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Do relative leverage and relative distress really explain size and book-to-market anomalies?

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Abstract

In a capital asset pricing model (CAPM) framework, Ferguson and Shockley [2003. Equilibrium ‘‘anomalies’’. Journal of Finance 58, 2549–2580] propose two factors constructed on relative leverage and relative distress, and show that the two factors subsume Fama and French's [1993. Common risk factors in the returns on stocks and bonds. Journal of Financial Economics 33, 3–56] factors constructed on size and book-to-market (BM) in explaining the cross-sectional average returns of the 25 size-BM portfolios. Based on tests on individual securities, we find that all factors fail to fully explain the common asset-pricing anomalies. In the spirit of Merton's [1973. An intertemporal capital asset pricing model. Econometrica 41, 867–887] intertemporal CAPM, we propose an augmented five-factor model, which incorporates Ferguson and Shockley's [2003. Equilibrium ‘‘anomalies’’. Journal of Finance 58, 2549–2580] factors into the Fama–French three-factor model. The empirical results show that a simple conditional version of the augmented model is able to explain most well-known asset-pricing anomalies.

JEL classification: G12; G14

Keywords: Anomalies; Asset pricing; Equilibrium anomalies; Relative distress; Relative leverage

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1. Introduction

The seminal work of Fama and French (1992), which identifies size and book-to-market (BM) ratio as two major determinants of stock returns, initiated a debate on the nature of size and BM anomalies. A central issue of this debate is whether the two asset-pricing anomalies are driven by rational or behavioral forces.

Rationality supporters argue that the higher average returns on small-size and high-BM firms are compensation for risk that is not properly captured by the standard single-factor Capital Asset Pricing Model (CAPM). Thus, the effort has been devoted to search for a good, if not the correct, asset-pricing model that accounts for both cross-sectional and time-series variations in stock returns. A focus of this quest is whether stock returns are better characterized by a single-factor model or a multi-factor model, in either a conditional or an unconditional setting. Of special concern is whether there exists an asset-pricing model that can explain the asset-pricing anomalies such as size, BM, and momentum. Among various competing models, Fama and French's (1993) three-factor model that augments the CAPM with two mimicking portfolios formed on size (i.e., SMB) and BM (i.e., HML) remains one of the most popular empirical asset-pricing models. Fama and French (1993, 1998) document that the model explains the cross-sectional variation in stock returns for both the U.S. and international markets.[1] However, the Fama–French three-factor model still fails to fully account for the cross-sectional regularities related to size and BM.

In contrast, behavioral supporters argue that the premium between high- and low-BM stocks is too large to be explained by rational pricing theories (see, for example, Lakonishok et al., 1994).[2] Daniel and Titman (1997) in turn propose a characteristics-based model, arguing that it is firms' characteristics rather than risks that drive the BM premium.

However, the rejection of the factor hypothesis does not necessarily imply that size and BM premiums are caused by irrational or behavioral forces. In a CAPM framework, Ferguson and Shockley (2003, hereafter FS) argue that size and BM anomalies are induced by the improper use of an equity-only market portfolio to estimate the equilibrium expected return. They argue that the true market portfolio should be composed of two sub-portfolios: the economy's debt claims and the economy's equity claims. The economy's debt claims, which have been neglected in empirical asset-pricing tests, are associated with firms' distress and leverage risks. Thus, the higher expected returns of stocks with higher loadings on SMB and HML factors are just compensation for their relatively higher risk.

To complement the unobserved debt claims in the market portfolio, FS (2003) construct two mimicking portfolios to capture the part of returns associated with relative leverage (based on the ratio of debt-to-equity, D/E) and the part of returns associated with relative distress (based on Altman's (1968) Z score). They show that the two proposed factors,

[1] To better explain the driving forces behind size and BM, Fama and French (1993, 1995, 1996) demonstrate that the returns on SMB and HML capture the distress factor, as documented in Chan and Chen (1991). Fama and French (1995) also suggest that lower-BM firms have persistently strong earnings, while higher-BM firms are associated with persistently low earnings.

[2] Lakonishok et al. (1994) and Haugen (1995) argue that the premium is due to investor overreaction. They suggest that investors who are in favor of low-BM (growth) stocks tend to overvalue the performance of low-BM firms and undervalue the performance of high-BM (value) firms. When the overreaction is corrected, the BM premium is induced because value stocks have high returns while growth stocks have low returns.

namely the D/E and Z factors, subsume the SMB and HML factors in explaining the cross-sectional variation in stock returns on the Fama–French 25 size-BM portfolios. Thus, the size and BM effects are "equilibrium anomalies" that are simply consistent with the single-factor CAPM.

