note that although the factor loadings are estimated with error, this error affects only the dependent variable, R^*_t. While the factor loadings will be correlated with the security characteristics, Z_t, there is no a priori reason to believe that the errors in the estimated loadings will be correlated with the security characteristics. This implies that the estimated coefficient vector $\hat{c}_t$ should be unbiased.

However, if the errors in the estimated factor loadings are correlated with the security characteristics, the monthly estimates of the coefficients will be correlated with the factor realizations and the Fama and Macbeth estimators will be biased by an amount that depends upon the mean factor realizations. Therefore, the purged estimator is obtained for each of the characteristics as the constant term from the regression of the monthly coefficient estimates on the time series of the factor realizations. This estimator, which was first developed by Black et al. (1972), purges the monthly estimates of the factor-dependent component. The standard errors of the estimators are taken from the time series of monthly estimates in the case of the Fama–Macbeth estimator, and from the standard error of the constant from the OLS regression in the case of the purged estimator.

We require a firm to satisfy the following two criteria in order to be included for analysis in a given month: (1) The firm must have return data in CRSP for the current month, t, and for 24 of the previous 60 months, and have sufficient data to calculate the size, price, and dividend yield as of month $t-2$; (2) The firm must have sufficient data available on the COMPUSTAT tapes to calculate the book-to-market ratio as of December of the previous year. We use a number of firm-specific characteristics as controls following Brennan et al. (1998). Specifically, the following variables are calculated for each stock and each month:

- *SIZE*: the natural logarithm of the market value of the equity of the firm as of the end of the second-to-last month.

- *BM*: the natural logarithm of the ratio of the book value of equity plus deferred taxes to the market value of equity, using the end of the previous year's market and book values. As in Fama and French (1992), the value of *BM* for July of year t to June of year $t+1$ is computed using accounting data at the end of year $t-1$, and book-to-market ratio values greater than the 0.995 fractile or less than the 0.005 fractile are set equal to the 0.995 and 0.005 fractile values, respectively.
- *TURN*: the natural logarithm of share turnover, measured by the number of shares traded divided by the number of shares outstanding in the second-to-last month.
- *SUE*: the most recent standardized unexpected earnings.
- *PRICE*: the natural logarithm of the reciprocal of the share price as reported at the end of the second-to-last month.
- *YLD*: the dividend yield, measured as the sum of all dividends paid over the previous 12 months divided by the share price at the end of the second-to-last month.
- *RET2-3*: the cumulative return over the two months ending at the beginning of the previous month.
- *RET4-6*: the cumulative return over the three months ending three months previously.
- *RET7-12*: the cumulative return over the six months ending six months previously.

The lagged return variables proxy for price momentum effects, as documented by Jegadeesh and Titman (1993). These are constructed to exclude the return during the immediate prior month in order to avoid any spurious association between the prior month return and the current month return caused by thin trading or bid-ask spread effects. In addition, all variables involving the price level are also lagged by one month in order to preclude the possibility that a linear combination of the lagged return variables, the book-to-market variable (which is related to the price level in the previous year), and the reciprocal of the price level could provide a noisy estimate of the return in the previous month, and thereby lead to bid-ask bounce effects.[12]

The results are presented in Table 4. Panel A presents results from regressions that exclude *SUE*, while Panel B presents results from regressions that include *SUE* as an additional explanatory variable. Let us first focus on Panel A. The second column presents the Fama and Macbeth (1973) coefficients when the dependent variable is excess returns. Consistent with prior results,[13] the coefficient on book-to-market is significantly positive and turnover is significantly negative. Firms with a high book-to-market ratio have higher expected returns than firms with a low book-to-market ratio. High turnover stocks have lower expected returns than low turnover stocks, suggesting that turnover is a proxy for liquidity.[14] High past returns also suggest high expected returns, consistent with past returns being a proxy for the momentum effect. The third and the fourth columns present results with risk-adjusted returns, with the risk adjustment being done using the Fama–French factors. The impact of liquidity and price momentum survives the use of the Fama–French factors. This is consistent with the evidence in Brennan et al. (1998). The size effect and the book-to-market effect also have an important impact on the cross-section of returns. The fifth and the sixth columns use the Carhart (1997) four-factor

[12]See Jegadeesh (1990). It is easy to show that thin trading will cause returns to exhibit first order negative serial correlation.

[13]See for instance, Brennan et al. (1998).

[14]Using dollar trading volume instead of turnover does not change any of the results. See Chordia et al. (2001) for the impact of using dollar trading volume on the other coefficients.

Table 4

Cross-sectional asset pricing tests This table presents the Fama–Macbeth estimates of monthly cross-sectional regressions. The dependent variable in the second column is simply the excess return, while in the third and fourth columns it is the factor-adjusted return using the Fama–French (FF) factors. In the fifth and the sixth (seventh and eighth) columns the dependent variable is the adjusted return using the FF factors along with the momentum-based factor *WML* (earnings-based factor *PMN*). The independent variables are measured as deviations from the cross-sectional mean in each period. *SIZE* represents the logarithm of market capitalization in billions of dollars. *BM* is the logarithm of the book-to-market ratio; book-to-market values greater than the 0.995 fractile or less than the 0.005 fractile are set equal to the 0.995 and the 0.005 fractile values, respectively. *TURN* is the logarithm of share turnover. *PRICE* is the logarithm of the reciprocal of the share price. *SUE* is the most recent standardized unexpected earning. *YLD* represents the dividend yield. *RET2-3*, *RET4-6*, *RET7-12* are the cumulative returns over the second through third, fourth through sixth, and seventh through twelfth months prior to the current month, respectively. The estimates in the column labeled "Raw" are the standard Fama–MacBeth coefficients, while the coefficients labeled "Purged" are obtained as the intercept term by regressing the time series of coefficients on the factors. The sample contains all NYSE-AMEX firms from January 1974 through December 1999. All coefficients are multiplied by 100. The *t*-statistics are in parentheses.

