Confounded Factors[*]

Joseph Gerakos   Juhani T. Linnainmaa

March 15, 2013

Abstract

Book-to-market (BE/ME) ratios explain variation in expected returns because they correlate with recent changes in the market value of equity. Although the remaining variation in BE/ME ratios captures comovement among stocks, it does not predict returns. Therefore, the HML factor is a sum of two parts: one with a positive price of risk ("priced part") and the other with a zero price of risk ("unpriced part"). The unpriced part confounds the HML factor and distorts inferences. First, portfolio managers can exploit the unpriced part—a portfolio long the priced and short the unpriced part has an annual three-factor model alpha of 7.7%. Second, the three-factor model subsumes the earnings-to-price and cashflow-to-price anomalies only because these anomalies covary with the HML's unpriced part. Third, the unpriced part leads to downwardly biased estimates of money managers' skill. The problem of confounded factors applies to all empirical risk factors.

1   Introduction

Empirical asset pricing models assume that factors have uniform risk premia. The problem of confounded factors arises when an empirical factor combines multiple factors with different risk premia. We examine the implications of this problem in the context of the HML factor, which is constructed by sorting stocks into portfolios by their book-to-market (BE/ME) ratios. We show that systematic variation in BE/ME ratios can be broken into two components: one with a positive price of risk and the other with a zero price of risk. The priced component is the part of BE/ME ratios that correlates with recent (approximately five years) changes in the market value of equity. The unpriced component is everything else. The unpriced component confounds the HML factor and distorts inferences.

The problem of confounded factors can be illustrated in terms of the arbitrage pricing theory (Ross 1976). Suppose the true model describing asset returns is

$$\tilde{r}_{it} - r_{ft} = \beta_{1,i}\tilde{F}_{1t} + \beta_{2,i}\tilde{F}_{2t} + \tilde{\varepsilon}_{it}, \tag{1}$$

in which the risk premia are $\lambda_1 = 5\%$ and $\lambda_2 = 0\%$, and the factors are orthogonal and have equal variances. What happens if an econometrician uses $\tilde{F}_t = \tilde{F}_{1t} + \tilde{F}_{2t}$ as the factor? If a stock's true betas are $\beta_{1,i} = 1, \beta_{2,i} = 0$, then its risk premium is $\mathrm{E}(\tilde{r}_{it}) - r_{ft} = (1)5\% + (0)0\% = 5\%$. But its beta against $\tilde{F}_t$ is $\frac{\mathrm{cov}(\tilde{r}_{it}-r_{ft},\tilde{F}_{1t}+\tilde{F}_{2t})}{\mathrm{var}(\tilde{F}_{1t}+\tilde{F}_{2t})} = \frac{1}{2}(\beta_{1,i} + \beta_{2,i}) = \frac{1}{2}$, so the econometrician would estimate the stock's risk premium as $(\frac{1}{2})(\lambda_1 + \lambda_2) = 2.5\% < 5\%$. Similarly, if the betas are $\beta_{1,i} = 0, \beta_{2,i} = 1$ instead, the stock's risk premium is now 0%, but the econometrician's estimate is unchanged, $2.5\% > 0\%$. The econometrician draws correct inferences only if $\beta_{1,i} = \beta_{2,i}$.

Our main result is that HML behaves very much like the composite $\tilde{F}$ in this example, with one component's risk premium significantly positive and the other's close to zero. We call the first

component the "priced" component and the other "unpriced." The composite nature of HML is a problem for multi-factor models. HML has just one price of risk, $\hat{\lambda}_{\text{hml}}$, but a strategy can covary with HML not only because $\beta_{1,i}$ in equation (1) is high, but also because $\beta_{2,i}$ is high. Although we use the HML factor to illustrate the confounded factors problem, our message is more general. If we as econometricians use composite risk factors that confound primitive risk factors with different prices of risk, we misestimate assets' risk premia. This issue applies not just to HML, but to any empirical risk factor such as size, momentum, or liquidity—we do not know whether these factors have uniform risk premia.

In four empirical applications we demonstrate the distortions that arise from confounded factors in the context of book-to-market ratios. First, we show that managers can cheat the three-factor model by trading the unpriced variation in book-to-market ratios. If stocks are sorted by the unpriced part of the BE/ME ratio, the average return and CAPM alpha on the high-minus-low strategy are close to zero. But the unpriced part correlates with the HML factor. Hence, this hedge portfolio's three-factor model alpha is -47 basis points per month ($t = -3.5$) and the regression's R^2 is 34%. A strategy that purchases stocks with high "true" BE/ME ratios and sells those with low "false" BE/ME ratios has a three-factor model alpha of 7.7%. This strategy is noteworthy because it chooses stocks based on just BE/ME ratios, yet the three-factor model fails because the long-short portfolio barely correlates with the HML factor.

Second, the presence of the unpriced component suggests that the 25 size- and BE/ME-sorted portfolios should not be used to evaluate asset pricing models. We show that the when moving from growth to value, the loadings on the unpriced component of the HML increase as fast (or even faster) than those on the priced component. Because both sets of loadings form a smooth surface, a model can price these portfolios not only by covarying with their priced components but also by covarying

with their unpriced components.

Third, the unpriced component distorts estimates of skill among money managers. We show that on average managers' strategies covary positively with the unpriced components thereby leading to downwardly biased estimates of skill based on the traditional three-factor model. When we estimate skill using a multi-factor model that includes separate factors for the priced and unpriced components, we find a significant shift to the right in the distribution of alphas.

Fourth, the unpriced component of the HML distorts inferences about anomalies. In particular, this unpriced component is shared by other price-scaled variables as well, such as earnings-to-price, cashflow-to-price, and dividend-to-price ratios. Hence, the three-factor model appears to explain the E/P and C/P anomalies, but only because it assigns the HML's one price of risk to these anomalies' covariation with the unpriced component. When we adjust the E/P and C/P ratios so they only reflect information accrued into these ratios over the past five years, the anomalies reappear with sizable three-factor model alphas: 50 basis points per month on an E/P strategy (t-value $= 3.06$) and 32 basis points per month on a C/P strategy (t-value $= 2.06$).

To identify the priced and unpriced components in HML, we decompose book-to-market ratios. Our decomposition relates to studies by Daniel and Titman (2006) and Fama and French (2008). Similar to those studies, we start from the identities governing the time series evolution of BE/ME ratios: one describes changes in the book value of equity and the other changes in the market value of equity. Although other decompositions are possible, changes in the book and market values of equity naturally correspond with theoretical models that differentiate between tangible and intangible assets. More generally, this decomposition demonstrates the problem of confounded risk factors. Our analysis does not assume that we capture the "exact" primitive risk factors.

Because the value premium is based on how the BE/ME ratio sorts stocks into portfolios, what

matters is the extent to which these components drive variation in BE/ME ratios. Although at the five-year mark changes in the market value of equity explain less than two-thirds of the variation in BE/ME ratios, they provide all of the predictive power. The predictive power of recent changes in the market value of equity is easy to demonstrate. Regressions (1), (4), and (7) in Table 1 report Fama-MacBeth regressions of monthly returns against size and BE/ME ratios for all stocks, All but Microcaps, and Microcaps. The BE/ME ratio is statistically significant among both large and small stocks. Regressions (2), (5), and (8) show that when the regressions also control for the five-year change in the market value of equity, BE/ME ratios lose their significance.[1] A disaggregation of the changes in the market value of equity (regressions (3), (6), and (9)) pushes the slopes on the BE/ME ratios even closer to zero. This result is not specific to the U.S. equity market. We show similar results for Japan—a market that is often used for out of sample validation of asset pricing models.

Our results do not imply that the changes in fundamentals have no information about expected returns. Novy-Marx (2012), for example, finds that a pure-fundamentals factor, gross profitability, also predicts future returns but is independent of the information contained in book-to-market ratios. Fama and French (2008) show that when they control simultaneously for the changes in the market and book value of equity, net issuances, and firms' historical BE/ME ratios, the two change variables are of equal importance in explaining variation in average returns among large stocks. What our results show is that when the BE/ME ratio is taken in isolation, it leads to an asset pricing factor with a non-uniform price of risk. The other information *is* useful for modeling expected returns—but it has to be introduced by adding new factors.

When we decompose book-to-market ratios why are only recent changes in the market value of

---

[1]We compute the five-year changes in the market value of equity up to December in year $t-1$ to match the timing of the ME term in BE/ME ratios. We use the same timing conventions as Fama and French (1992) to construct BE/ME ratios.

equity informative about expected returns? One explanation starts by delineating the firm's assets into tangible and intangible assets.[2] Assume that tangible asset values change infrequently and are captured noisily by the book value of equity. Furthermore, assume that intangible asset values change frequently and are captured accurately in the market value of equity but not by the firm's book value.

If tangible and intangible assets have different risk exposures, then the current book-to-market ratio predicts returns because it proxies for the composition of the firm's assets. Moreover, changes in the market value of equity accurately capture changes in the firm's asset base and hence are informative about risk exposures. In contrast, book values of equity contribute to the overall variation in book-to-market ratios but are less informative about risk exposures because they are only a noisy measure of tangible assets. In models with such a mechanism, a factor based on BE/ME ratios would suffer from the same confounded factor problem that we find in the data.

2   Data

We take stock returns from CRSP and accounting data from Compustat. Our sample starts with all firms traded on NYSE, Amex, and Nasdaq. For these firms, we calculate the book value of equity (shareholder equity, plus balance sheet deferred taxes, plus balance sheet investment tax credits, minus preferred stock). We set missing values of balance sheet deferred taxes and investment tax credit equal to zero. To calculate the value of preferred stock, we set it equal to the redemption value if available, or else the liquidation value, or the carrying value. If shareholders' equity is missing, we set it equal to the value of common equity if available, or total assets minus total liabilities. We then use the Davis, Fama, and French (2000) book values of equity from Ken French's website to fill in missing values of

---

[2]See, for example, Berk, Green, and Naik (1999), Gomes, Kogan, and Zhang (2003), and Carlson, Fisher, and Giammarino (2004) for such models.

the book value of equity. Because we require the return on equity, we start our sample in January 1963. We end it in December 2011.