Some interesting observations emerge from FS's (2003) findings. First, many researchers have been trying to explain the pricing anomalies with risk-based multifactor models, either in a conditional or an unconditional setting, but FS (2003) show that the static one-factor CAPM alone suffices to explain the pricing anomalies. Second, while the FS (2003) three-factor model (hereafter FS model), which augments the single-index market model with the D/E and Z factors, better explains the cross section of stock returns, the Fama–French three-factor model (hereafter FF model) remains a better model in explaining the time-series variation in the 25 size-BM portfolio returns. A natural question arises: Does the FS model really fully explain the size and BM anomalies? It seems that although the FS model provides a good account of the cross-sectional stock returns, it has yet failed to capture some time-series patterns related to size and BM. If this is indeed the case, then maybe a static CAPM does not suffice to fully account for the asset-pricing anomalies.

Hence, the objective of this paper is threefold. First, we examine whether relative distress and relative leverage fully explain the asset-pricing anomalies. Although FS (2003) suggest that variables that correlate strongly with leverage, such as size, BM, and earnings-to-price, should not be considered as anomalies, they do not particularly examine whether the size and BM anomalies are completely eliminated when returns are evaluated based on their model.[3]

Second, since in the size-BM portfolios there is yet time-series variation that is not captured by the FS model but captured by the FF model, we explore whether there exists a model that is more complete in explaining stock returns both in the cross section and in time series. In the spirit of Merton's (1973) intertemporal CAPM (ICAPM), we propose a five-factor model by augmenting the FS model with SMB and HML. Merton's ICAPM indicates that over time equilibrium expected returns are determined by a two-factor model, the first being the market portfolio and the second a hedging portfolio that reflects the changes in the investment opportunity set. The FS (2003) factors, in combination with the market index, appear to proxy the market portfolio well, while SMB and HML factors appear to reflect the changes in the opportunity set *over time*. We demonstrate that the augmented five-factor model succeeds in explaining most existing asset-pricing anomalies, including the size, BM, and momentum effects.

Third, to investigate the performance of the competing asset-pricing models, we advocate an empirical methodology proposed by Brennan, Chordia, and Subrahmanyam, et al. (1998, hereafter BCS) that is free of the errors-in-variables (EIV) problem and yet can be implemented on individual securities.[4] Like many models that claim to outperform the

[3]In a similar attempt, Petkova (2006) proposes an alternative macroeconomic-based factor model and claims that the new model subsumes the FF model in explaining the cross-sectional variation in stock returns. Neither FS (2003) nor Petkova (2006) examines whether their models fully explain the size and BM anomalies.

[4]The errors-in-variables (EIV) problem, which has been widely documented in the asset-pricing test literature, states that the time-series estimation of the market beta or the factor loading in the two-stage method may suffer from the measurement errors. As a result, the coefficients estimated in the second-stage cross-sectional regression will be biased towards zero. Following Fama and MacBeth (1973), most studies group firms into portfolios to minimize the EIV problem. The rationale for this procedure is that if the errors in the betas for different firms are

FF model, the FS model is also tested with the Fama–French 25 size-BM portfolios. Lewellen, Nagel, and Shanken (2006, hereafter LNS) argue that inference based on these portfolios is misleading because the size-BM portfolios exhibit a strong factor structure; more than 90% of the time-series variation of the portfolio returns and more than 75% of the cross-sectional variation in the average return are explained by the FF model. Since the size-BM portfolios contain very few idiosyncratic components, it is easy to construct a factor model that yields significant factor risk premium(s), as long as the new factor(s) are related to the Fama–French factors. LNS (2006) find that the performance of most models is disappointing when the tests are implemented on samples expanded beyond the size-BM portfolios. They therefore suggest that researchers use larger samples beyond the size-BM portfolios, e.g., the individual securities, to allow for more powerful inference.

Since the BCS approach is implemented using individual securities, it has several potential advantages. First, it is more statistically powerful than the analysis based on portfolio samples, as is argued by LNS (2006). Second, it avoids the portfolio formation process problem that is commonly encountered in the Fama–MacBeth cross-sectional regression.[5] Third, it allows for testing different versions of asset-pricing models; both conditional and unconditional pricing models are easily tested with the BCS approach. We will discuss this in more detail in Section 4.

The empirical results are briefly summarized in the following. Based on a sample of stock returns available from CRSP/Compustat for 1966–2004, we find that the explanatory power of some non-risk firm characteristics, including size, BM, and past returns, remains significant, regardless of the model specifications used to calculate the risk-adjusted returns. Specifically, the empirical results reveal that even after stock returns are adjusted by FS's (2003) CAPM-based model, or by Fama and French's (1993) three-factor model, the explanatory power of various firm characteristics is still significant. In particular, the BM premium remains strongly significant, suggesting that neither the FS model nor the FF model can fully capture the BM anomaly.