	Excess returns	Returns adjusted using FF factors		Returns adjusted using FF factors and *WML*		Returns adjusted using FF factors and *PMN*	
		Raw	Purged	Raw	Purged	Raw	Purged
Panel A: without SUE							
Intercept	0.914	−0.070	−0.116	0.029	0.044	0.099	0.117
	(2.97)	(−1.09)	(−1.87)	(0.67)	(0.85)	(1.61)	(1.80)
SIZE	−0.035	−0.048	−0.078	−0.040	−0.021	−0.027	−0.058
	(−1.05)	(−1.60)	(−2.71)	(−1.58)	(−0.74)	(−0.88)	(−1.66)
BM	0.216	0.100	0.117	0.122	0.077	0.110	0.017
	(4.00)	(2.00)	(2.38)	(2.61)	(1.53)	(2.10)	(0.30)
PRICE	0.095	−0.094	−0.197	0.022	0.146	0.223	0.270
	(0.79)	(−0.89)	(−1.87)	(0.26)	(1.61)	(2.22)	(2.40)
TURN	−0.127	−0.149	−0.170	−0.192	−0.198	−0.173	−0.137
	(−2.58)	(−4.30)	(−4.83)	(−5.01)	(−4.53)	(−4.85)	(−3.41)
YLD	−0.925	0.723	0.293	0.868	1.460	0.220	0.072
	(−0.72)	(0.75)	(0.30)	(0.76)	(1.10)	(0.22)	(0.06)
RET2-3	0.386	0.550	0.645	0.805	0.130	0.800	0.221
	(1.32)	(1.78)	(2.06)	(3.00)	(0.44)	(2.43)	(0.59)
RET4-6	0.825	0.948	1.050	0.840	0.408	0.639	−0.197
	(3.35)	(3.23)	(3.58)	(3.89)	(1.69)	(1.51)	(−0.42)
RET7-12	1.155	0.760	0.963	0.847	0.393	0.637	0.237
	(7.68)	(3.46)	(4.66)	(6.15)	(2.94)	(2.53)	(0.88)
Panel B: with SUE							
Intercept	0.959	−0.040	−0.086	0.029	0.044	0.092	0.097
	(3.08)	(−0.62)	(−1.40)	(0.67)	(0.85)	(1.48)	(1.47)
SIZE	−0.032	−0.047	−0.080	−0.039	−0.018	−0.038	−0.058
	(−0.92)	(−1.57)	(−2.77)	(−1.54)	(−0.64)	(−1.28)	(−1.70)
BM	0.215	0.091	0.097	0.134	0.090	0.055	−0.041
	(3.23)	(1.53)	(1.66)	(2.89)	(1.77)	(0.87)	(−0.58)
PRICE	0.175	−0.010	−0.112	0.082	0.196	0.299	0.359
	(1.45)	(−0.09)	(−1.01)	(0.99)	(2.16)	(2.80)	(3.01)

Table 4 (*continued*)

	Excess returns	Returns adjusted using FF factors		Returns adjusted using FF factors and *WML*		Returns adjusted using FF factors and *PMN*	
		Raw	Purged	Raw	Purged	Raw	Purged
TURN	−0.114	−0.130	−0.140	−0.153	−0.165	−0.135	−0.084
	(−2.00)	(−3.16)	(−3.33)	(−4.04)	(−3.81)	(−3.19)	(−1.77)
YLD	2.250	4.460	3.954	2.548	2.877	4.527	2.848
	(1.33)	(3.51)	(3.07)	(2.24)	(2.18)	(3.33)	(1.82)
SUE	0.305	0.302	0.299	0.289	0.248	0.319	0.280
	(20.53)	(19.62)	(18.81)	(19.67)	(15.10)	(19.42)	(15.03)
RET2-3	0.094	0.256	0.305	0.274	0.312	0.452	−0.129
	(0.31)	(0.79)	(0.93)	(1.03)	(1.06)	(1.31)	(−0.33)
RET4-6	0.286	0.373	0.470	0.410	0.044	0.060	−0.720
	(1.16)	(1.29)	(1.63)	(1.90)	(0.18)	(0.14)	(−1.49)
RET7-12	0.825	0.428	0.626	0.642	0.317	0.338	−0.103
	(5.59)	(1.97)	(3.08)	(4.78)	(1.86)	(1.31)	(−0.37)

model, i.e., Fama–French factors along with *WML* for risk adjustment. The purged estimates suggest that the coefficients on past returns and their significance are substantially reduced. This is not surprising since *WML* is designed to capture the impact of past returns. Even so, the estimated coefficient of *RET7-12* is still significant.

Finally, in the last two columns of Panel A of Table 4, we augment the Fama–French factor by *PMN*. Consistent with the time-series results in Table 2, the momentum effect is considerably weakened. The purged estimator shows that the coefficients on past returns are all insignificant, confirming that the price momentum effect is entirely captured in the cross-section by *PMN*.

Panel B of Table 4 presents results after including *SUE* as one of the characteristics in the monthly cross-sectional regressions. The important result is that the coefficient of *SUE* is highly significant with *t*-statistics always above 15, regardless of whether or how the left-hand side returns are risk-adjusted in the cross-sectional regressions. Also, note that upon inclusion of *SUE*, the impact of past returns is considerably attenuated.[15] The coefficient estimates on past returns are far smaller than in Panel A and are often statistically insignificant. With excess returns as the dependent variable, both *RET2-3* and *RET4-6* are statistically insignificant and the coefficient on *RET7-12* is smaller than in Panel A but is still significant. The same pattern is repeated when we use Fama–French risk-adjusted returns. When we use the Fama–French factors along with *WML*, the purged estimate of *RET7-12* is significant at only the 10% level (although the raw estimates are still significant). Only when the risk adjustment is done with the Fama–French factors and *PMN* does the impact of past returns disappear in the cross-section. Thus, while *SUE* does capture some of the impact of past returns (*RET2-3* and *RET4-6*), the entire impact

[15]This finding is consistent with Chan et al. (1996) who also find that upon inclusion of *SUE* the impact of past returns is considerably attenuated.

(*RET2-3*, *RET4-6*, and *RET7-12*) is captured only when *PMN* is used for risk-adjustment. In other words, while the characteristic *SUE* captures part of momentum effect, the entire momentum effect is explained only when the common factor, *PMN*, along with the Fama–French factors, is used to risk-adjust returns.