For BE/ME, we use the market value of equity as per the fiscal year end and calculate it as the CRSP month-end share price times the Compustat shares outstanding if available, or else the CRSP shares outstanding. We follow prior research and lag BE/ME ratios by at least six months so that companies have released their annual financial statements. For example, if a firm's fiscal year ends in December, we begin using the December information at the end of June. When we calculate the BE/ME, E/P, C/P, and D/P ratios, we align the numerator, the denominator, and all the components of the decomposition at the same point in time. For example, the five-year changes in the book and market values of equity are the five-year change up to the date when the BE/ME ratio is computed. Thus, our decompositions are exact.[3]

3 Decomposing BE/ME ratios

The BE/ME ratio can be decomposed using the following algebraic identity:

$$bm_t \equiv bm_{t-k} + \sum_{\tau=t-k+1}^{t} dbe_\tau - \sum_{\tau=t-k+1}^{t} dme_\tau, \tag{2}$$

where bm_t is the log-BE/ME ratio at time t, $dbe_t = \ln(\text{BE}_t/\text{BE}_{t-1})$ is the change in the book value of equity, and $dme_t = \ln(\text{ME}_t/\text{ME}_{t-1})$ is the change in the market value of equity. This identity implies that in a regression of bm_t against the components on the right-hand side of equation (2), the slopes on bm_{t-k} and dbe_τs would equal one and those on dme_τs would equal negative one.

---

[3]We trim BE/ME ratios at 0.005 and 0.995 levels each month to discard outliers. The literature alternatively winsorizes or trims outlying BE/ME ratios. We trim because winsorization breaks the exact decomposition identities. See Fama and French (2008, footnote 1).

Components of bm_t can, however, differ significantly from each other in their contribution to the *variation* of current BE/ME ratios. This variation is what ultimately matters when we sort stocks by BE/ME ratios to measure value and growth or to construct the HML factor. Because stock returns are more volatile than accounting variables, changes in the market value of equity drive more of the cross-sectional variation in BE/ME ratios. Changes in the market and book values of equity are also not independent of each other: a change in the book value of equity usually reflects in market valuations, either contemporaneously or at lead or lag.

3.1 Which components explain cross-sectional variation in BE/ME ratios?

Our cross-sectional decomposition of BE/ME ratios starts from the identity that a variable's covariance with itself equals its variance:

$$
\begin{aligned}
\mathrm{var}(bm_t) &= \mathrm{cov}(bm_t, bm_{t-k} + \sum_{\tau=t-k+1}^{t} dbe_\tau - \sum_{\tau=t-k+1}^{t} dme_\tau) \\
&= \mathrm{cov}(bm_t, bm_{t-k}) + \sum_{\tau=t-k+1}^{t} \mathrm{cov}(bm_t, dbe_\tau) + \sum_{\tau=t-k+1}^{t} \mathrm{cov}(bm_t, -dme_\tau).
\end{aligned} \tag{3}
$$

Dividing both sides of this equation through by $\mathrm{var}(bm_t)$ gives each term's percentage contribution to the variance of today's BE/ME ratios.

Our use of the term "variance decomposition" is consistent with its usage in prior research.[4] Our decompositions, however, measure the covariation between today's BE/ME ratios and their components, and can therefore be negative. These covariances have the same interpretation as the analysis of Fama and French (1995), who, for example, show that value firms experienced low profitability for the prior five years. These estimates tell us what type of firms end up in different portfolios when sorted

---

[4]See Cochrane (1992).

by their BE/ME ratios. In particular, if a component's covariance with today's BE/ME ratios is zero, then any information contained in this component is lost in the univariate portfolio sorts, because this component does not vary across portfolios.[5]

Table 2 Panel A presents the variance-decomposition estimates. We estimate the covariances of equation (3) with year fixed effects. The main result here is that most of the variation in BE/ME ratios arises from lagged BE/ME ratios and changes in the market value of equity. In the one-year decomposition, 81.84% of the variation is due to the prior year's BE/ME ratio, 23.1% is due to (minus) the change in the market value of equity, and the rest, -4.93%, is due to the change in the book value of equity. The negative sign on the change in the book value indicates that when the book value of equity increases, the market value of equity generally increases even more, thereby resulting in lower BE/ME ratios in the cross-section.[6]

The importance of one-year changes in the book and market values of equity decreases from year to year. This result may seem unexpected. If a random variable is decomposed into k *independent* components with equal variances, then each component contributes $1/k$th of the overall variance. The importance of one-year changes must then decrease because the changes in the book and market values of equity are significantly autocorrelated or cross-serially correlated. If the market value of equity increases in year $t - k$, Table 2 suggests that this increase is often offset by increases in the book value of equity in years $t - k$, $t - k + 1$, ..., or, alternatively, by decreases in the market value of equity in years $t - k + 1 \leq \tau \leq t - 1$.

The covariances between the changes in the book and market value are such that changes in market

---

[5]This argument relates to Lewellen, Nagel, and Shanken's (2010) critique of using 25 size- and BE/ME-sorted portfolios to test asset pricing models. They argue that a sort of stocks in these two dimensions imposes a rigid factor structure—whatever asset pricing factors were in stock returns prior to sorting are largely gone after the sort.

[6]The covariance term $\mathrm{cov}(bm_t, dbe_\tau)$ can be written as the variance of dbe_τ and its covariances with all other terms of the decomposition. The results here suggest that the sum of these other covariances is large enough to more than offset dbe_τ's own variance. These results are similar to that observed in the price-dividend ratio decompositions where future returns appear to account for more than 100% of the variation in the price-dividend ratios (Cochrane 2005, p. 400).

values of equity older than five years contribute negligible amounts of variation to today's BE/ME ratios. At this horizon, over half of the variation is due to the old BE/ME ratios, 63% is due to the cumulative changes in market value of equity, and the difference (-15%) is due to cumulative changes in the book value of equity.

These decomposition results address the question of what information BE/ME ratio-sorted portfolios pick up from the data. The algebraic decomposition could lead one to believe that, because bm_t is the cumulation of past changes in the book and market values, every part of the history plays as prominent a role. Table 2 rejects this view. The BE/ME ratios observed in the cross-section today are almost entirely due to what these ratios were in the past plus an adjustment for the changes in the market value of equity during the intermittent years. Thus, a sort on today's BE/ME ratios mostly sorts stocks by their old BE/ME ratios and the changes in the market value of equity.

3.2 Which BE/ME components predict future returns?

The BE/ME ratio correlates with future returns because some or all of its components correlate with those returns. A regression of stock returns against a firm's log-market capitalization (me_t) and log-BE/ME ratio (bm_t) constrains the regression slopes on the BE/ME components to equal each other:

$$
\begin{aligned}
r_{j,t+1} &= b_0 + b_1 me_{j,t} + b_2 bm_{j,t} + e_{j,t+1} \\
&= b_0 + b_1 me_{j,t} + b_2 bm_{j,t-k} + b_2 \left(\sum_{\tau=t-k+1}^{t} dbe_\tau \right) + b_2 \left(- \sum_{\tau=t-k+1}^{t} dme_\tau \right) + e_{j,t+1}
\end{aligned}
\tag{4}
$$

where the second row replaces today's BE/ME ratio with its decomposition.

We use regression (4) to assess which components of the BE/ME ratio contribute to its ability

to predict future returns in the absence of all other conditioning information except size. We first estimate a baseline specification, which includes only the log-size and today's BE/ME ratio. Let b^* denote the estimated slope on today's BE/ME ratio from this baseline specification. We then estimate univariate regressions of these first-stage residuals against each component. This second-stage regression measures how close each component's optimal slope $(\hat{b}_j)$ is to the common slope (b^*) when all other components' slopes remain fixed at b^*. As an illustration, suppose the slope on the BE/ME ratio is 0.25 in the baseline specification and that the slope from regressing the residuals against component j is -0.2. This estimate suggests that if all other components' slopes are kept at 0.25, then a slope of $\hat{b}_j = 0.25 - 0.2 = 0.05$ on component j maximizes the amount of variance the model explains.

We first test whether the component's optimal slope differs from zero, $H_0 : \hat{b}_j = 0$. In the example above, the test is whether $\hat{b}_j = 0.05$ is statistically different from zero. Second, we test whether the component's optimal slope differs from the common slope, $H_0 : \hat{b}_j = b^*$. In the example above, the test now is whether -0.2 differs from zero. These two sets of t-values measure whether, at a first approximation, one would be better off excluding a variable from the BE/ME ratio when we do not condition on any additional information besides size. Component j's optimal slope maximizes the regression's predictive power. Therefore, any deviation from this value hurts the model. If for a specific component the zero benchmark is statistically closer than the common-slope benchmark, then the regression suggests that the BE/ME ratio performs worse than it would with the component removed from the ratio.

Table 3 Panel A shows the baseline regression, which is the same as the regression in column (1) of Table 1. Panel B decomposes the BE/ME ratio into three parts: the BE/ME ratio five years ago, and the cumulative log-changes in the book and market values of equity. The three univariate regressions

summarized here show that the optimal slopes on these three components are very different: 0.133 (bm_{t-k}), 0.087 $(\sum dbe_\tau)$, and 0.439 $(-\sum dme_\tau)$. The first column of t-values shows that whereas the change in the market value of equity is significant (t-value of 4.98), the other two are not. Although the slope on $\sum dme_\tau$ exceeds the common slope of 0.261 from Panel A, the other components' optimal slopes are closer to zero. The significant t-values here (-2.80 and -4.04) suggest that the BE/ME ratio would perform better in Fama-MacBeth regressions after excluding one (or possibly both) of these components from the ratio.[7]

Panel C disaggregates log-changes in the book and market values of equity by year. The message here is similar to that in Panel B. None of the one-year changes in the book value of equity are significantly different from zero and, except in one case, the estimated slope is statistically closer to zero than to the common slope. These computations suggest that today's BE/ME ratio, as an explanatory variable in Fama-MacBeth regressions only with log-size, does not derive its power from either old BE/ME ratios or changes in the book value of equity. The slopes on the changes in the market value of equity are significantly above both the zero and common-slope benchmarks.

Table 3 Panel D further decomposes the changes in the book and market values of equity into three components each. The change in the book value of equity arises from (1) the log-return on equity (roe) with tax adjustments, (2) dividends, which we measure by the difference in the with- and without-dividends log-returns on equity $(r^b_{\text{div},\tau})$, and (3) the remainder term, which represents net issuance $(r^b_{\text{iss},\tau})$ and is the difference between the log-change in the book value of equity and the without-dividends log-return on equity. Similarly, the change in the market value of equity is traceable

---

[7]Asness and Frazzini (2011) find that adjusting BE/ME ratios to use more timely price data increases the value premium. Their results are consistent with those in Table 3. Our Fama-MacBeth regressions indicate that the BE/ME ratio would perform better without the book value of equity because its optimal slope is closer to zero than to the common slope. Asness and Frazzini (2011) increase the value premium by using more recent market values of equity, thereby reducing the role that recent changes in the book value of equity play in the BE/ME ratios. In fact, one obtains a similar result not only by using more timely records of market values of equity, but also by using *older* records of book value of equity.

to (1) total stock return, r_τ, (2) the dividend yield, $r_{\mathrm{div},\tau}$, measured as the difference in with- and without-dividends stock returns, and (3) the remainder term, which represents net issuance ($r_{\mathrm{iss},\tau}$) and is the difference between the log-change in the market value of equity and the without-dividends stock return. Year-τ change in the BE/ME ratio can thus be written as

$$dbm_\tau = dbe_\tau - dme_\tau = (\mathrm{roe}_\tau - r_\tau) - \left(r^b_{\mathrm{div},\tau} - r_{\mathrm{div},\tau}\right) + \left(r^b_{\mathrm{iss},\tau} - r_{\mathrm{iss},\tau}\right). \tag{5}$$

Equation (5) gives the appearance that the change in the BE/ME ratio is approximately equal to the difference in returns on book and market equity, roe_τ and r_τ. This approximation, however, is only accurate if a firm's pre-issuance (or pre-dividend) BE/ME ratio is close to one. Otherwise, equity issuances, for example, pull BE/ME ratios toward one.