In contrast, we find that a simple conditional version of the augmented five-factor model is able to explain almost all asset-pricing anomalies, including the BM, which is not captured by either the FF model or the FS model alone. The augmented five-factor model also fully explains the momentum effect, suggesting that the momentum effect is not due to mispricing. The only anomaly that remains is that several firm characteristics still have abnormal explanatory ability for January stock returns, suggesting that the January seasonality remains a puzzle that needs further investigation. Overall, our model suggests that most asset-pricing anomalies are consistent with a conditional multifactor model.

The rest of the paper is organized as follows. Section 2 presents a brief review of the literature on asset-pricing models. Section 3 presents the augmented five-factor model, with a review of FS's (2003) model. Section 4 presents the empirical methodology, and Section 5 describes the data. Section 6 presents the empirical results, along with some robustness checks. The last section concludes the paper.

(*footnote continued*)
not perfectly correlated, the betas of the portfolios can provide more precise estimates of true beta than the betas for individual securities due to noise cancellation.

[5]The portfolio formation process may induce two distinct problems: the under-rejection bias argued by Roll (1977) and the data-snooping bias suggested by Lo and MacKinlay (1990a).

2. Asset-pricing models and asset-pricing anomalies

The search for a risk-based asset-pricing model that accounts for both cross-sectional and time-series variations in stock returns is one of the most important issues in finance. A focus of this question is whether stock returns are better characterized by a single-factor model or a multi-factor model, in either a conditional or an unconditional setting. Of special concern is whether there exists a risk-based asset-pricing model that can explain the asset-pricing anomalies related to size, BM, and momentum.

In this section, we provide a brief review of the literature on asset-pricing models and discuss how they relate to the pricing anomalies. As the literature on asset-pricing models is too extensive to be reviewed in a single article (see, for example, Cochrane, 2001), the discussion here is very selective; the purpose is only to locate the position of our paper in the literature.

2.1. *Single-factor asset-pricing models*

There are two major single-factor asset-pricing models: the capital asset-pricing model (CAPM) of Sharpe (1964), Lintner (1965), and Black (1972) and the consumption-based CAPM (CCAPM) of Breeden (1979). While theoretically the CCAPM is preferable to the CAPM because the former takes into account the dynamic nature of investors' decisions and links asset prices to the fundamentals, empirically the CCAPM often performs worse than the CAPM (see, for example, Breeden et al., 1989).

But can the CAPM explain the pricing anomalies? Fama and French (1992) document that the "post-ranking" beta provides no incremental explanatory power beyond size and BM, thus suggesting that the CAPM fails to explain the asset-pricing anomalies unconditionally. Kim (1995) shows that the market beta still carries significant risk premium once the EIV problem in the Fama–MacBeth regression is corrected. Nevertheless, the size premium remains significant. In a static setting, FS (2003) show that size and BM may have explanatory power when an asset's beta is measured with respect to a market portfolio proxy composed of only equities. The failure to include the economy's debt claims in the market portfolio gives rise to the explanatory power of firm characteristics, such as relative leverage and relative distress on stock returns over that of the stock-only market index. By augmenting the single-index market model with two mimicking factor portfolios formed on relative distress and leverage, FS (2003) show that their alternative three-factor model outperforms the FF model in explaining the cross-sectional variation of returns on the 25 size-BM portfolios. The question remains, however, whether the FS model can fully explain the size and BM effects.

Researchers also attempt to explain the asset-pricing anomalies with the conditional CAPM. Theoretically, the conditional CAPM would be better in explaining the pricing anomalies than the unconditional CAPM because time-varying betas and price of risk in combination would generate a larger return spread that is more consistent with the magnitude of the observed anomalies.

In a general equilibrium setting that takes into account firms' production, Gomes et al. (2003) show that size and BM are correlated with the market beta. Petkova and Zhang (2005) empirically confirm that BM is related to the conditional beta. Avramov and Chordia (2006) show that when the beta is allowed to vary as suggested by Gomes et al. (2003), the size and value effects are often explained, but the explanatory power of past returns remains significant. Lewellen and Nagel (2006), however, show that the conditional CAPM cannot fully explain the asset-pricing anomalies because the variation in conditional betas would have been implausibly large to generate premiums that match the asset-pricing anomalies associated with BM and momentum.

To briefly sum up, it seems that researchers still hold discrepant views concerning whether the CAPM can really explain the asset-pricing anomalies. It appears that although FS's (2003) model outperforms the FF (1993) model in explaining cross-sectional stock returns, the FS (2003) model may not fully explain the size and BM effects because the model is a static one. Lewellen and Nagel (2006) show that even the conditional CAPM cannot fully explain the asset-pricing anomalies. Indeed, as we shall show later, the FS (2003) model does not fully explain the pricing anomalies, either conditionally or unconditionally. Instead, we find that a conditional multifactor model that augments the FS (2003) model with the two Fama-French factors (i.e., SMB and HML) provides a very good account of most asset-pricing anomalies.