4.3. Reconciliation with Chan et al. (1996)

Our results are consistent with Chan et al. (1996), in that *SUE* does capture some of the impact of past returns on future returns. However, if price momentum is related to systematic variables as documented by Chordia and Shivakumar (2002), Ahn et al. (2003) and Avramov and Chordia (2005), then proxies for firm-specific news should be insufficient to capture price momentum. Thus, the above cross-sectional tests also consider the exposure of firms to the systematic component in earnings momentum, as measured by *PMN*. In other words, while Chan et al. (1996) study whether the firm-specific component of earnings momentum subsumes price momentum, we focus on the systematic component of earnings momentum.

To explain the differences between the two studies, Table 5 reports results from cross-sectional regressions similar to those reported in Chan et al. (1996).[16] Over our sample period, firm size is negatively related to excess returns, although the relation is insignificant in regressions that include the book-to-market ratios. The book-to-market ratio has a significantly positive coefficient in all regressions. The most significant characteristic in all the regressions is *SUE*, with *t*-statistics in excess of 10.00. The positive coefficient on *SUE* is consistent with the drift in returns that occurs following earnings announcements. Price momentum, as measured by the lagged six-month return, *R6*, is also significant, when either the book-to-market ratio is not included in the regression or a one-month lag is allowed between the independent variable *R6* and the dependent variable, which is the excess stock return.[17] Allowing for the one-month lag increases the magnitude of the coefficient on *R6* and the associated *t*-statistics. Thus, consistent with Chan et al. (1996), we also find both *SUE* and lagged six-month returns are important characteristics in explaining the cross-section of returns.

In Table 5, Panel B, we regress the time series of the coefficients of the lagged six-month returns on *PMN* and find that the intercept is insignificant and often negative, while the coefficient on *PMN* is large and statistically significant. This suggests that the time series of coefficients on past returns essentially captures the same information as *PMN*. This result is robust to controls for small firms and penny stocks in the regressions. The adjusted-R^2 ranges from 19% to 34%. Once again these results confirm our findings that price momentum is primarily attributable to *PMN*. To examine whether the converse is also true, in Panel C of Table 5 we regress the time series of coefficients for *SUE* on the momentum factor, *WML*. The intercept from this regression is highly significant while the coefficient on *WML* is generally indistinguishable from zero. Moreover, the adjusted-R^2 is low and often negative. Thus, *WML* is insufficient to explain the significantly positive coefficient on *SUE*, and we find again that while earnings momentum subsumes price momentum, the converse is not true, a result that is consistent with Tables 2 and 4.

[16]To avoid overlapping returns, instead of using six-month returns for the dependent variable as in Chan et al. (1996), we used returns measured over the following one-month. However, our conclusions remain unaffected when the returns are measured over a six-month period.

[17]Chan et al. (1996) do not include book-to-market ratio in their analyses.

Table 5
Fama–Macbeth cross-sectional regressions with earnings and price momentum characteristics Panel A of this table reports the Fama–Macbeth coefficients from monthly cross-sectional regressions of individual stock returns in excess of the risk-free rate on the various firm characteristics. For regressions in month t, *SIZE* is measured as the logarithm of market capitalization as of the beginning of that month, *BM* is the book-to-market ratio that is computed using the book value of equity from financial statements ending at least three months prior to month t and the market value of equity at the beginning of month t, *SUE* is the standardized unexpected earnings based on ranking firms as of month $t-1$ and *R6* is the compounded return in the six-months prior the regression month. Panel B (C) reports results from regressing the monthly coefficient of past returns (*SUE*) from cross-sectional regressions on *PMN* (*WML*). *PMN* and *WML* are defined in Table 1. *Jandum* is a dummy variable that equals one for January and zero otherwise. The smallest decile of firms are excluded in regressions III and IV while the penny stocks are excluded in regressions V and VI. The regressions also allow for *R6* to be estimated over the six-month period $t-8$ to $t-7$ (i.e., allowing for a one-month gap between the regression month, i.e., month t, and the computation of *R6*). The Fama–Macbeth t-statistics are in parentheses.

Panel A: *Fama– Macbeth coefficients from monthly cross-sectional regressions*

	Regn I	Regn II	Regn III	Regn IV	Regn V	Regn VI
Exclusion criterion:	None	None	Smallest decile	Smallest decile	Price < \$1	Price < \$1
1-month gap:	No	No	No	Yes	No	Yes
Intercept	3.31	1.77	1.38	1.43	1.51	1.50
	(2.83)	(1.72)	(1.35)	(1.41)	(1.51)	(1.50)
$SIZE_{t-1}$	−0.13	−0.07	−0.05	−0.05	−0.06	−0.06
	(−2.47)	(−1.55)	(−1.12)	(−1.20)	(−1.30)	(−1.31)
BM_{t-1}		0.44	0.39	0.42	0.42	0.45
		(5.01)	(3.78)	(3.97)	(4.92)	(5.12)
SUE_{t-1}	0.14	0.14	0.13	0.12	0.14	0.13
	(12.70)	(11.26)	(10.84)	(10.72)	(11.38)	(11.19)
$R6_{t-1}$	0.54	0.41	0.44	0.67	0.42	0.68
	(2.18)	(1.61)	(1.66)	(2.58)	(1.62)	(2.76)
Adjusted-R^2 (%)	2.75	3.56	3.62	3.60	3.50	3.47

Panel B: *second stage time-series regression of the coefficients on R6 obtained from monthly cross-sectional regressions*