It is easy to summarize Panel D's results on the book side. The optimal slopes on the return on equity are positive but insignificant and the slopes on book dividend yields, although large in magnitude, are noisy and not significant either way. The first two optimal slopes on the book issuance variables are significant but in the wrong direction, and therefore work *against* the BE/ME ratio in predicting future returns. These net issuance variables are thus also responsible for the slightly negative slopes on the first two changes in the book value of equity in Panel B. With respect to the market-side variables, stock market returns and net issuances contribute power to the BE/ME ratio. In fact, the slopes on all net-issuance variables are significantly above the common slope. The market-side dividend yield variables, similar to their book-side counterparts, are thoroughly insignificant.

Table 3 thus indicates that when we estimate regressions of returns on BE/ME ratios the significantly positive slope can be traced back to past stock returns and net issuances. The other variables do not contribute to the BE/ME ratio's ability to predict returns.

4 The priced component of BE/ME ratios

4.1 Fama-MacBeth regressions

Table 4 Panel A reports Fama-MacBeth regressions that use different variables to capture information in BE/ME ratios. We use five alternative definitions of value: bm_{t-5} is the BE/ME ratio from five years ago; $-\sum dme_\tau$ is minus the change in the market value of equity over the past five years; $\sum dbe_\tau$ is the change in the book value of equity over the past five years; $-\sum r_\tau$ is minus the five-year stock returns; $\widehat{bm}_t$ is the part of bm_t that is due to the change in the market value of equity over the past five years—we call $\widehat{bm}_t$ the priced component.

We construct $\widehat{bm}_t$ by first computing for each monthly cross-section the covariances between current BE/ME ratios and $X_t = [dme_t \cdots dme_{t-5}]$. We denote the column vector of these five covariances by C_t. We then compute the cross-sectional covariance matrix of X_t, denoted by Σ_t^X, and construct a vector of weights $w_t = (\Sigma_t^x)^{-1} C_t$. The alternative repackaging $\widehat{bm}_t$ is then $\widehat{bm}_t = X_t w_t$. By construction, the covariances of this variable with the changes in the market value of equity are identical to what they are for the BE/ME ratio itself; that is, $\text{cov}(bm_t, dme_{t-k}) \equiv \text{cov}(\widehat{bm}_t, dme_{t-k})$ for $0 \leq k \leq 4$.[8] This construction therefore discards any information not contained in the changes in the market value of equity, and only uses these five variables in the same proportions as they are present in bm_t.[9] We use only the information from each cross-section to construct this variable, and so it uses no future information. Hence, when we sort on $\widehat{bm}_t$, we sort on the changes in the market value of equity to the same extent as when we sort on bm_t. The difference is that $\widehat{bm}_t$ throws out

---

[8]We created a similar variable based on the five-year stock returns and two-year net issuances because the low-level decomposition suggests that these variables may be ultimately responsible for the significance of the changes in the market value of equity. Results using this alternative measure are very similar to $\widehat{bm}_t$ across all tests.

[9]Because $\widehat{bm}_t$ is constructed to replicate the variance due to the changes in the market values of equity, $\widehat{bm}_t$ is not the same as summing the past five years of changes in the market value of equity. For such a sum, in general, $\text{cov}(bm_t, dme_{t-k}) \neq \text{cov}(\sum_{\tau=t-4}^{0} dme_\tau, dme_{t-k})$.

variation that we suspect does not spread returns. Our construction of $\widehat{bm_t}$ is equivalent to projecting

today's BE/ME ratios against a constant and the changes in the market value of equity.

The first row of Table 4 Panel A reports the baseline regression that only includes the market

capitalization and the current BE/ME ratio as the explanatory variables. We report two specifications

for all repackaged value variables, one with and the other without the current BE/ME ratio.

The old BE/ME ratio is not very informative. It is insignificant by itself and significantly negative

when used in conjunction with the current BE/ME ratio. This result suggests that the old BE/ME

ratio differences out fresher information from the current BE/ME ratio. The remaining rows of Table 4

Panel A examine what this fresh information might be. The cumulative log-change in the market value

of equity is significant by itself with a t-value of 5.7 and it retains most of its significance when the

regression also controls for the current BE/ME ratio. Whatever information is in bm_t is primarily

embedded in these log-changes. The next row shows that the same is not true for the changes in the

book value of equity. The sign on this variable is negative because, as Table 2 shows, a firm is more

likely to be a "growth" firm after experiencing positive changes in the book value of equity. When

the regression also includes the current BE/ME ratio, the two variables are of approximately equal

importance.

A comparison of the $-\sum dme_\tau$ and $-\sum r_\tau$ rows shows that past stock returns are not as powerful

predictors of future returns as are changes in the market value of equity. The last row shows that

in a regression of returns against both the current BE/ME ratio and the same ratio stripped out of

all the variation that is not driven by the changes in the market value of equity, bm_t is statistically

and economically insignificant.[10] This result is in contrast to Fama and French (1996), who show that

---

[10]Table 1 shows that, in a multivariate Fama-MacBeth regression of returns on the year-by-year changes in market value of equity, the slopes are all statistically significant but slightly different. This finding explains why the current BE/ME ratio has a higher t-value (although it is statistically insignificant) in the $-\sum dme$ regression than in the $\widehat{bm_t}$ regression. Because bm_t and $-\sum dme$ load differently on the changes in the market values of equity, bm_t helps the first

BE/ME ratios subsume the five-year reversal in returns found by De Bondt and Thaler (1985). We show the opposite: *changes in the market value of equity* subsume BE/ME ratios. As we show below, the difference arises from net issuances.

4.2 Portfolio sorts

Table 4 Panel B shows average excess returns, CAPM alphas, and Fama and French (1993) three-factor model alphas for portfolios sorted by alternative value factors. The results here are similar to the Fama-MacBeth regressions. The portfolios sorted on the five-year changes in the market value of equity spread out returns more than the current BE/ME ratio. The excess-return spread between the top and bottom decile is 50 basis points (t-value = 2.48) per month for bm_t but 56 basis points (t-value = 2.85) for the market-value-of-equity changes. A CAPM adjustment increases the spreads to 52 and 65 basis points. The stripped-down version of bm_t, which uses dme_τs in the same proportions they are present in bm_t, generates similar return spreads.

The average monthly return (35 basis points) and CAPM alpha (39 basis points) are lower for the high-minus-low portfolio based on $-\sum r_\tau$ than they are for the portfolio based on $-\sum dme_\tau$. The largest return differences are in the tails of the distribution. Although the returns on deciles 2 and 9 are similar for these two sorts, those on deciles 1 and 10 are not. This behavior is expected because the difference between the two sorts is (mostly) about net issuances. The unconditional covariation between net issuances and today's BE/ME ratios is modest (see Table 2 Panel B) because issuances are relatively rare. Conditional on issuing or retiring equity, however, a firm's BE/ME ratio can change substantially. Table 4 Panel B shows indirectly that firms with recent net issuances are mostly in deciles 1 or 10 when firms are sorted by their BE/ME ratios. The results here explain why five-year

---

of these regressions span the optimal slopes. Because $\widehat{bm}_t$, in contrast, loads in exactly the same way on the changes in the market value of equity as bm_t, bm_t is not useful even for this purpose in Table 4's last regression.

reversal (De Bondt and Thaler 1985) is a good proxy for value,[11] yet the return spread on BE/ME-sorted portfolios is higher than the spread based on five-year reversals. The difference is that a sort on BE/ME ratios gets an additional return boost from net issuances.

An important result in Panel B is that the explanatory power of the three-factor model is significantly higher for the current BE/ME ratio-sorted hedge portfolio (73.8%) than for alternative value factors such as $\widehat{bm}_t$ (42.1%). Moreover, the explanatory power for the bm_{t-5}-sorted portfolio is 36.2%. The implication here, together with the negative alpha on the high-minus-low hedge portfolio for bm_t, is that HML captures some systematic variation in returns that is not priced, or that at least has a lower price than the rest of HML.

5 Implications and applications

5.1 Cheating the three-factor model

Table 5 tests whether there is unpriced, but systematic, variation in HML. We begin by creating a variable that picks up only the variation in bm_t that is not already part of $\widehat{bm}_t$. This variable, defined as the difference $bm_t^e = bm_t - \widehat{bm}_t$, does not covary with the past five-year change in the market value of equity. For example, $\text{cov}(bm_t^e, dme_t) = \text{cov}(bm_t, dme_t) - \text{cov}(\widehat{bm}_t, dme_t)$, but this difference is zero by the definition of how we constructed $\widehat{bm}_t$. We sort stocks into deciles based on this residual component of bm_t and estimate both CAPM and three-factor model regressions for returns on these deciles as well as the return earned by the high-minus-low portfolio.[12]

---

[11]See, for example, Fama and French (1996) and Asness, Moskowitz, and Pedersen (2012).

[12]One significant difference between the priced component, $\widehat{bm}_t$, and the residual component, $bm_t - \widehat{bm}_t$, is that the former does not vary significantly by industry. The R^2 from regressing $\widehat{bm}_t$ on indicator variables for the 49 Fama and French industries (with year fixed effects) is 1.23%. By contrast, the R^2 from regressing the residual component on these same indicator variables is 11.96%. Thus, industries explain a significant amount of the variation in BE/ME ratios that does not appear to be priced but that explains comovement in returns.

Excess returns earned by different deciles of bm_t^e are similar. The return spread between the highest and the lowest decile is only -7 basis points per month and insignificant with a t-value of -0.39. This residual component also does not covary significantly with the market. The market betas on different deciles range from 0.89 to 1.04, but the beta on the high-minus-low strategy is only 0.13 (t-value $=$ 3.64). The CAPM alpha is thus -12 basis points per month with a t-value of -0.70.