2.2. Multi-factor asset-pricing models

There are also two major theory-based multifactor asset-pricing models: the intertemporal CAPM (ICAPM) of Merton (1973) and the arbitrage pricing theory (APT) of Ross (1976). Although the two models have different theoretical foundations (one based on the equilibrium argument and the other based on the no-arbitrage argument), empirically researchers have mostly treated the two models as the same, as it seems difficult to distinguish the two empirically. Perhaps a distinctive difference is that the ICAPM yields specific "predictions" concerning the nature of the state variables, whereas in the APT the common factors are only ambiguously specified. More specifically, the ICAPM suggests that when there is stochastic variation in the investment opportunity set, there will be risk premia associated with innovations in the state variables that describe the investment opportunities.

Connor (1995) classifies the empirical asset-pricing models into three broad categories: macroeconomic, statistical, and fundamental. The macroeconomic model of Chen et al. (1986) seems to be consistent with both the spirit of APT and ICAPM, but Petkova's (2006) macroeconomic-based model appears to conform more to the ICAPM. Statistical factors such as Roll and Ross's (1980) factors extracted from factor analysis and Connor and Korajczyk's (1988) factors extracted from the asymptotic principal components clearly are the common factors that conform to the APT in spirit. Examples of "fundamental" factor models include the FF model and Carhart's (1997) four-factor model, which augments the FF model with a momentum factor. Such models, however, are empirically motivated because they are not derived from economic fundamentals but from accounting fundamentals.

The ICAPM has been more popular than the APT in recent years because many factors are proposed based on state variables that supposedly reflect the state of the economy; see LNS (2006) for a brief survey of the alternative factor models. Most models claim to explain the returns on the Fama–French 25 size-BM portfolios well. But based on simulations and real data, LNS (2006) find that most models perform poorly in explaining the asset-pricing anomalies when the test assets are expanded beyond the Fama–French size-BM portfolios. Also, BCS (1998) find that neither the Fama–French three-factor model nor an APT model based on Connor and Korajczyk's (1988) asymptotic principal components explains the size and BM effects well when the analysis is implemented on individual assets. Thus, the quest for a risk-based multifactor model that accounts for the asset-pricing anomalies continues.

3. An augmented five-factor model

3.1. Ferguson and Shockley's (2003) model: a review

This subsection briefly reviews the FS model, which is derived from Merton's ICAPM. The ICAPM suggests the following specification:

$$R_{it} - R_{ft} = \frac{\sigma_{im}}{\sigma_M^2}(R_{mt} - R_{ft}) + \frac{\sigma_{io}}{\sigma_o^2}(R_{ot} - R_{ft}), \tag{1}$$

where the subscript i indicates asset i, f denotes the riskless asset, M indicates the market portfolio, and O indicates the portfolio that hedges against changes in the opportunity set.

Assuming that investors have logarithmic utility so that they do not care about the changes in the investment opportunity set, FS (2003) show that the expected return can be related to relative leverage and relative distress even when the real world is "generated" by a single-factor CAPM. Specifically, suppose that the market is in equilibrium and the CAPM holds, the market portfolio M can be divided into two subportfolios: the economy's debt claims (D) and the economy's equity claims (E). In this equilibrium, firm i's true β can be written as

$$\beta_{S_i} = \frac{\sigma_{S_i,M}}{\sigma_M^2} = \frac{E}{M}\frac{\sigma_{S_i,E}}{\sigma_M^2} + \frac{D}{M}\frac{\sigma_{S_i,D}}{\sigma_M^2}, \tag{2}$$

where $\sigma_{S_i,M}$, $\sigma_{S_i,E}$, $\sigma_{S_i,D}$, and σ_M^2 are the stock's return covariances with the asset, equity, debt markets, and the return variance of the asset market, respectively. The proxy β of firm i's equity calculated against the economy's equity claims can be written as a transformation of the true β:

$$\hat{\beta}_{S_i}^E = \frac{\sigma_{S_i,E}}{\sigma_E^2} = \Phi^{-1}[\beta_{S_i} - \Omega\hat{\beta}_{S_i}^D], \tag{3}$$

where

$$\Phi = \frac{E}{M}\frac{\sigma_E^2}{\sigma_M^2}, \quad \Omega = \frac{D}{M}\frac{\sigma_D^2}{\sigma_M^2},$$

σ_E^2 and σ_D^2 are the return variances of the equity and debt markets, and $\hat{\beta}_{S_i}^D$ is the β of the firm's equity calculated against the economy's debt claims only.