DEPENDENT VARIABLE = Coefficient on *R*6 obtained from:	Exclusion criterion	1-month gap	Intercept	*PMN*	Jandum	Adj-R^2 (%)
Regn I	None	No	−0.40	0.95		20.23
			(−1.62)	(8.87)		
Regn I	None	No	−0.22	0.89	−1.61	20.87
			(−0.83)	(8.04)	(−1.85)	
Regn II	Smallest decile	No	−0.32	0.97		18.75
			(−1.23)	(8.46)		
Regn II	Smallest decile	No	−0.24	0.94	−0.70	18.63
			(−0.87)	(7.90)	(−0.76)	
Regn III	Smallest decile	Yes	−0.07	0.94		18.93
			(−0.29)	(8.51)		
Regn III	Smallest decile	Yes	−0.04	0.93	−0.29	18.69
			(−0.15)	(8.06)	(−0.33)	
Regn IV	Price < \$1	No	−0.38	0.96		21.93
			(−1.58)	(9.32)		
Regn IV	Price < \$1	No	−0.21	0.91	−1.53	22.48
			(−0.82)	(8.44)	(−1.78)	

Table 5 (*continued*)

Panel B: second stage time-series regression of the coefficients on R6 obtained from monthly cross-sectional regressions

DEPENDENT VARIABLE = Coefficient on *R*6 obtained from:	Exclusion criterion	1-month gap	Intercept	*PMN*	Jandum	Adj-R^2 (%)
Regn V	Price < $1	Yes	−0.10	0.94		22.44
			(−0.42)	(9.46)		
Regn V	Price < $1	Yes	0.00	0.91	−0.90	22.48
			(0.01)	(8.74)	(−1.09)	

Panel C: second-stage time-series regression of the coefficients on SUE obtained from monthly cross-sectional regressions

DEPENDENT VARIABLE = Coefficient on *SUE* obtained from:	Exclusion criterion	1-month gap	Intercept	*WML*	Jandum	Adj-R^2 (%)
Regn I	None	No	0.14	0.00		−0.31
			(11.16)	(−0.21)		
Regn I	None	No	0.13	0.00	0.08	0.11
			(9.85)	(0.48)	(1.51)	
Regn II	Smallest decile	No	0.12	0.00		−0.28
			(10.65)	(0.36)		
Regn II	Smallest decile	No	0.12	0.00	0.03	−0.44
			(9.61)	(0.64)	(0.72)	
Regn III	Smallest decile	Yes	0.12	0.00		0.09
			(10.19)	(1.13)		
Regn III	Smallest decile	Yes	0.12	0.00	0.03	−0.10
			(9.14)	(1.30)	(0.66)	
Regn IV	Price < $1	No	0.14	0.00		−0.33
			(10.98)	(0.02)		
Regn IV	Price < $1	No	0.13	0.00	0.06	−0.08
			(9.65)	(0.51)	(1.33)	
Regn V	Price < $1	Yes	0.13	0.00		−0.06
			(10.28)	(0.91)		
Regn V	Price < $1	Yes	0.12	0.00	0.05	0.04
			(9.06)	(1.24)	(1.14)	

In summary, both *SUE* and past returns are important firm characteristics that are related to the cross-section of stock returns. However, *PMN* captures the cross-sectional impact of past returns as well as the payoffs to *WML*. This result supports the argument that the systematic component of earnings momentum is important in explaining price momentum.

The results also suggest that price momentum is a manifestation of earnings momentum, or post-earnings-announcement-drift. This conclusion is consistent with evidence in other studies. For instance, Hong et al. (2003) examine earnings and price momentum in 11 international equity markets and find that price momentum exists only in those countries in which earnings momentum is profitable. Moreover, the payoffs for both anomalies are known to follow a very similar pattern in that for both earnings momentum and price momentum, disproportionately large payoffs occur at earnings announcements subsequent to portfolio formation. For instance, Jegadeesh and Titman (1993) report (on page 88) for their sample that, about 25% of the six-month payoffs occur at earnings announcements. Chan et al. (1996) find that 36% of the six-month payoffs to the momentum strategy occur at earnings announcements subsequent to portfolio formation. If the momentum payoffs were evenly distributed across the six-month holding period, we would expect to observe less than 5% of the payoffs occurring around earnings announcement days in the six-month holding period. Bernard and Thomas (1989) report similar findings for earnings momentum. They show that, on average, about 25% (40%) of the post-earnings announcement drifts of large (small) firms occur at the subsequent earnings announcement.

The rest of the paper is devoted to understanding the zero-investment portfolio, *PMN*, and to answering the question of why *PMN* captures price momentum.

5. Properties of *PMN*

Table 6 documents seasonality in payoffs to the *SUE* portfolios P_{10} and P_1 as well as the zero-investment strategy, *PMN*. The returns on the zero-investment portfolio, *PMN*, are highest in April, July, October and December. It is positive in all months except in January when it is −1.42%. In the non-January months the average $P_{10}-P_1$ returns is 1.11%. This seasonality in payoffs to *PMN* is mainly attributable to the low-*SUE* portfolio P_1, for which the average return is 6.17% in January compared to 0.30% in non-January months. Further, for this portfolio as well as for *PMN*, the null of equal holding-period returns across months is rejected. Although not the focus of our paper, we speculate that negative returns in January and large returns in October through December for *PMN* are consistent with the tax loss selling hypothesis, according to which investors sell poor performing stocks in October through December and buy them back in January. The strong negative January return obtains for price momentum strategies as well, as documented by Jegadeesh and Titman (1993) and Chordia and Shivakumar (2002).

Having seen from Table 3 that the mean *PMN* portfolio returns are not explained by the Fama–French and momentum factors, we now investigate the impact of *PMN* on these factors. To set the stage, Panel A of Table 7 initially reports the mean monthly returns of the various factors. The mean monthly return on the market is 0.62%, on *SMB* is 0.07%, on *HML* is 0.34%, on *PMN* is 0.90%, and on the momentum factor, *WML*, is 0.76%. All the average monthly returns are statistically and economically significant except for that of *SMB*. The insignificant return for *SMB* is consistent with an absence of a small firm effect in recent data.

Table 7, Panel B reports the correlations among the various factors. The correlation of *SMB* with *WML* and *PMN* is −0.38 and −0.34, respectively, suggesting that small stocks comprise a larger fraction of the losers and of the portfolio with negative or low earnings surprises. The main result is that the correlation between *WML* and *PMN* (0.66) is the highest in the table. Thus, price momentum and earnings momentum are highly correlated. Also, note from Table 1 that the standard deviations on *WML* are higher than those on *PMN*. For instance, in the overall sample from January 1972 to December 1999, the monthly standard deviation of returns for *PMN* is 2.21% while that for *WML* is 5.62%. These results are consistent with *WML* being a noisy measure of *PMN*.