The three-factor alphas differ markedly across deciles. The loadings on the SMB and HML factors increase almost monotonically in bm_t^e, and the high-minus-low portfolio has SMB and HML loadings of 0.30 (t-value $=$ 6.41) and 0.74 (t-value $=$ 15.22). The three-factor model alpha on a strategy that purchases "value" stocks and sells "growth" stocks, as classified according to bm_t^e, is -49 basis points. This estimate has a t-value of -3.48. The high R^2 in the three-factor model regression for the high-minus-low portfolio (33.6%) implies that the unpriced component is an economically important part of the overall variation in HML.

The results in Table 5 suggest that a value factor constructed from the BE/ME ratios is a problematic variable for risk adjustment. The problem is that bm_t has two types of systematic variation, but only the systematic variation related to the changes in the market value of equity is priced. If a strategy, such as the one highlighted here, correlates only with the unpriced part, HML still assigns the $\hat{\lambda}_{\mathrm{hml}}$ price of risk to this strategy, thereby giving it a low risk-adjusted alpha. Hence, the three-factor model alpha is not a fair representation of a strategy's risk-adjusted performance. Here the *seemingly* profitable strategy is to purchase false growth stocks and short false value stocks, where false growth and value stocks are those defined using the variation in bm_ts not related to five-year change in the market value of equity.

A manager could also earn the value premium while hiding the true source of these profits from the three-factor model by buying the true value strategy (based on $\widehat{bm_t}$) and shorting the false value

strategy (based on bm_t^e). This strategy has an excess return of $61.2 - (-0.07) = 67.8$ basis points per month (t-value = 2.74), a three-factor model alpha of 63.8 basis points (t-value = 2.64), and a four-factor model alpha of 59 basis points (t-value = 2.39). Multi-factor models fail to explain this strategy's profits because its three-factor model loading on the HML factor is just 0.29 (t-value = 3.47). This true-minus-false value strategy is notable because it is more profitable than the traditional BE/ME strategy even in terms of excess returns (see the first row in Table 4 Panel B).

5.2   Using 25 Fama-French portfolios as test assets

The 25 Fama-French portfolios are used extensively to test asset pricing models.[13] If the unpriced component varies across BE/ME-sorted portfolios, then the 25 Fama-French portfolios can generate misleading inferences when used as test assets. Table 6 measures how prevalent the priced and unpriced components are within each of the 25 portfolios by regressing their returns on the market factor, SMB factor, and the priced and unpriced components of the HML factor. We construct the (priced) $\widehat{bm}_t$-based HML factor in the same way as the standard HML factor, except we replace bm_t with the priced component, $\widehat{bm}_t$. We construct the (unpriced) bm_t^e-based HML factor by replacing bm_t with the unpriced component, bm_t^e.

The estimates in Table 6 show that although the loadings on the priced component increase steadily as we move from low BE/ME portfolios to high BE/ME portfolios, so do the loadings on the unpriced component. For example, in the highest-size quintile, the loading on the priced component increases from BE/ME quintile 1's -0.12 (t-value $= -6.0$) to quintile 5's 0.44 (t-value = 10.05), and the loading on the unpriced component increases monotonically from -0.50 (t-value $= -20.5$) to 0.68 (t-value = 13.10).

---

[13]See Lewellen, Nagel, and Shanken (2010, p. 175) for a list of studies.

The unpriced-component loadings are problematic. Table 6 implies that an asset pricing model can price the 25 Fama-French portfolios not only because it covaries correctly with these portfolios' priced components, but also because it covaries with their unpriced components. A model's ability to price these portfolios thus does not show conclusively that the model explains the value premium. Hence, BE/ME-sorted portfolios, such as the 25 Fama-French portfolios, can be ill-suited for testing asset pricing models for a reason distinct from that detailed in Lewellen, Nagel, and Shanken (2010). A better set of test assets would be those sorted only by the priced component of BE/ME ratios or by the five-year change in the market value of equity.

5.3   Measuring skill among mutual fund managers

We next evaluate whether the unpriced component of HML affects inferences about money manager skill. Studies extensively use multi-factor models to search for skill.[14] If managers' strategies covary with the unpriced component of the HML factor, then estimated alphas are biased.

To evaluate money manager skill, we start with Morningstar's mutual fund database. We keep only U.S. equity funds and exclude index funds to isolate active managers. To be consistent with prior research, we start our sample in 1984 and require a minimum of 36 months of returns. Funds only enter our sample when their assets under management reach a minimum of $5 million in December 2000 dollars. We use this filter to remove bias that can arise from fund incubation documented by Evans (2010). These filters lead to final sample of 3,694 U.S. equity mutual funds.

To estimate alphas, we use both the traditional three-factor model and a multi-factor model that includes separate factors based on the priced and unpriced components of HML. This alternative multi-factor model allows for different prices of risk for the two systematic components of HML. Specifically,

---

[14]See, for example, Fama and French (2010), Kosowski, Timmermann, Wermers, and White (2006), and Linnainmaa (2012).

for the alternative model we estimate the following regressions

$$r_{j,t} - r_{f,t} = \alpha_j + b_j(R_{m,t} - r_{f,t}) + s_j SMB_t + h_j HML_t + e_{j,t} \tag{6}$$

$$r_{j,t} - r_{f,t} = \alpha_j + b_j(R_{m,t} - r_{f,t}) + s_j SMB_t + p_j HMLP_t + u_j HMLU_t + e_{j,t} \tag{7}$$

where $r_{j,t}$ is fund j's month-t net of fees return, $r_{f,t}$ is the risk-free rate, $R_{m,t}$ is the month-t return on the value weighted CRSP index, SMB_t is the month-t return on a long-short size portfolio, HML_t is the month t return on a long-short book-to-market portfolio, and $HMLP_t$ and $HMLU_t$ are month-t returns on long-short portfolios of the priced and unpriced components of HML.

To compare the estimates of skill under the two models, we evaluate the distributions of the t-statistics of the estimated alphas. We evaluate the t-statistics instead of the estimated alphas to control for differences in the standard errors of the estimates. As shown in Figure 1 the distribution of t-statistics for alphas estimated with the multi-factor model that splits HML into priced and unpriced components is shifted to the right compared to the distribution based on the traditional three-factor model. This rightward shift implies that the traditional three-factor model that uses HML underestimates money manager skill in the economy.

The differences in the distributions are significant at both the mean and the median for the estimated alphas and t-statistics. For the traditional three-factor model, the means (medians) are $\alpha = -1.318$ and $t(\alpha) = -0.919$ ($\alpha = -0.992$ and $t(\alpha) = -0.586$). When we allow for separate prices of risk for the two systematic HML components, the means (medians) rise to $\alpha = -0.642$ and $t(\alpha) = -0.353$ ($\alpha = -0.624$ and $t(\alpha) = -0.354$). At the mean and the median, these changes are statistically significant at the 0.001 level. Moreover, alphas increase for over 67% of the funds.

Overall, we find that the unpriced component of HML distorts estimates of money manager skill

based on the traditional three-factor model. The mean loading on HML in the traditional three-factor model is 0.021. When we split HML into the two factors, we find that the loadings go in the opposite directions. They negatively covary with the priced component: the median loading on $HMLP_t$ is -0.029. And managers' strategies positively covary with the unpriced component: the median loading on $HMLU_t$ is 0.076 with over 62% of the loadings greater than zero. This positive covariance with the unpriced component of HML causes the three-factor model to underestimate money manager skill.

5.4   Resurrecting anomalies

Any price-scaled variable can be decomposed in the same way as the BE/ME ratio. Daniel and Titman (2006), for example, decompose the log-earnings-to-price ratio as

$$ep_t = ep_{t-k} + \sum_{\tau=t-k+1}^{t} de_\tau - \sum_{\tau=t-k+1}^{t} dme_\tau \tag{8}$$

where de_τ is the log-change in earnings in year τ. This decomposition requires that earnings are positive over the span of the decomposition.

Table 7 Panel A shows five-year variance decompositions for BE/ME, E/P, C/P, and D/P ratios. The E/P, C/P, and D/P decompositions indicate that changes in the market value of equity play a smaller role than for the BE/ME ratio. Both E/P and C/P, for example, reflect to a significant extent the most recent changes in earnings and cash flows. The reason is that whereas the book value of equity keeps track of the accumulation of earnings and therefore usually changes slowly, both E/P and C/P can change abruptly in response to earnings or cash-flow shocks.

Panel B uses the same method as Table 3 to evaluate the extent to which the different components of price-scaled variables explain cross-sectional variation in returns in the absence of other conditioning

information. The baseline specification regresses returns against log-size and the current price-scaled variable. The bottom rows estimate univariate regressions of the first-stage residuals against each component. The first column reports each component's optimal slope $(\hat{b}_j)$ when all other components' slopes are held at the common slope b^*. The second column is the deviation between the optimal and common slope. The E/P and C/P regressions show that both variables are significant in explaining variation in returns. The D/P ratio enters with a positive sign but is insignificant. The "fundamentals" of E/P and C/P behave differently than those of BE/ME in the analysis of deviations. The optimal slopes on the changes in earnings and cash flows are always significant and also statistically closer to the common-slope benchmark than to zero. This analysis suggests that E/P and C/P ratios derive some of their predictive power from the fundamentals.

Table 7 Panel C shows that the E/P, C/P, and D/P ratios covary not only with the priced part of BE/ME ratios, but also with the unpriced part. E/P and C/P, for example, covary significantly more with the unpriced part than with the priced part. The fact that these price-scaled variables covary positively with the unpriced part of BE/ME ratio is not surprising. The decompositions in Panel A suggest that if we were to extend the decompositions backwards in time, we would find significant loadings on the changes in the market value of equity in years $t-5$, $t-6$, and so forth. This pattern is common across all price-scaled variables, including BE/ME, and this commonality shows up in Panel C as positive covariances between E/P, C/P, D/P, and the unpriced part of bm_t.

Table 8 evaluates the extent to which the three-factor model prices the E/P, C/P, and D/P anomalies through the unpriced-risk channel. The first three columns show average excess returns, CAPM alphas, and three-factor model alphas for portfolios sorted using each ratio. The three-factor model well describes the returns on the extreme portfolios sorted by these variables, and so the high-minus-low portfolio earns excess returns close to zero: 7 basis points per month for E/P-, 6 basis points per

month for C/P-, and -15 basis points per month for D/P-sorted high-minus-low portfolios. None of these high-minus-low portfolio returns are close to being statistically significant.