Eq. (3) shows that the error in the proxy equity β ($\beta_{S_i}^E$) has two components: a scaling error (Φ^{-1}) that is common across all equities and a firm-specific error ($-\Omega\hat{\beta}_{S_i}^D$) that reflects the stock's covariance with the omitted debt claims. Because the total debt claims in the economy are not observable, FS (2003) suggest that factors formed on relative leverage and relative distress should provide the best complements to the equity market index for explaining the cross-sectional variation in stock returns. To capture this, they construct two portfolios to mimic the part of common return associated with relative leverage (based on D/E ratio) and the part of return associated with relative distress (based on Altman's (1968) Z), resulting in an alternative three-factor model as follows:

$$R_{it} - R_{ft} = \alpha_i + \beta_{i,RMRF} R_{RMRF,t} + \beta_{i,D/E} R_{D/E,t} + \beta_{i,Z} R_{Zt} + e_{it}, \tag{4}$$

where $R_{RMRF,t}$ is the excess return on the stock-only market portfolio, and $R_{D/E,t}$, and R_{Zt} are, respectively, returns on the two mimicking portfolios on leverage and distress. In the following, we refer to the two factors as the FS's (2003) factors, and the three-factor model as the FS's (2003) three-factor model or simply the FS model for short.

FS (2003) show that the two new factors subsume the explanatory power of SMB and HML in explaining the cross-sectional variation in the average returns on the 25 size-BM portfolios, thus supporting their viewpoint that asset-pricing anomalies such as the size and BM effects can be consistent with the single-factor CAPM.

The major contribution of FS (2003) is twofold. First, they show that the size and BM effects may arise as a result of measurement error in the market portfolio. Second, their new model outperforms the Fama–French model in explaining the cross-sectional average returns, suggesting that a static (unconditional) one-factor model alone can explain the asset-pricing anomalies; there is no need to rely on a multi-factor explanation.

Some interesting questions emerge from FS's (2003) findings. First, while the FS model outperforms the FF model cross-sectionally, does it really fully explain the size and BM effects? The question is important because although the FS factors outperform the FF factors in explaining the cross-sectional variation in stock returns, the FF model still has higher time-series explanatory power than the FS model (see FS, 2003, Table 5, p. 2572). If the FS model is a good pricing model, then the intercepts in (4) should be zero for all assets. However, FS (2003) find that the zero-intercept hypothesis of both the FS model and the FF model is rejected by the Gibbons, Ross, and Shanken (1989, hereafter GRS) test, suggesting that neither model is complete in describing stock returns.

Since the FF model, which is essentially a time-series model, is constructed based on empirical observations on cross-sectional regularities, one would have expected it to perform well both in the cross section and in time series. It is therefore surprising that the FF factors underperform the FS factors in terms of cross-sectional explanatory power. This gives rise to a second question: Why do SMB and HML capture more time-series variation in returns than do the FS factors? FS (2003) conjecture that the FF factors may have either reflected other state variables unrelated to relative leverage and distress or captured the time-varying nature of risk. In other words, they suggest that either a conditional CAPM or a multifactor model may provide a better account of both cross-sectional and time-series variation in stock returns. We explore this possibility in the next section.

A final question is related to LNS's (2006) comment: FS's (2003) empirical analysis is performed on the Fama–French size-BM portfolios; LNS (2006) show that empirical analysis based on the Fama–French size-BM portfolios is misleading because the portfolios have a strong factor structure. In this article, we advocate the use of an EIV-free approach developed by BCS (1998). Thus, we are able to explore further how the FS model performs in explaining the size and BM anomalies, in comparison with the FF model.

3.2. A leverage and distress augmented ICAPM

The empirical evidence indicates that the FS model performs better in explaining the cross-sectional variation of the 25 size-BM portfolios, whereas the FF model performs better in capturing the time-series variation. This suggests that neither model fully captures the patterns related to size and BM.

The size and BM effects seem to have two facets: in one aspect they reflect cross-sectional variation in returns that can be explained by a static single-factor model (i.e., the FS model) and in the other they also exhibit time-series patterns, which are captured by the FF model but not by the FS model. A problem that remains is therefore whether the time-series variation related to size and BM is associated with risk or due to mispricing. Rationally, this part of variation can be related to a time-varying market premium and/or an additional risk factor; otherwise it can be attributed to mispricing.

Note that although the FS model is derived from Merton's ICAPM, it is a static one-period model because the changes in the investment opportunity set are ignored (due to the assumption of logarithmic utility). Thus, it is possible that the FF factors (i.e., SMB and HML) may in fact proxy for changes in the investment opportunity set over time, while the FS factors capture the innovations in the market portfolio. In other words, without assuming logarithmic utility, there should in fact be a two-factor world: the first the market portfolio R_{mt} and the second the hedging portfolio R_{ot}.