Table 7, Panel C reports results from the regression of the Fama–French factors and momentum factor on *PMN* and a January dummy. From the market return regression, we

Table 6
Payoffs to *SUE* portfolios in each calendar month. The *SUE* portfolios are formed as detailed in Table 1. This table presents the holding-period returns to the extreme *SUE* portfolios and to the portfolio that is long the high *SUE* portfolio and short the low *SUE* portfolio (*PMN*), classified by calendar months. The table also presents the *p*-value from the *F*-test that the payoffs are equal across the months.

	P_1		P_{10}		$PMN = P_{10} - P_1$	
	Mean (%)	*t*-stat	Mean (%)	*t*-stat	Mean (%)	*t*-stat
January	6.17	5.62	4.75	4.68	−1.42	−3.63
February	1.56	1.42	2.14	2.11	0.58	1.49
March	1.22	1.11	1.82	1.79	0.60	1.54
April	0.55	0.51	1.92	1.89	1.37	3.51
May	1.30	1.19	1.58	1.55	0.28	0.70
June	0.55	0.51	1.61	1.58	1.05	2.70
July	0.02	0.01	1.45	1.43	1.44	3.67
August	−0.23	−0.21	0.50	0.49	0.73	1.87
September	−0.92	−0.84	0.23	0.23	1.15	2.94
October	−2.04	−1.86	−0.30	−0.30	1.74	4.46
November	0.95	0.86	2.22	2.19	1.28	3.27
December	0.30	0.27	2.28	2.24	1.98	5.06
Non-January	0.30	0.90	1.40	4.60	1.11	9.30
F-test (*p*-value)	0.00		0.11		0.00	

observe that while the market return exhibits a January effect, it is essentially uncorrelated with *PMN*. Both *SMB* and *HML* exhibit a significant positive intercept when *PMN* and the January dummy are used as regressors. The intercept on *SMB* is a significant 0.37% per month and on *HML* is a significant 0.51% per month. The most striking result of Table 7 is that the price momentum effect, which is about 76 basis points per month, is statistically indistinguishable from zero once *PMN* is used as a regressor.[18] This is consistent with the earlier results from asset pricing tests and suggests that the price momentum strategy of buying winners and selling losers is a manifestation of the systematic component of the post-earnings announcement drift.

We now study the characteristics of firms in the two extreme *SUE* portfolios, P_{10} and P_1. Table 8 presents the average firm characteristics across the *SUE* portfolios. Stocks in the

[18]Kraft (2001) examines the relations among various market anomalies, including the price momentum and the post-earnings announcement drift, by regressing hedge portfolio returns of different strategies on each other. His results for the price momentum and post-earnings announcement drift are consistent with those reported in this paper.

Table 7
Analysis of the Fama–French factors and the winners minus losers momentum portfolio This table presents coefficient estimates from regressions of the Fama–French factors as well as momentum on *PMN* and a January dummy. *MKT* denotes the CRSP value-weighted market excess return. *SMB* and *HML* are the Fama–French factors related to size and book-to-market. *WML* is the winners minus loser momentum portfolio. *PMN* is defined in Table 1. Panel A presents the mean monthly returns on each of the portfolios and Panel B presents the Pearson correlation coefficients across the factors. Panel C presents the factors regression coefficient estimates. The sample is from January 1972 to December 1999. The *t*-statistics are in parenthesis.

Panel A: mean monthly returns on Fama–French factors and momentum (WML)

	MKT_t	SMB_t	HML_t	WML_t	PMN_t	$JANDUM_t$
Mean (%)	0.62	0.07	0.35	0.76	0.90	0.08
t-stat	(2.51)	(0.47)	(2.24)	(2.48)	(7.46)	(5.52)

Panel B: correlation matrix

	MKT_t	SMB_t	HML_t	WML_t	PMN_t
MKT_t	1.00	0.27	−0.45	−0.10	−0.04
SMB_t	0.27	1.00	−0.15	−0.38	−0.34
HML_t	−0.45	−0.15	1.00	−0.18	−0.27
WML_t	−0.10	−0.38	−0.18	1.00	0.66
PMN_t	−0.04	−0.34	−0.27	0.66	1.00

Panel C: regression of monthly Fama–French factor returns on the PMN payoff

	MKT_t			SMB_t			HML_t			WML_t		
	I	II	III	IV	V	VI	VII	VIII	IX	X	XI	XII
INTERCEPT	0.48	0.70	0.50	−0.08	0.47	0.37	0.19	0.65	0.51	1.55	−0.75	−0.07
	(1.86)	(2.61)	(1.72)	(−0.52)	(2.95)	(2.13)	(1.20)	(4.08)	(3.00)	(5.41)	(−2.98)	(−0.26)
PMN_t		−0.08	−0.02		−0.44	−0.41		−0.34	−0.30		1.69	1.46
		(−0.74)	(−0.14)		(−6.58)	(−5.77)		(−5.09)	(−4.22)		(15.99)	(14.01)
$JANDUM_t$	1.71		1.67	1.89		0.86	1.85		1.10	−9.41		−5.73
	(1.92)		(1.77)	(3.38)		(1.52)	(3.42)		(1.98)	(−9.49)		(−6.90)
Adj-R^2(%)	0.79	−0.14	0.50	3.02	11.22	11.57	3.09	6.91	7.72	21.01	43.20	50.16

lowest *SUE* portfolio, P_1, have negative earnings on average, whereas stocks in the highest *SUE* portfolio, P_{10}, have positive average earnings as evidenced by the earnings-to-price ratio. (Hence the notation *PMN* for positive minus negative.) The book-to-market ratio of the P_1 portfolio is 1.05 while that of P_{10} is 0.76. In other words, the P_{10} portfolio behaves more like a growth portfolio and the P_1 portfolio behaves more like a value portfolio. Also, the P_{10} portfolio stocks are larger as measured by market capitalization and have higher prices than the P_1 portfolio stocks, confirming the negative correlation between *SMB* and *PMN*. While the book-to-market and size effects might suggest that the returns to *PMN* are negative, the impact of momentum overwhelms the size and the book-to-market effects. The returns in the six months prior to the portfolio formation month are significantly different across the two portfolios. The P_1 portfolio has an average past six-month return of −0.87% whereas the P_{10} portfolio has an average past six-month return of 15.74%. Thus, consistent with the momentum return classifications, the P_{10} portfolio returns are higher than the P_1 portfolio returns.