The second set of columns sorts stocks into portfolios based on the information in E/P, C/P, and D/P ratios, but does so based on the variation accrued into these variables over the past five years. At the end of each June, when stocks are assigned to portfolios, we first estimate the covariances between each price-scaled variable and the five changes in fundamentals (such as $de_t, \ldots, de_{t-4}$) and the five changes in the market value of equity. We then use these covariances to create a linear combination of the 10 variables (the five changes in fundamentals and the five changes in market values) such that the combination's covariances with the 10 variables are the same as they are for the original price-scaled variable. For the earnings-to-price ratio, for example, we solve for weights such that $\text{cov}(ep_t, df_{t-k}) = \text{cov}(\widehat{ep}_t, df_{t-k})$ and $\text{cov}(ep_t, dme_{t-k}) = \text{cov}(\widehat{ep}_t, dme_{t-k})$ for $0 \le k \le 4$. These variables, which we call the adjusted price-scaled ratios, capture the information accrued into the original ratios over the past five years but are free of the "old" information that covaries significantly with the unpriced part of HML.[15]

The results for the portfolios sorted by adjusted variables are striking. The three-factor model no longer prices portfolios sorted by E/P and C/P. Consider the portfolios sorted by E/P. Whereas the three-factor model alphas for the lowest and highest unadjusted-E/P deciles were -1 basis point (t-value $= -0.13$) and 6 basis points (t-value $= 0.58$) per month, the alphas on these two portfolios diverge significantly when stocks are sorted into portfolios based on the adjusted E/P. The three-factor model alphas for the low and high portfolios are now -22 basis points (t-value $= -2.29$) and 28 basis

---

[15]This computation results in a variable that differs from the one that would be obtained by subtracting a five-year-old price-scaled variable from the current price-scaled variable. The results in Table 7 show that such differencing would create an oddly behaving variable. In the case of E/P and C/P, for example, such differencing would create a variable that loads significantly on both year t and $t-5$ changes in fundamentals. The covariance-matching method, by contrast, properly purges variation that is associated with the "old" price-scaled variable.

points (t-value $= 2.70$) per month. Hence a high-minus-low portfolio based on the adjusted E/P ratio earns a per-month three-factor model alpha of 50 basis points with a t-value of 3.06. The return on the high-minus-low portfolio based on C/P also increases significantly, from 3 basis points (t-value $= 0.25$) to 32 basis points (t-value $= 2.06$) after adjustment.

Although the return on the D/P strategy is affected as well, the return on the high-minus-low portfolio goes from being moderately negative (-15 basis points) to moderately positive (18 basis points) but remains statistically insignificant because, as Table 7 shows, D/P behaves very similarly to BE/ME in terms of not only the sources of variation (mostly the changes in the market value of equity) but also the source of profits (these same changes). The three-factor model prices the D/P-sorted portfolios because their returns arise from the same source as those of the priced part of the HML factor.

Table 8's results are important because they reflect returns on two sets of strategies that are equally implementable. We do not use any forward-looking information to construct the adjusted versions of E/P, C/P, and D/P. The key observation is that the three-factor model cannot explain the returns of "pure" E/P and C/P anomalies. The three-factor model appears to explain these anomalies because BE/ME, E/P, and C/P all reflect some variation in the very old changes in the market value of equity. Because of this shared variation, the HML factor covaries positively with the high unadjusted-E/P portfolio, and it takes credit for this portfolio's high return. However, when E/P (or C/P) is adjusted so that it is only about recent changes in earnings (or cashflows) and market values of equity, the three-factor model fails.

6   Evidence from Japan

Daniel and Titman (2001) use Japanese data to pit the Fama and French (1993) three-factor model against the characteristic model. They use Japanese data to re-evaluate the question from Daniel and Titman (1997) because Japanese stock returns are even more closely related to their book-to-market ratios than their U.S. counterparts, and because the Japanese market has the largest number of listed stocks after the U.S.—over 2,000 as of 2010.

We also use Japanese data to test the role that the changes in market value of equity play in regressions of returns on book-to-market ratios. We use monthly Datastream return data on all Japanese companies listed on the Tokyo Stock Exchange in the 1981–2010. We use the same conventions as Daniel and Titman (2001) to time the variables. Book-to-market ratios and market values of equity are updated once a year in September. Book values of equity are from the fiscal year that ends on or before March of year t and market value of equity in September.

Table 9 estimates the same regressions as those in Table 1 using Japanese stocks. The average regression coefficient on the book-to-market ratio is significant in the all stocks, big stocks, and small stocks samples, and this significance is comparable to that found in the U.S. sample. Regressions (2), (5), and (8) show that when the regressions also control for the five-year change in the market value of equity, BE/ME ratios lose their significance. Among small stocks, BE/ME ratio is marginally significant when the regression only controls for the cumulative change in the market value of equity— but only because BE/ME correlates unevenly with past changes in the market value of equity. Similar to the U.S. sample, a disaggregation of the changes in the market value of equity (regressions (3), (6), and (9)) pushes the slopes on the BE/ME ratios even closer to zero. The increases on the average adjusted R^2s are also non-trivial. In the full sample this average increases from 3.7% to 7.3% when

moving from the baseline regression (1) to the regression with year-by-year changes in the market value of equity. Much like in the U.S., BE/ME ratios predict returns on Japanese stocks because they correlate with changes in the market values of equity.

7 Conclusions

All empirical risk factors are subject to the problem of confounded factors. If an empirical factor combines multiple factors with different risk premia, we misestimate assets' risk premia. We examine the implications of this problem in the context of the HML factor, but the problem is more general. It applies not just to the HML factor, but to any empirical risk factor, such as size, momentum, or liquidity—we do not know whether these factors have one-to-one correspondence with a primitive risk factor or whether they confound multiple risk factors.

In four empirical applications, we demonstrate how the confounded risk factor problem in the HML distorts inferences. First, we show that trading strategies can circumvent the three-factor model. Second, the unpriced component presents a problem in using BE/ME-sorted portfolios (such as the 25 Fama-French portfolios) to test asset pricing models. Third, the unpriced component leads to downwardly biased estimates of money manager skill based on the traditional three-factor model. Fourth, we show that the three-factor model prices the E/P and C/P anomalies in part because these anomalies covary with the unpriced part of the HML.

In addition to demonstrating the problem of confounded factors, our results provide a clear testable prediction for theories of value: changes in the market value of equity should subsume the value premium. If "value" within a model is something that exists independent of these changes, such a model is inconsistent with the data. Moreover, to match the data, a theory that models the behavior

of BE/ME ratios should produce ratios that decompose into priced and unpriced components. Models that generate the value premium by assuming that tangible and intangible assets have different risk exposures appear to hold the most promise.

REFERENCES

Asness, C. and A. Frazzini (2011). The devil in HML's details. AQR Capital Management working paper.

Asness, C., T. J. Moskowitz, and L. H. Pedersen (2012). Value and momentum everywhere. *Journal of Finance*. Forthcoming.

Berk, J. B., R. C. Green, and V. Naik (1999). Optimal investment, growth options, and security returns. *Journal of Finance 54*(5), 1553–1607.

Carlson, M., A. Fisher, and R. Giammarino (2004). Corporate investment and asset price dynamics: Implications for the cross-section of returns. *Journal of Finance 59*(6), 2577–2603.

Cochrane, J. H. (1992). Explaining the variance of price-dividend ratios. *Review of Financial Studies 5*(2), 243–280.

Cochrane, J. H. (2005). *Asset Pricing, Revised Edition.* Princeton, New Jersey: Princeton University Press.

Daniel, K. and S. Titman (1997). Evidence on the characteristics of cross sectional variation in stock returns. *Journal of Finance 52*(1), 1–33.

Daniel, K. and S. Titman (2001). Explaining the cross-section of stock returns in Japan: Factors or characteristics? *Journal of Finance 56*(2), 743–766.

Daniel, K. and S. Titman (2006). Market reactions to tangible and intangible information. *Journal of Finance 61*(4), 1605–1643.

Davis, J. L., E. F. Fama, and K. French (2000). Characteristics, covariances, and average returns: 1929 to 1997. *Journal of Finance 55*(1), 389–406.

De Bondt, W. F. M. and R. H. Thaler (1985). Does the stock market overreact? *Journal of Finance 40*(3), 379–395.

Evans, R. B. (2010). Mutual fund incubation. *Journal of Finance 65*(4), 1581–1611.

Fama, E. F. and K. French (1992). The cross-section of expected stock returns. *Journal of Finance 47*(2), 427–465.

Fama, E. F. and K. French (1993). Common risk factors in the returns of stocks and bonds. *Journal of Financial Economics 33*(1), 3–56.

Fama, E. F. and K. French (1995). Size and book-to-market factors in earnings and returns. *Journal of Finance 50*(1), 131–155.

Fama, E. F. and K. French (1996). Multifactor explanations of asset pricing anomalies. *Journal of Finance 51*(1), 55–84.

Fama, E. F. and K. French (2008). Average returns, B/M, and share issues. *Journal of Finance 63*(6), 2971–2995.

Fama, E. F. and K. French (2010). Luck versus skill in the cross section of mutual fund returns. *Journal of Finance 65*(5), 1915–1947.

Gomes, J., L. Kogan, and L. Zhang (2003). Equilibrium cross section of returns. *Journal of Political Economy 111*(4), 693–732.

Kosowski, R., A. Timmermann, R. Wermers, and H. White (2006). Can mutual fund "stars" really pick stocks? New evidence from a bootstrap analysis. *Journal of Finance 61*(6), 2551–2595.

Lewellen, J., S. Nagel, and J. Shanken (2010). A skeptical appraisal of asset pricing tests. *Journal of Financial Economics 96*(2), 175–194.

Linnainmaa, J. (2012). Reverse survivorship bias. *Journal of Finance*. Forthcoming.

Novy-Marx, R. (2012). The other side of value: The gross profitability premium. *Journal of Financial Economics*. Forthcoming.

Ross, S. A. (1976). The arbitrage theory of capital asset pricing. *Journal of Economic Theory 13*(3), 341–360.

Table 1: Average returns, BE/ME ratio, and changes in the market value of equity

This table shows average Fama-MacBeth regression slopes and their t-values from cross-section regressions to predict monthly returns for all individual stocks, All but Microcaps stocks (above the 20th percentile of market capitalization for NYSE stocks), and Microcaps (below the 20th NYSE market capitalization percentile). The regressions, estimated monthly, use variables that are updated six months after the firm's fiscal year end. The regressions are estimated using data from July 1968 through December 2011. In these regressions, me_t is the natural log of market capitalization, bm_t is the natural logarithm of the BE/ME ratio, and dme_t is the change in the market value of equity over fiscal year t, $\ln(\text{ME}_t/\text{ME}_{t-1})$.