Since the market portfolio is unobservable, it can be proxied by the FS model. In addition, since the hedging portfolio reflects the changes in the investment opportunity set *over time*, it might be proxied by SMB and HML, which better account for time-series variation in stock returns than the FS factors. In combination, this gives rise to an alternative five-factor model, which augments the Fama–French three-factor model with the two factors on relative distress and relative leverage. The *augmented five-factor model* has the following specification:

$$R_{jt} - R_{ft} = \alpha_j + \beta_{j1} R_{RMRF,t} + \beta_{j2} R_{SMB,t} + \beta_{j3} R_{HML,t} + \beta_{j4} R_{Z,t} + \beta_{j5} R_{D/E,t} + \varepsilon_{jt}. \quad (5)$$

If the model is correct, we expect the intercept to be zero. Also, the explanatory power of firm characteristics associated with size, BM, and momentum should disappear, once returns are properly adjusted by the model. This augmented five-factor model is only a slight extension of the FS model. In fact, FS (2003) also consider a similar model in which the fourth and fifth factors are SMB and HML, but the factors are orthogonalized so that they are independent of the FS factors (see the last panel of Table 5 in FS, 2003, p. 2573). However, the GRS test, which is essentially a test of unconditional pricing models, still rejects their five-factor model, thus suggesting that the unconditional five-factor model fails to fully account for stock returns. We examine whether an unconditional version or a conditional version of the augmented five-factor model can explain the asset-pricing anomalies using the BCS approach.

4. Econometric methodology

To test the competing asset pricing models, we follow the approach of BCS (1998), briefly described and discussed as follows. Suppose that returns are generated by an L-factor asset-pricing model:

$$\tilde{R}_{jt} = E(\tilde{R}_{jt}) + \sum_{k=1}^{L} \beta_{jk} \tilde{f}_{kt} + \tilde{e}_{jt}, \quad (6)$$

where the expected return is

$$E(\tilde{R}_{jt}) = R_{0t} + \sum_{k=1}^{L} \lambda_{kt}\beta_{jk}. \tag{7}$$

$\tilde{R}_{jt}$ is the return on security j, R_{0t} is the risk-free interest rate, β_{jk} is security j's loading of on factor k, and λ_{kt} is the risk premium associated with factor k. Substituting Eq. (7) into Eq. (6) gives the following result:

$$\tilde{R}_{jt} = R_{0t} + \sum_{k=1}^{L} \beta_{jk}\tilde{F}_{kt} + \tilde{e}_{jt}, \tag{8}$$

where $\tilde{F}_{kt} \equiv \lambda_{kt} + \tilde{f}_{kt}$ is the sum of the factor realization and its associated risk premium at month t.

A standard application of the Fama–MacBeth (1973) procedure would involve estimation of the following cross-sectional regression:

$$\tilde{R}_{jt} - R_{0t} = c_{0t} + \sum_{k=1}^{L} \lambda_{kt}\hat{\beta}_{jk} + \sum_{m=1}^{M} c_{mt}Z_{mjt} + \tilde{e}_{jt}, \tag{9}$$

where $\hat{\beta}_{jk}$ is estimated from Eq. (8), Z_{mjt} $(m = 1, 2, \ldots, M)$ is the value of (non-risk) characteristic m for security j in month t, and c_{mt} is the premium per unit of characteristic m in month t.

The null hypothesis is that expected returns of stocks depend only on factor loadings, while the alternative hypothesis is that security characteristics have incremental explanatory power over the factor model. Under the null hypothesis that expected returns depend only on the factor loadings, β_{jk}, the coefficients, c_{mt}'s $(m = 0, 1, \ldots, M)$, of security characteristics will be equal to zero.

The standard Fama–MacBeth two-pass approach suffers from the errors-in-variables (EIV) problem in testing Eq. (9) because the estimations of factor loadings are measured with errors. As a result, the risk premiums associated with the factors (λ_{kt}'s), and the coefficients associated with the characteristics (c_{mt}'s), will be biased. One way to deal with this measurement error problem is to use the information from the first-stage regressions to correct the coefficient estimates in the second-stage regressions (e.g., Fama and French, 1992.) BCS (1998), on the other hand, propose a novel approach to correct the bias. First, for each year, the factor loadings, β_{jk}'s, are estimated for all securities that have at least 24 return observations over the previous 60 months. As in BCS (1998), we estimate the factor loadings with a one-period-lag adjustment to correct for the nonsynchronous problem. The estimated risk-adjusted return on each of the securities, $\tilde{R}^*_{jt}$, for each month of the following year is then calculated as