Table 8
Average characteristics across *SUE* portfolios The *SUE* decile portfolios are formed as in Table 1 and are held for the following six-month period. The characteristics are obtained for each month for each portfolio by averaging the relevant variable across all stocks in that portfolio in that month. These means are then averaged across months; the table reports the time-series averages. P_1 is the lowest *SUE* portfolio. P_{10} is the highest *SUE* decile portfolio. MID consists of all other portfolios. All accounting values (namely, book value of equity, earnings, and total assets) are taken from most recent quarter that ends at least four months prior to the formation month. All market variables (namely, market value of equity, turnover, volume, and price) are as of the end of quarter prior to formation month. The variable definitions are as follows: *BM* is the book-to-market ratio, *SIZE* represents market capitalization, *EP* represents the earnings-to-price ratio, *PRICE* is the share price, and $RET_{t-6,t-1}$ is the cumulative return in the six-months prior to formation month. The table also presents the *p*-value from *F*-tests of equality across the portfolios.

	P_1	*MID*	P_{10}	H_0: $P_1 = MID$ (*p*-value)	H_0: $MID = P_{10}$ (*p*-value)	H_0: $P_1 = P_{10}$ (*p*-value)	H_0: $P_1 = P_{10} = MID$ (*p*-value)
BM	1.05	0.96	0.76	0.00	0.00	0.00	0.00
SIZE ($ million)	1459.11	1352.62	1712.63	0.16	0.00	0.01	0.00
Log (SIZE)	18.84	18.68	19.35	0.70	0.00	0.00	0.00
EP (%)	−3.86	0.77	2.45	0.00	0.00	0.00	0.00
PRICE ($)	24.60	26.71	33.67	0.16	0.00	0.00	0.00
$RET_{t-6,t-1}$ (%)	−0.87	7.70	15.74	0.00	0.00	0.00	0.00

The firm characteristics of the high- and low-*SUE* portfolio stocks suggest, once again, that price momentum and earnings momentum are strongly related. Since Chordia and Shivakumar (2002), Ahn et al. (2003) and Avramov and Chordia (2005) argue that price momentum is related to the macroeconomy, *PMN* should also be related to the business cycle. In the next section, we study the relation between *PMN* and the macroeconomy.

6. *PMN* and the macroeconomy

PMN could be related to the macroeconomy for two reasons. First, *PMN* could proxy for a risk factor in an intertemporal asset pricing model, such as Merton (1973). In this case, the payoffs to *PMN* would reflect future macroeconomic activities as these activities contain information about the future investment opportunity set of investors. Liew and Vassalou (2000) use this approach to examine whether the Fama–French factors proxy for risk factors. Alternatively, *PMN* could reflect mispricing of macroeconomic variables that impact earnings, or capture the aggregate investor sentiment to corporate earnings, that is, the market's overreaction or underreaction to corporate earnings. To the extent that aggregate investor confidence is related to economic performance, *PMN* would again be related to future macroeconomic conditions.[19]

Following Chen (1991) and Liew and Vassalou (2000), we regress future GDP growth on lagged values of the Fama–French factors as well as *PMN*. The dependent variable is the continuously compounded growth in real GDP over months $t+1$ through $t+12$ and the explanatory variables include the value-weighted excess market return, *SMB*, *HML*, and *PMN*, also compounded over months $t-11$ through t. Since the GDP data are available only at the quarterly frequency, the regressions use quarterly data. Due to overlapping data, the t-statistics are based on the autocorrelation-consistent Newey-West standard errors.

Table 9 presents the results. Over the entire sample period from January 1972 to December 1999, the coefficient on *PMN* is significantly negative, irrespective of whether Fama–French factors are included as additional explanatory variables in the regression. Further, *PMN* tends to be the most significant variable in the regressions. The adjusted-R^2 of the regression is about 29% when *PMN* is included by itself, but this figure increases by only about 3% when Fama–French factors are also added to the regression. The negative coefficient on *PMN* suggests that variation in *PMN* is countercyclical to the business cycle. A countercyclical portfolio should not earn the premium of 90 basis points that is documented in Table 1. Thus, *PMN* is not likely to be related to business cycle risk and is unlikely to be a risk factor.[20]

The coefficient on the value-weighted market return is the only other slope coefficient that is significant in the regressions. The positive coefficient on the market return is consistent with the results documented in Chen (1991). To verify the robustness of our results, we repeat the regressions using GDP growth over months $t+1$ to $t+3$, rather than over 12 months, and we also use returns to *PMN* and Fama–French factors that are compounded over months $t-2$ to t, instead of $t-11$ to t. These modifications yield qualitatively similar results.

Note that our results are in contrast with those in Liew and Vassalou (2000), who find a significantly positive coefficient for *HML*. However, their sample covers the period January

[19] Gervais and Odean (2001) develop a dynamic model in which aggregate investor overconfidence varies with economic performance.

[20] Independently, Young (2001) also shows that an earnings-based factor, similar to *PMN*, is negatively related to future growth in real output.

Table 9
Regression of 12-month-ahead growth in real GDP on 12-month compounded *PMN* and Fama–French factors
This table presents the regression coefficients from regressing growth in real GDP on the Fama–French factors and *PMN*. *PMN* is the zero-investment portfolio that is long stocks with the highest earnings changes and short stocks with the lowest earnings changes. The earnings change portfolios are as defined in Table 1. *MKT* is the CRSP value-weighted market return in excess of the risk-free rate. *SMB* and *HML* are the Fama–French zero-investment portfolios. The regression uses quarterly data, since data on GDP is available only on a quarterly basis. The dependent variable is the continuously compounded growth in real GDP over months t to $t+12$. Since the regressions use overlapping data, the t-statistics, which are reported in parentheses, are based on Newey-West standard errors. This table reports the regression results for the full sample period as well as for the sample period January 1978 to December 1996.