Explanatory	All stocks			All but Microcaps			Microcaps		
variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
me_t	-0.074	-0.053	-0.060	-0.050	-0.058	-0.063	-0.261	-0.183	-0.182
	(-1.74)	(-1.33)	(-1.57)	(-1.29)	(-1.50)	(-1.72)	(-3.53)	(-2.66)	(-2.71)
bm_t	0.261	0.104	0.051	0.172	0.059	0.005	0.323	0.165	0.113
	(3.51)	(1.40)	(0.71)	(2.08)	(0.80)	(0.07)	(4.02)	(1.84)	(1.22)
$\sum_{\tau=t-4}^{t} dme_\tau$		-0.272			-0.173			-0.311	
		(-5.17)			(-3.20)			(-4.81)	
dme_t			-0.339			-0.084			-0.429
			(-2.91)			(-0.69)			(-3.23)
dme_{t-1}			-0.299			-0.202			-0.369
			(-3.37)			(-2.09)			(-3.40)
dme_{t-2}			-0.280			-0.162			-0.313
			(-3.78)			(-1.80)			(-3.55)
dme_{t-3}			-0.214			-0.126			-0.262
			(-3.38)			(-1.60)			(-3.63)
dme_{t-4}			-0.207			-0.213			-0.217
			(-3.55)			(-2.71)			(-3.11)
Average R^2	2.15%	2.56%	3.42%	2.38%	3.09%	4.96%	1.28%	1.62%	2.25%

Table 2: Cross-sectional decompositions of BE/ME ratios

This table decomposes the BE/ME ratio in year t into the BE/ME ratio in year $t - k$ plus the changes in the book (dbe_τ) and market ($-dme_\tau$) values of equity from year $t - k$ to t. The first row presents a one-year decomposition, the second row a two-year decomposition, and so forth. Each estimate is the percentage of variation in today's BE/ME ratios explained by the component indicated by the column. Columns dbe_τ and $-dme_\tau$ indicate the amount of variation explained by the change in the book or market value of equity τ years ago; columns $\sum dbe_\tau$ and $-\sum dme_\tau$ indicate the amount explained by the cumulative change in the book or market value of equity from year $t - k + 1$ to t. As the lag of the decomposition grows, data from fewer firms are available. Hence, the amount of variation due to the cumulative changes in the book and market values of equity (shown in the last two columns) does not exactly match the sums of the one-year estimates (shown in the "one-year changes" columns). A firm is used for estimating year $t - k$ covariances if the necessary information for it is available in years $t - k$ and t. The sample is from 1963 through 2011 for horizon $= 1$, from 1964 through 2011 for horizon $= 2$, and so forth. The covariances used to measure these percentages (see equation (3)) are estimated with year fixed effects. Standard errors, reported in square brackets, are block bootstrapped by year.

| Horizon | bm_{t-k} | One-year changes | | | Cumulative changes | |
		dbe_τ	$-dme_\tau$		$\sum dbe_\tau$	$-\sum dme_\tau$
1	81.84	−4.93	23.09		−4.93	23.09
	[1.41]	[0.60]	[1.54]		[0.60]	[1.54]
2	70.36	−4.23	14.80		−9.09	38.72
	[1.89]	[0.46]	[1.11]		[0.77]	[1.98]
3	62.61	−2.74	9.80		−11.49	48.88
	[1.92]	[0.41]	[1.09]		[0.93]	[2.08]
4	56.74	−2.27	7.58		−13.43	56.69
	[1.96]	[0.34]	[0.77]		[1.07]	[2.06]
5	51.73	−1.90	6.60		−14.85	63.13
	[1.48]	[0.31]	[0.91]		[1.15]	[1.68]
6	48.61	−1.38	4.38		−15.81	67.20
	[1.41]	[0.30]	[0.86]		[1.35]	[1.78]
7	46.33	−1.45	3.89		−16.69	70.36
	[1.33]	[0.34]	[0.86]		[1.53]	[1.92]
8	44.51	−1.52	3.37		−17.31	72.81
	[1.20]	[0.28]	[0.85]		[1.41]	[1.76]
9	42.70	−1.50	3.17		−17.89	75.19
	[1.19]	[0.24]	[0.74]		[1.44]	[1.96]
10	41.74	−1.49	2.59		−18.58	76.85
	[1.45]	[0.23]	[0.68]		[1.43]	[2.14]

Table 3: Fama-MacBeth regressions with BE/ME ratio components

This table reports Fama-MacBeth regressions that assess which components of the BE/ME ratio contribute to its ability to explain cross-sectional variation in monthly returns. Panel A presents the baseline regression. Panels B, C, and D report estimates from univariate regressions of first-stage residuals against each component. These regressions measure how close each component's optimal slope $(\hat{b}_j)$ is to the common slope (b^*) when all other components' slopes remain fixed at b^*. The first t-value tests whether each component's optimal slope is different from zero. The second t-value tests whether the optimal component slope is significantly different from the common slope reported in Panel A. The regressions presented in Panel B decompose the BE/ME ratio into the BE/ME ratio five years ago and the cumulative log-changes in the book and market values of equity. Panel C disaggregates the log-changes by year. Panel D presents a low-level decomposition in which the annual changes in the book and market values of equity are further disaggregated into total return, dividend yield, and net issuance components.

Panel A: Baseline regression

Regressors	EST	t-value
me_t	-0.074	-1.86
bm_t	0.261	3.63
R^2	2.15%	

Panel B: Five-year changes in the book and market value of equity

		t-values	
Regressors	EST	$H_0 : \hat{b}_j = 0$	$H_0 : \hat{b}_j = b^*$
bm_{t-k}	0.133	1.40	-2.80
$\sum_{\tau=t-k+1}^{t} dbe_\tau$	0.087	1.32	-4.04
$-\sum_{\tau=t-k+1}^{t} dme_\tau$	0.439	4.98	4.98

Panel C: Disaggregated five-year changes in the book and market value of equity

Regressors	EST	t-values $H_0 : \hat{b}_j = 0$	$H_0 : \hat{b}_j = b^*$
bm_{t-k}	0.133	1.40	-2.80
dbe_t	-0.026	-0.24	-2.83
dbe_{t-1}	-0.054	-0.58	-3.94
dbe_{t-2}	0.131	1.40	-1.71
dbe_{t-3}	0.133	1.48	-1.67
dbe_{t-4}	0.161	1.71	-1.26
$-dme_t$	0.447	3.68	2.04
$-dme_{t-1}$	0.442	4.16	2.51
$-dme_{t-2}$	0.415	4.22	2.29
$-dme_{t-3}$	0.387	3.69	2.03
$-dme_{t-4}$	0.421	4.09	2.71

Panel D: Five-year low-level decomposition of BE/ME ratios

Regressors	EST	t-values $H_0 : \hat{b}_j = 0$	$H_0 : \hat{b}_j = b^*$	Regressors	EST	t-values $H_0 : \hat{b}_j = 0$	$H_0 : \hat{b}_j = b^*$
bm_{t-k}	0.133	1.40	-2.80				
Book side:				Market side:			
roe_t	0.247	1.60	-0.11	$-r_t$	0.348	2.79	0.86
roe_{t-1}	0.202	1.50	-0.63	$-r_{t-1}$	0.382	3.57	1.46
roe_{t-2}	0.213	1.66	-0.57	$-r_{t-2}$	0.387	3.97	1.66
roe_{t-3}	0.204	1.49	-0.56	$-r_{t-3}$	0.363	3.60	1.55
roe_{t-4}	0.203	1.66	-0.72	$-r_{t-4}$	0.412	4.23	2.52
$-r^b_{\text{div},t}$	-0.509	-0.52	-0.75	$r_{\text{div},t}$	-1.337	-0.75	-0.92
$-r^b_{\text{div},t-1}$	-0.377	-0.39	-0.63	$r_{\text{div},t-1}$	-0.193	-0.10	-0.23
$-r^b_{\text{div},t-2}$	-0.100	-0.10	-0.36	$r_{\text{div},t-2}$	-0.128	-0.06	-0.19
$-r^b_{\text{div},t-3}$	0.355	0.39	0.10	$r_{\text{div},t-3}$	-0.186	-0.10	-0.24
$-r^b_{\text{div},t-4}$	-0.179	-0.19	-0.45	$r_{\text{div},t-4}$	-0.755	-0.39	-0.54
$r^b_{\text{iss},t}$	-0.265	-2.20	-4.18	$-r_{\text{iss},t}$	1.617	7.54	7.65
$r^b_{\text{iss},t-1}$	-0.256	-2.09	-4.09	$-r_{\text{iss},t-1}$	1.402	6.23	6.11
$r^b_{\text{iss},t-2}$	0.083	0.63	-1.34	$-r_{\text{iss},t-2}$	0.949	4.01	3.31
$r^b_{\text{iss},t-3}$	-0.008	-0.06	-1.98	$-r_{\text{iss},t-3}$	0.816	3.52	2.84
$r^b_{\text{iss},t-4}$	0.110	1.05	-1.41	$-r_{\text{iss},t-4}$	0.664	2.91	2.07

Table 4: The priced component of BE/ME ratios

This table reports Fama-MacBeth regressions and portfolio sorts that compare the information content of the current BE/ME ratio (bm_t) to alternative BE/ME constructions. Panel A reports the average coefficients, the t-values associated with these averages, and average adjusted R^2s from monthly Fama-MacBeth regressions. All coefficients are multiplied by 100. Panel B sorts stocks into 10 portfolios based on NYSE breakpoints on the alternative BE/ME ratio construct at the end of each June and holds the portfolio for the next year. The first row of each block reports the average monthly excess returns over the one-month Treasury bill rate, the second row reports the CAPM alphas, and the third row reports the three-factor model alphas. The R^2 in the last column is the adjusted R^2 from a time-series regression of high-minus-low $(10-1)$ hedge portfolio returns against the market factor or the three-factor model factors. The five alternative BE/ME constructs are (1) the BE/ME ratio from five years earlier, (2) minus the sum of log-changes in the market value of equity over the past five years, (3) the sum of log-changes in the book value of equity over the past five years, (4) the negative of the five-year stock returns, and (5) a modified version of the current BE/ME ratio that only includes variation due to the five-year change in the market value of equity.