$$\tilde{R}^*_{jt} \equiv \tilde{R}_{jt} - R_{0t} - \sum_{k=1}^{L} \hat{\beta}_{jk}\tilde{F}_{kt}, \tag{10}$$

for all j. The risk-adjusted returns from Eq. (10) are then used to test whether non-risk characteristics can describe the cross-sectional variation in expected returns. The estimation equation now becomes

$$\tilde{R}^*_{jt} = c_{0t} + \sum_{m=1}^{M} c_{mt} Z_{mjt} + \tilde{e}'_{jt}, \tag{11}$$

for all t. BCS (1998) show that the measurement error is now transferred to the dependent variable. Thus the estimates of $\hat{c}_{mt}$'s $(m = 0, 1, \ldots, M)$ from the OLS regression are unbiased.

Using data from the CRSP/Compustat, BCS (1998) find that non-risk firm characteristics retain significant explanatory power relative to the arbitrage pricing theory benchmark, with the common factors being the Fama and French's (1993) three factors, or determined using the Connor and Korajczyk's (1988) asymptotic principal component approach. In particular, BCS (1998) find that the momentum and trading volume effects persist when the analysis is repeated using the Fama–French factors, while the size and BM effects attenuate.

To summarize, the BCS approach has several advantages over the classical Fama–MacBeth (1973) cross-sectional regression. First, the BCS approach does not suffer from the EIV problem as in the Fama–MacBeth cross-sectional regression because the *estimated* factor loadings or market betas are not used as the independent variables. Second, the BCS methodology is implemented using single securities, thus avoiding the data-snooping biases that are inherent in portfolio-based approaches, as noted by Lo and MacKinlay (1990a). LNS (2006) also advise researchers to expand the set of test portfolios beyond size and BM portfolios. Without relying on the portfolio grouping procedure, the BCS approach retains the information in the individual securities. As a result, theoretically the BCS approach is more powerful statistically because a larger sample is used in the second-pass cross-sectional regression. Finally, the BCS methodology allows for tests of a wide array of asset-pricing models, both conditionally and unconditionally (see Avramov and Chordia, 2006). Notice that the original BCS approach is a test of the conditional asset-pricing model because the factor loadings are estimated over an "estimation window" prior to the second-pass cross-sectional regression. It can be easily extended to allow for testing unconditional asset-pricing models by, for example, replacing the estimates of factor loadings with the postranking estimates, as in Fama and French (1992) (see also Avramov and Chordia, 2006).

We consider three models: the FF model, the FS model, and the augmented five-factor model that combines the FF and the FS models. Under the null hypothesis, the coefficients on the non-risk security characteristics should be insignificantly different from zero. We consider eight non-risk security characteristics as the candidate variables in the alternative hypothesis. The firm size is included because of the importance of the size effect mentioned by Banz (1981) and Fama and French (1992). The BM ratio is used to capture the value effect documented by Fama and French (1992) and other literature. Amihud and Mendelson (1986) and Brennan and Subrahmanyam (1996) suggest that expected returns are affected by liquidity. We therefore include the dollar volume of trading, which is associated with liquidity. Moreover, because of the differences in trading mechanisms between the NYSE/AMEX and NASDAQ, we separate the variable into NYSE/AMEX-only and NASDAQ-only dollar volumes. Also, we add a dummy variable for stocks listed on the NASDAQ in the regression. Finally, we include three past-return variables as in BCS (1998), because numerous studies on contrarian/momentum strategies find that stock returns are correlated to their past returns (e.g., Lo and MacKinlay, 1990b; Jegadeesh and Titman, 1993).

Thus, we can rewrite our testing equation as

$$\tilde{R}^*_{jt} = c_0 + c_1\,\mathrm{NASDUM} + c_2\,\mathrm{NADVOL} + c_3\,\mathrm{NYDVOL} + c_4\,\mathrm{SIZE} \\ + c_5\,\mathrm{BM} + c_6\,\mathrm{RET2\text{-}3} + c_7\,\mathrm{RET4\text{-}6} + c_8\,\mathrm{RET7\text{-}12} + \varepsilon_{jt}, \quad (12)$$

where $\tilde{R}^*_{jt}$ denotes the risk-adjusted return estimated from either of the three competing asset pricing models. For example, for the augmented five-factor model, $\tilde{R}^*_{jt} = \tilde{R}_{jt} - \tilde{R}_{0t} - \hat{\beta}_{j1}R_{RMRF,t} + \hat{\beta}_{j2}R_{SMB,t} + \hat{\beta}_{j3}R_{HML,t} + \hat{\beta}_{j4}R_{Z,t} + \hat{\beta}_{j5}R_{D/E,t}$. The independent variables are defined as follows.