	Jan 1972–Dec 1999			Jan 1978–Dec 1996		
	I	II	III	IV	V	VI
INTERCEPT	4.71	2.36	4.19	4.16	0.07	3.12
	(10.87)	(5.04)	(7.46)	(6.38)	(3.05)	(4.41)
$PMN_{t-11,\ t}$	−0.18		−0.15	−0.15		−0.10
	(−5.09)		(−3.72)	(−3.24)		(−2.06)
$MKT_{t-11,\ t}$		0.06	0.03		1.68	0.05
		(3.95)	(2.13)		(4.33)	(2.47)
$SMB_{t-11,\ t}$		0.00	−0.01		−0.01	−0.01
		(0.12)	(−0.32)		(−0.17)	(−0.40)
$HML_{t-11,\ t}$		0.03	0.01		0.08	0.05
		(0.89)	(0.33)		(3.91)	(2.22)
No. of obs	108	108	108	75	75	75
Adj-R^2(%)	28.74	15.18	31.95	21.16	20.55	26.73

1978 through December 1996. We are able to replicate their result for *HML* over their sample period but not over our entire sample period from January 1972 to December 1999.

In Table 10 we test the robustness of the above relation between GDP growth and returns to the *PMN* portfolio by using alternative measures for future business conditions. Specifically, this table presents results from regressing future values of growth in industrial production (IPG), consumption growth (RCG), growth in labor income (RLIG), inflation, and the three-month T-bill return on the Fama–French factors and *PMN*. In Panel A of Table 10 we use 12-month-ahead data for the dependent variables. Due to overlapping regressions, the t-statistics are based on Newey-West standard errors. When the dependent variable is growth in industrial production, consumption growth, and growth in labor income, the coefficient on *PMN* is significantly negative and is consistent with the results of Table 9. The coefficient on *PMN* is positive when inflation or the T-bill return is the dependent variable. The impact of the Fama–French factors, particularly *SMB* and *HML*, on IPG, RCG, and RLIG is essentially zero.[21] Panel B of Table 10 repeats the above

[21]We use *WML* instead of *PMN* in Tables 9 and 10 and find that *WML* has no impact on future GDP growth, industrial production growth, growth in real consumption, growth in labor income, and inflation. *WML* does impact future nominal T-bill returns.

regression for nonoverlapping data and in Panel C of Table 10 we repeat the above exercise for regressions of three-month-ahead economic activity on three-month-lagged Fama–French factors and *PMN*. The conclusions are essentially unchanged.

Table 10
Regression of future economic activity on *PMN* and Fama–French factor returns This table presents coefficient estimates from regression of macroeconomic variables on *PMN*, *MKT*, *SMB*, and *HML*. *PMN* is the zero-investment portfolio that is long stocks with the highest earnings changes and short stocks with the lowest earnings changes. The earnings change portfolios are as defined in Table 1. *MKT* is the CRSP value-weighted market return in excess of the risk-free rate. *SMB* and *HML* are the Fama–French zero-investment portfolios. $IPG_{t,t+k}$ represents the growth in industrial production over months t to $t+k$, measured as the change in the log of total production. $RCG_{t,t+k}$ is the growth in real consumption over months t to $t+k$, measured as the change in the log of seasonally adjusted real consumption of services and nondurable goods. $RLIG_{t,t+k}$ is the growth in real labor income over months t to $t+k$, measured as the change in the log of personal income from wages and salaries, seasonally adjusted, minus inflation. $INF_{t,t+k}$ represents inflation over months t to $t+k$, measured as the change in the log of CPI and is not seasonally adjusted. $TBR_{t,t+k}$ represents nominal Treasury bill returns in months $t+1$ to $t+k$, measured as continuously compounded returns on one-month Treasury-bills. Panel A presents results from overlapping regressions; there are 336 observations in each regression. The t-statistics are corrected for serial correlation using Newey-West standard errors (with number of lags = 12). Panel B uses non-overlapping annual data and has 28 observations in each regression. Panel C uses nonoverlapping quarterly data and has 112 observations in each regression.

Panel A: regression of 12*-month-ahead economic activity on PMN and Fama–French factor returns*

	$IPG_{t,t+12}$		$RCG_{t,t+12}$		$RLIG_{t,t+12}$		$INF_{t,t+12}$		$TBR_{t,t+12}$	
	I	II	III	IV	V	VI	VII	VIII	IX	X
INTERCEPT	3.21	3.12	3.11	3.09	2.26	2.22	4.95	4.91	6.50	6.47
	(4.44)	(4.67)	(13.23)	(13.17)	(4.88)	(5.12)	(8.76)	(9.16)	(14.58)	(15.31)
PMN_t	−0.35	−0.37	−0.11	−0.11	−0.15	−0.18	0.12	0.20	0.11	0.17
	(−4.19)	(−3.37)	(−2.98)	(−2.89)	(−2.43)	(−2.57)	(1.86)	(2.58)	(1.83)	(2.25)
MKT_t		0.21		0.05		0.12		−0.09		−0.05
		(2.53)		(2.70)		(2.66)		(−2.61)		(−1.83)
SMB_t		−0.01		−0.01		−0.07		0.18		0.12
		(−0.12)		(−0.20)		(−1.09)		(2.18)		(2.01)
HML_t		−0.08		−0.01		−0.01		0.02		0.02
		(−0.77)		(−0.25)		(−0.17)		(0.32)		(0.27)
Adj-R^2(%)	2.51	7.05	2.20	3.82	1.13	3.96	0.47	3.01	0.70	1.84

Panel B: regression of 12*-month ahead economic activity on* 12*-month compounded PMN and Fama–French factors*