Panel A: Fama-MacBeth regressions

Alternative repackaging of bm_t	Regressors						
	me_t		bm_t		Alternative bm_t		
	EST	t-value	EST	t-value	EST	t-value	R^2
Baseline	-0.074	-1.74	0.261	3.51			2.15%
bm_{t-5}	-0.106	-2.55			-0.018	-0.28	1.97%
	-0.075	-1.80	0.344	4.51	-0.176	-2.86	2.51%
$-\sum dme_\tau$	-0.059	-1.60			0.311	5.70	2.11%
	-0.053	-1.33	0.104	1.40	0.272	5.17	2.56%
$\sum dbe_\tau$	-0.089	-2.25			-0.219	-4.36	1.82%
	-0.062	-1.46	0.244	3.33	-0.186	-3.95	2.40%
$-\sum r_\tau$	-0.060	-1.73			0.254	4.14	2.09%
	-0.051	-1.34	0.159	1.97	0.197	3.02	2.66%
$\widehat{bm}_t$	-0.055	-1.48			0.730	5.17	2.25%
	-0.053	-1.32	0.056	0.76	0.679	4.74	2.63%

Panel B: Portfolio sorts

Sorting variable	$\hat{\alpha}$	Deciles				10-1	10-1 R^2
		1	2	9	10		
Current	$\bar{r}^e$	0.283	0.452	0.705	0.780	0.496	
BE/ME:		(1.23)	(2.07)	(3.30)	(3.16)	(2.48)	
bm_t	CAPM	−0.144	0.035	0.334	0.381	0.524	0.31%
		(−1.74)	(0.58)	(3.19)	(2.63)	(2.62)	
	FF	0.088	0.103	0.021	−0.059	−0.147	73.79%
		(1.50)	(1.77)	(0.31)	(−0.64)	(−1.41)	
Old	$\bar{r}^e$	0.390	0.463	0.548	0.578	0.188	
BE/ME:		(1.81)	(2.13)	(2.54)	(2.47)	(1.05)	
bm_{t-5}	CAPM	0.001	0.049	0.165	0.180	0.179	−0.12%
		(0.01)	(0.76)	(1.67)	(1.48)	(0.99)	
	FF	0.181	0.083	−0.016	−0.059	−0.240	36.20%
		(2.36)	(1.28)	(−0.19)	(−0.56)	(−1.64)	
Five-year changes	$\bar{r}^e$	0.251	0.460	0.650	0.813	0.563	
in ME:		(0.91)	(1.99)	(2.95)	(3.21)	(2.85)	
$-\sum dme_\tau$	CAPM	−0.257	0.022	0.274	0.390	0.647	4.40%
		(−2.49)	(0.31)	(2.40)	(2.81)	(3.34)	
	FF	−0.054	0.093	0.000	0.182	0.235	33.07%
		(−0.60)	(1.35)	(0.00)	(1.43)	(1.43)	
Five-year changes	$\bar{r}^e$	0.703	0.688	0.331	0.387	−0.316	
in BE:		(3.03)	(3.34)	(1.39)	(1.44)	(−1.84)	
$\sum dbe_\tau$	CAPM	0.298	0.322	−0.124	−0.114	−0.411	7.65%
		(2.66)	(3.47)	(−1.83)	(−1.21)	(−2.48)	
	FF	0.110	0.180	−0.052	0.091	−0.018	41.44%
		(1.08)	(2.08)	(−0.79)	(1.18)	(−0.14)	
Five-year	$\bar{r}^e$	0.386	0.419	0.647	0.737	0.351	
stock returns:		(1.42)	(1.97)	(2.74)	(2.63)	(1.68)	
$-\sum r_\tau$	CAPM	−0.115	0.020	0.241	0.273	0.388	0.58%
		(−1.13)	(0.28)	(2.01)	(1.74)	(1.85)	
	FF	0.091	0.067	−0.023	0.060	−0.030	35.11%
		(1.03)	(0.93)	(−0.24)	(0.46)	(−0.18)	
Variation due to five-year	$\bar{r}^e$	0.244	0.383	0.760	0.856	0.612	
changes in ME:		(0.90)	(1.67)	(3.35)	(3.18)	(2.87)	
$\widehat{bm}_t$	CAPM	−0.259	−0.043	0.375	0.414	0.673	1.86%
		(−2.57)	(−0.52)	(3.16)	(2.72)	(3.17)	
	FF	−0.054	0.033	0.074	0.095	0.149	42.10%
		(−0.62)	(0.41)	(0.81)	(0.75)	(0.90)	

Table 5: Cheating the three-factor model

This table sorts stocks into 10 portfolios at the end of every June based on an alternative version of the BE/ME ratio. This alternative ratio, bm_t^e, picks up the variation in bm_t that is not part of $\widehat{bm}_t$. It is defined as the difference $bm_t^e = bm_t - \widehat{bm}_t$ and therefore does not covary with the past five-year change in the market value of equity, but its remaining covariances are unaffected. This variable is reconstructed at the end of every June by estimating the required covariances and linear combinations from the cross-section. This table shows the average monthly excess returns, CAPM betas (m) and alphas (a), and three-factor model loadings (m for the market factor; s for the size factor; and h for the value factor) and alphas (a) for value-weighted deciles based on this variable. The t-values are reported in parentheses.

Model	Low	2	3	4	5	6	7	8	9	High	High − Low $\hat{\alpha}$	R^2
Excess return	0.40	0.45	0.49	0.44	0.49	0.46	0.46	0.57	0.47	0.34	−0.07	
	(1.98)	(2.00)	(2.26)	(2.02)	(2.21)	(2.18)	(2.24)	(2.79)	(2.21)	(1.40)	(−0.39)	
CAPM, m	0.91	1.04	1.01	0.98	1.01	0.94	0.91	0.89	0.92	1.03	0.13	2.21%
	(51.00)	(71.25)	(67.15)	(56.90)	(56.91)	(53.51)	(50.75)	(45.87)	(41.93)	(41.94)	(3.58)	
a	0.03	0.02	0.08	0.04	0.08	0.07	0.09	0.21	0.10	−0.09	−0.12	
	(0.39)	(0.35)	(1.18)	(0.47)	(0.96)	(0.90)	(1.03)	(2.26)	(0.97)	(−0.73)	(−0.70)	
FF3 model, m	0.90	1.05	1.00	1.03	1.07	1.01	0.97	0.98	1.01	1.12	0.22	33.56%
	(54.57)	(68.54)	(61.71)	(58.41)	(60.45)	(59.56)	(55.71)	(58.56)	(62.36)	(47.86)	(6.73)	
s	−0.24	−0.13	−0.03	−0.05	−0.11	−0.07	−0.04	−0.03	0.14	0.05	0.30	
	(−10.13)	(−5.74)	(−1.27)	(−1.80)	(−4.15)	(−3.05)	(−1.45)	(−1.14)	(5.80)	(1.62)	(6.41)	
h	−0.28	−0.07	−0.05	0.20	0.21	0.28	0.29	0.43	0.60	0.46	0.74	
	(−11.12)	(−3.16)	(−2.20)	(7.60)	(7.94)	(10.76)	(10.77)	(16.96)	(24.46)	(13.00)	(15.22)	
a	0.18	0.07	0.11	−0.05	−0.01	−0.05	−0.04	0.01	−0.19	−0.31	−0.49	
	(2.53)	(1.02)	(1.55)	(−0.70)	(−0.15)	(−0.67)	(−0.58)	(0.08)	(−2.73)	(−2.99)	(−3.48)	

Portfolio decile

Table 6: BE/ME- and size-sorted portfolios and their loadings on priced and unpriced components of HML

This table measures how prevalent the priced and unpriced components of BE/ME ratios are in the 25 portfolios sorted by size and BE/ME ratios ("the 25 Fama-French portfolios"). We estimate factor loadings on priced and unpriced components by regressing each portfolio's monthly returns on the market factor, the SMB factor, and the priced and unpriced components of the HML factor. We construct the (priced) $\widehat{bm}_t$-based HML factor in the same way as the standard HML factor, except we replace bm_t with the priced component, $\widehat{bm}_t$. The (unpriced) bm_t^e-based HML factor is constructed by replacing bm_t with the unpriced component, bm_t^e.

	Priced long-short factor based on $\widehat{bm}_t$										
	Factor Loadings						t-values				
	BE/ME quintile						BE/ME quintile				
Size	Low	2	3	4	High		Low	2	3	4	High
Small	−0.26	0.06	0.22	0.32	0.52		−6.05	1.83	8.33	11.73	16.66
2	−0.31	0.03	0.19	0.31	0.56		−9.82	1.17	7.04	10.58	18.26
3	−0.32	0.06	0.15	0.31	0.50		−10.78	1.73	4.83	9.34	13.29
4	−0.26	0.03	0.21	0.29	0.50		−8.93	0.91	6.13	8.67	12.76
Big	−0.12	0.09	0.05	0.33	0.44		−6.03	3.24	1.77	11.09	10.05

	Unpriced long-short factor based on bm_t^e										
	Factor Loadings						t-values				
	BE/ME quintile						BE/ME quintile				
Size	Low	2	3	4	High		Low	2	3	4	High
Small	−0.13	0.00	0.16	0.28	0.43		−2.46	−0.10	5.15	8.55	11.64
2	−0.21	0.14	0.37	0.45	0.57		−5.58	4.19	11.42	12.86	15.45
3	−0.27	0.22	0.49	0.55	0.60		−7.51	5.82	12.94	13.69	13.29
4	−0.32	0.27	0.43	0.49	0.66		−9.10	6.98	10.30	12.52	14.31
Big	−0.50	0.11	0.41	0.50	0.68		−20.52	3.32	11.16	13.89	13.10

Table 7: Other price-scaled variables

This table decomposes BE/ME ratio, earnings-to-price ratio (E/P), cashflow-to-price ratio (C/P), and dividend-to-price ratio (D/P) into what these variables were five years prior and how the fundamentals and market values of equity have changed between the two points in time. Panel A shows variance decomposition estimates. The first row (X_{t-5}) shows the amount of variation in today's variable that is due to the variation in the old variable. The rows labeled df show how much the changes in fundamentals (book value of equity, earnings, cashflows, or dividends) contribute to this variance, and the rows labeled $-dme$ show the importance of changes in the market values of equity. The covariances used to measure these percentages are estimated with year fixed effects. Standard errors, reported in square brackets, are clustered by year. Only firms that have existed for five years and for which the required fundamentals are positive are included in the sample. Panel B reports average coefficients (t-values associated with these averages) from monthly Fama-MacBeth regressions. The baseline specification explains returns using size and the current price-scaled variable. The analysis of deviations reports univariate regressions of the first-stage residuals against each component. The first column reports each component's optimal slope ($\hat{b}_j$) when all other components' slopes remain fixed at the common slope (b^*) and the second column reports the deviation between the component's optimal slope and the common slope. Panel C reports the pairwise cross-sectional covariances between BE/ME ratio, its two components, and E/P, C/P, and D/P. All variables here are in logs. The two BE/ME components are the priced and unpriced component of BE/ME (see text for details). All variables are demeaned cross-sectionally before estimating the covariances from the panel.

Panel A: Variance decompositions of other price-scaled variables

	Price-scaled variable			
Component	BE/ME	E/P	C/P	D/P
X_{t-5}	52.22	28.85	34.05	68.44
	[1.63]	[1.40]	[1.30]	[1.75]
df_t	−4.05	46.66	39.79	1.94
	[0.58]	[1.89]	[1.70]	[0.76]
df_{t-1}	−2.52	9.22	5.71	0.09
	[0.53]	[0.99]	[0.93]	[0.70]
df_{t-2}	−1.35	2.10	0.32	−1.07
	[0.47]	[0.95]	[0.80]	[0.25]
df_{t-3}	−1.35	−1.61	−3.29	−1.24
	[0.43]	[0.74]	[0.69]	[0.30]
df_{t-4}	−1.26	−2.74	−2.79	−1.90
	[0.40]	[1.05]	[0.63]	[0.47]
$-dme_t$	22.27	5.24	7.85	12.05
	[1.51]	[1.52]	[1.54]	[0.99]
$-dme_{t-1}$	14.00	0.74	2.71	7.02
	[1.11]	[1.20]	[1.15]	[0.84]
$-dme_{t-2}$	9.21	2.63	4.59	5.26
	[1.02]	[1.01]	[0.95]	[0.78]
$-dme_{t-3}$	6.92	4.00	5.14	4.88
	[0.82]	[0.92]	[0.86]	[0.61]
$-dme_{t-4}$	5.90	4.91	5.92	4.52
	[0.94]	[0.85]	[0.83]	[0.64]

Panel B: Fama-MacBeth regressions with other price-scaled variables

		Price-scaled variable:							
		BE/ME		E/P		C/P		D/P	
	Regressors	b_j	$b_j - b^*$	b_j	$b_j - b^*$	b_j	$b_j - b^*$	b_j	$b_j - b^*$
Baseline regression	me	-0.073		-0.067		-0.073		-0.045	
		(-1.71)		(-2.05)		(-2.22)		(-1.46)	
	X_t	0.255		0.180		0.280		0.061	
		(3.41)		(2.94)		(4.57)		(0.93)	
	Adj. R^2	2.20%		2.11%		2.03%		2.69%	
Deviations	X_{t-5}	0.130	-0.125	0.098	-0.082	0.219	-0.060	0.031	-0.030
		(1.36)	(-2.75)	(1.18)	(-2.12)	(2.68)	(-1.42)	(0.43)	(-1.04)
	df_t	-0.008	-0.263	0.136	-0.045	0.188	-0.092	-0.009	-0.070
		(-0.07)	(-2.51)	(2.42)	(-0.93)	(3.15)	(-1.69)	(-0.09)	(-0.93)
	df_{t-1}	-0.069	-0.324	0.121	-0.059	0.193	-0.087	-0.102	-0.163
		(-0.72)	(-3.86)	(1.96)	(-1.63)	(2.97)	(-2.17)	(-1.16)	(-2.59)
	df_{t-2}	0.113	-0.142	0.152	-0.028	0.227	-0.052	-0.078	-0.139
		(1.17)	(-1.83)	(2.48)	(-0.92)	(3.62)	(-1.43)	(-0.94)	(-2.16)
	df_{t-3}	0.109	-0.146	0.121	-0.059	0.222	-0.058	-0.077	-0.138
		(1.17)	(-1.84)	(1.98)	(-1.57)	(3.42)	(-1.41)	(-0.90)	(-2.14)
	df_{t-4}	0.140	-0.115	0.170	-0.011	0.241	-0.039	0.009	-0.051
		(1.44)	(-1.38)	(2.84)	(-0.31)	(3.77)	(-0.94)	(0.12)	(-0.83)
	$-dme_t$	0.434	0.179	0.206	0.026	0.304	0.024	-0.060	-0.121
		(3.56)	(1.94)	(1.55)	(0.25)	(2.34)	(0.24)	(-0.46)	(-1.06)
	$-dme_{t-1}$	0.434	0.178	0.380	0.199	0.479	0.199	0.108	0.047
		(4.04)	(2.45)	(3.28)	(2.38)	(4.32)	(2.43)	(0.94)	(0.52)
	$-dme_{t-2}$	0.410	0.154	0.302	0.122	0.415	0.135	0.218	0.158
		(4.14)	(2.27)	(2.65)	(1.49)	(3.83)	(1.68)	(2.12)	(1.95)
	$-dme_{t-3}$	0.378	0.123	0.311	0.130	0.393	0.113	0.287	0.226
		(3.61)	(1.98)	(2.78)	(1.76)	(3.66)	(1.58)	(2.82)	(2.99)
	$-dme_{t-4}$	0.426	0.171	0.343	0.163	0.441	0.161	0.113	0.052
		(4.12)	(2.91)	(3.17)	(2.35)	(4.20)	(2.37)	(1.07)	(0.66)

Panel C: Covariances between bm_t, $\widehat{bm}_t$, bm_t^e, and other price-scaled variables

	bm_t	$\widehat{bm}_t$	bm_t^e	ep_t	cp_t	dp_t
bm_t	0.497	0.145	0.351	0.150	0.171	0.202
$\widehat{bm}_t$		0.152	-0.007	0.034	0.045	0.112
bm_t^e			0.359	0.115	0.126	0.090

Table 8: Resurrecting anomalies

This table estimates monthly average excess returns, CAPM alphas, and three-factor model alphas for portfolios sorted on E/P, C/P, and D/P ratios. The columns labeled "Sorts on unadjusted variables" use each ratio as such directly to compute the deciles. The columns labeled "Sorts on adjusted variables" replace original price-scaled variables with repackaged versions that are based on five-year changes in fundamentals (E, C, or D) and market values of equity. This repackaging is done every June at the same time the portfolios are formed by first computing (from the cross-section) the covariance of the price-scaled variable with the five fundamental and five market value variables. The adjusted version of a price-scaled variable is a linear combination of these 10 components such that its covariances with these components are the same as what they are for the original price-scaled variable. This table shows the average excess returns, CAPM alphas, and three-factor model alphas for deciles 1 and 10 and for a high-minus-low portfolio. The t-values are reported in parentheses.

Anomaly variable	Model	Sorts on unadjusted variables				Sorts on adjusted variables			
		Deciles		High − Low		Deciles		High − Low	
		1	10	$\hat{\alpha}$	R^2	1	10	$\hat{\alpha}$	R^2
E/P	$\bar{r}^e$	0.25	0.78	0.53		0.17	0.86	0.69	
		(0.98)	(3.27)	(2.76)		(0.70)	(3.76)	(3.95)	
	CAPM	−0.22	0.38	0.60	2.42%	−0.28	0.46	0.74	2.13%
		(−2.22)	(3.01)	(3.11)		(−2.85)	(4.12)	(4.28)	
	FF3	−0.01	0.06	0.07	52.30%	−0.22	0.28	0.50	13.70%
		(−0.13)	(0.58)	(0.49)		(−2.29)	(2.70)	(3.06)	
C/P	$\bar{r}^e$	0.29	0.76	0.47		0.23	0.77	0.54	
		(1.15)	(3.28)	(2.55)		(0.90)	(3.38)	(3.15)	
	CAPM	−0.18	0.36	0.54	3.16%	−0.24	0.37	0.61	3.86%
		(−1.95)	(3.03)	(2.95)		(−2.52)	(3.35)	(3.60)	
	FF3	0.04	0.07	0.03	51.85%	−0.14	0.18	0.32	22.16%
		(0.59)	(0.78)	(0.25)		(−1.52)	(1.76)	(2.06)	
D/P	$\bar{r}^e$	0.29	0.52	0.23		0.17	0.64	0.47	
		(1.14)	(2.48)	(0.96)		(0.66)	(2.78)	(2.41)	
	CAPM	−0.19	0.24	0.43	19.47%	−0.30	0.27	0.56	5.80%
		(−2.11)	(1.55)	(2.05)		(−2.72)	(1.99)	(2.97)	
	FF3	−0.04	−0.18	−0.15	62.94%	−0.21	−0.03	0.18	31.95%
		(−0.44)	(−1.66)	(−1.00)		(−1.94)	(−0.31)	(1.07)	

Table 9: Average returns, BE/ME ratio, and changes in the market value of equity: Japanese stocks

This table shows average Fama-MacBeth regression slopes and their t-values from cross-section regressions to predict monthly returns for all stocks, big stocks (above the median market capitalization), and small stocks (below the median market capitalization) listed on the Tokyo Stock Exchange. Book-to-market ratios and market values of equity are updated once a year in September. Book values of equity are from the fiscal year that ends on or before March of year t and market value of equity in September. We winsorize book-to-market ratios every year at the 1% level in both tails. The regressions are estimated using data from October 1981 through December 2010. The number of stocks in the regressions ranges from a low of 304 in February 1982 to a high of 2,078 in October 2010. In these regressions, me_t is the natural log of market capitalization, bm_t is the natural logarithm of the BE/ME ratio, and dme_t is the change in the market value of equity from September of year $t-1$ to September of year t, $\ln(\mathrm{ME}_t/\mathrm{ME}_{t-1})$.

Explanatory	All stocks			Big stocks			Small stocks		
variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
me_t	−0.135	−0.086	−0.080	0.010	0.020	0.011	−0.409	−0.310	−0.278
	(−1.93)	(−1.31)	(−1.27)	(0.13)	(0.26)	(0.16)	(−4.02)	(−3.23)	(−3.02)
bm_t	0.258	0.088	0.029	0.273	0.054	−0.002	0.274	0.157	0.099
	(3.66)	(1.54)	(0.51)	(3.23)	(0.82)	(−0.03)	(3.49)	(2.16)	(1.33)
$\sum_{\tau=t-4}^{t} dme_\tau$		−0.505			−0.472			−0.467	
		(−3.91)			(−3.31)			(−3.23)	
dme_t			−0.858			−0.734			−0.859
			(−3.14)			(−2.62)			(−2.80)
dme_{t-1}			−0.595			−0.527			−0.617
			(−2.88)			(−2.27)			(−2.75)
dme_{t-2}			−0.446			−0.333			−0.489
			(−2.94)			(−1.80)			(−2.82)
dme_{t-3}			−0.370			−0.463			−0.257
			(−2.63)			(−2.60)			(−1.58)
dme_{t-4}			−0.281			−0.280			−0.232
			(−2.03)			(−1.55)			(−1.62)
Average R^2	3.71%	4.74%	7.30%	3.31%	4.98%	9.24%	1.61%	2.47%	4.53%

Figure 1: **Cumulative density functions for $t(\alpha)$ estimated using a three-factor and a multi-factor model that includes separate factors for the priced and unpriced components of HML.** This figure plots the cumulative distribution functions of t-statistics for alphas estimated with the traditional three-factor model ("HML") and a multi-factor model that splits HML into priced and unpriced components ("HML Priced & Unpriced"). We estimate these models using monthly returns for U.S. equity mutual funds obtained from the Morningstar database. We start our sample in 1984 and require a minimum of 36 months. Fund enter our sample after their asset under management reach a minimum of $5 million in December 2000 dollars. The final sample consists of 3,694 funds.