	$IPG_{t,t+12}$		$RCG_{t,t+12}$		$RLIG_{t,t+12}$		$INF_{t,t+12}$		$TBR_{t,t+12}$	
	I	II	III	IV	V	VI	VII	VIII	IX	X
INTERCEPT	5.72	4.62	4.08	4.53	3.96	3.16	3.65	3.49	5.58	4.83
	(5.18)	(2.82)	(10.01)	(7.64)	(5.02)	(2.74)	(3.98)	(2.70)	(7.87)	(5.04)
$PMN_{t-11,t}$	−0.29	−0.22	−0.11	−0.13	−0.18	−0.14	0.13	0.17	0.10	0.17
	(−3.46)	(−1.97)	(−3.46)	(−3.24)	(−3.06)	(−1.75)	(1.92)	(1.87)	(1.89)	(2.59)
$MKT_{t-11,t}$		0.06		−0.03		0.05		−0.03		0.00
		(1.01)		(−1.36)		(1.35)		(−0.73)		(0.12)
$SMB_{t-11,t}$		0.01		0.00		−0.02		0.10		0.10
		(0.13)		(0.10)		(−0.35)		(1.82)		(2.43)
$HML_{t-11,t}$		0.02		0.00		0.01		−0.01		−0.01
		(0.33)		(−0.09)		(0.21)		(−0.13)		(−0.34)
Adj-R^2(%)	28.92	23.12	28.91	26.42	23.64	20.81	9.07	11.93	8.75	18.59

Table 10 (*continued*)

Panel C: regression of 3*-month ahead economic activity on* 3*-month compounded PMN and Fama–French factors*

	$IPG_{t,t+3}$		$RCG_{t,t+3}$		$RLIG_{t,t+3}$		$INF_{t,t+3}$		$TBR_{t,t+3}$	
	I	II	III	IV	V	VI	VII	VIII	IX	X
INTERCEPT	1.14	0.96	0.85	0.83	0.64	0.50	1.19	1.20	1.58	1.56
	(5.86)	(4.30)	(12.80)	(10.84)	(3.45)	(2.28)	(11.77)	(10.69)	(22.15)	(18.81)
$PMN_{t-2,t}$	−0.16	−0.12	−0.04	−0.04	−0.04	−0.03	0.03	0.05	0.02	0.04
	(−3.83)	(−2.44)	(−2.56)	(−2.37)	(−1.07)	(−0.58)	(1.40)	(1.89)	(1.62)	(2.14)
$MKT_{t-2,t}$		0.06		0.02		0.05		−0.04		−0.01
		(2.43)		(2.62)		(2.06)		(−3.14)		(−1.62)
$SMB_{t-2,t}$		0.01		−0.02		−0.04		0.06		0.03
		(0.35)		(−2.02)		(−1.09)		(3.40)		(2.43)
$HML_{t-2,t}$		−0.01		0.00		0.04		−0.01		0.00
		(−0.17)		(0.18)		(1.29)		(−0.44)		(0.37)
Adj-R^2(%)	10.97	16.84	4.78	10.22	0.14	1.72	0.84	12.27	1.43	5.44

Although our evidence suggests that *PMN* is related to the macroeconomy, the negative correlation between *PMN* and real activity (namely, GDP growth, IPG, RCG, and RLIG) is inconsistent with *PMN* being a proxy for macroeconomic risk. So, why is *PMN* related to the business cycle?

Chordia and Shivakumar (2005) offer one potential explanation, arguing that underreaction to earnings surprises partly results from an "inflation illusion." According to the inflation illusion argument, first proposed by Modigliani and Cohn (1979), stock market investors fail to incorporate the effect of inflation on nominal earnings growth rates when valuing stocks. Thus, when inflation rises, investors do not adjust the future earnings growth, even though they fully adjust the discount rates. A direct implication of this hypothesis is that, if the earnings growth in response to inflation varies across stocks, inflation illusion would induce misvaluation in the cross-section: Stocks whose earnings growth are positively related to inflation would be undervalued, while those with earnings growth negatively related to inflation would be overvalued. Chordia and Shivakumar (2005) show that the effect of inflation on earnings growth increases monotonically across the *SUE* portfolios and that this cross-sectional variation in the exposure of earnings to inflation along with the inflation illusion causes the post-earnings announcement drift.

7. Conclusions

Two robust and persistent anomalies over the last four decades that have defied rational explanations are the post-earnings announcement drift, or earnings momentum, and the short-run return continuations, or price momentum. In this paper we ask whether the two are related. A zero-investment portfolio (denoted *PMN*) that is long the highest earnings surprise portfolio and short the lowest earnings surprise portfolio captures the price momentum phenomenon in time-series and cross-sectional asset pricing tests. *PMN* is formed on the basis of stock portfolios that are sorted on the firm characteristic, *SUE*, which is defined as the seasonal change in earnings standardized by its standard deviation in prior quarters. We confirm the result of Chan et al. (1996) that although earnings surprises and past returns are related, they have separate explanatory power for future returns. The contribution of this paper is that the characteristics based factor *PMN* captures price momentum in asset pricing tests. Thus, while the two characteristics, *SUE* and past returns, have independent explanatory power for future returns, the factor loadings on the common factor *PMN* capture the impact of past returns on future returns. Our results support the argument that price momentum is primarily subsumed by the systematic component of earnings momentum and that price momentum is merely a manifestation of the earnings momentum.

The return on *PMN* is correlated with future growth in GDP, industrial production, consumption, labor income, inflation, and T-bill returns. These correlations persist even after controlling for the Fama–French factors. Interestingly, we find that *PMN* has a greater predictive power for future business conditions than the Fama–French factors. Because *PMN* is related to business cycle conditions, this possibly explains why it captures price momentum in asset pricing tests. These findings point to a systematic component of the post-earnings announcement drift that reflects fundamental macroeconomic information not contained in any of the commonly used stock return factors.

On average, *PMN* returns are about 90 basis points a month. Moreover, *PMN* is countercyclically related to the business cycle. This suggests that *PMN* is unlikely to be a risk factor in an asset pricing context. In any case, it is interesting to find that *PMN* captures the price momentum phenomenon. The fact that price momentum is subsumed by a common factor that is related to the macroeconomy suggests little role for idiosyncratic component of returns in explaining price momentum. Our findings help narrow the search for an explanation for price momentum and suggest that theories that explain price momentum must also explain the role of *PMN* in stock returns.

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