- NASDUM: A dummy variable equals one if the security trades on the NASDAQ, and zero otherwise.
- NADVOL: The natural logarithm of the dollar volume of trading in the security in the second to last month if the security trades on the NASDAQ, and zero otherwise.[6]
- NYDVOL: The natural logarithm of the dollar volume of trading in the security in the second to last month if the security trades on the NYSE/AMEX, and zero otherwise.
- 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 market and book values. As in Fama and French (1992), the value of BM for July of year t to June of year $t+1$ was computed using accounting data at the end of year $t-1$, and BM 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.
- RET2-3: The natural logarithm of the cumulative return over the two months ending at the beginning of the previous month.
- RET4-6: The natural logarithm of the cumulative return over the three months ending three months previously.
- RET7-12: The natural logarithm of the cumulative return over the six months ending six months previously.

We partition the past one-year return into three shorter-interval return variables following BCS (1998).

For each month, we regress the risk-adjusted returns on the eight characteristics to estimate the coefficients $\hat{c}_m$'s for $m = 0, 1, \ldots, 8$. Then the time-series averages of the coefficients associated with the characteristics are computed and tested based on the Newey–West robust t-statistic. If a factor-pricing model fully explains the expected stock returns, the coefficients $\hat{c}_m$'s should all be insignificantly different from zero.

Note that since BCS (1998) estimate betas each year in a first-pass regression using returns over past 60 months, the test based on the rolling regressions can be viewed as a test of the conditional multifactor model because the factor loadings are allowed to change over time. For a test of an unconditional version of the model, the factor loadings in

[6] To account for trading between market makers, we divide the trading volume of NASDAQ firms by 2 (see, for example, Atkins and Dyl, 1997).

Table 1
Summary statistics of variables.

Variable	Mean	Median	Std. dev.
Return (%)	1.34	1.33	6.27
Firm size (in millions of dollars)	0.80	0.49	0.67
Book-to-market ratio	0.86	0.78	0.35
Dollar volume (in millions of dollars)	0.79	0.23	1.23
Number of firms	2754.08	3182.00	1039.00

The summary statistics represent the time-series averages of cross-sectional means over 456 months from January 1967 through December 2004. Each stock satisfies the following criteria: (1) its return in the current month and in 24 of the previous 60 months is available from CRSP, and sufficient data are available to calculate the size, price and dollar volume as of the second to previous month; (2) sufficient data are available on the Compustat tapes to calculate the book-to-market ratio as of December of the previous year; and (3) we do not include firms until they have appeared on Compustat for two years. The variables are defined as follows: Return is the raw return. Firm size is the market value of the equity of the firm as of the end of the second to the last month. BM ratio is the ratio of book equity and market equity at the end of the second to last month. Dollar volume is the total dollar volume of trading in the security in the second to last month. We also report the average number of firms that are included in our sample.

Eq. (10) are replaced with the time-series estimates over the whole sample period. The procedure, however, differs from that of Fama and French (1992) in that here the factor loadings are estimated on each security, whereas Fama and French's procedure estimates the factor loadings on the 100 size-BM portfolios, and assigns the portfolios' estimates to the individual securities.

5. Data

The original data consist of monthly returns and other firm characteristics for a sample of common stocks of U.S. companies for the period January 1964 to December 2004. A stock to be included in a given month must satisfy the following criteria. First, its returns in the current month and in at least 24 out of the previous 60 months must be available from CRSP, and sufficient data must be available to calculate the size and dollar volume as of the second to last month. Second, sufficient data must be available on the Compustat tapes to calculate the book-to-market ratio as of December of the previous year. Third, as in Fama and French (1992), financial firms are excluded from our sample. Finally, as mentioned in Fama and French (1993), we do not include firms until they have appeared on Compustat for two years to avoid the survival bias inherent in the way Compustat adds firms to its tapes. This screening process yields an average of 2,754 stocks per month beginning from January 1967.

Since the data on NASDAQ-listed companies are only available in CRSP after 1972, tests based on the NASDAQ-related variables (e.g., NASDUM, NADVOL, and NYDVOL) are restricted to a sample period of 1975–2004.[7] Table 1 reports the time-series descriptive statistics of the cross-sectional means of the security characteristics and returns.

[7]This is because we require at least 24 out of the previous 60 months to estimate the factor loadings.

Table 2
Regression results on excess returns.

	Full period		Jan. only		Non-Jan.	
	Model 1 (1)	Model 2 (2)	Model 1 (3)	Model 2 (4)	Model 1 (5)	Model 2 (6)
Intercept	2.243***	2.029***	12.729***	9.763***	1.303***	1.335***
	(5.26)	(5.74)	(8.90)	(4.95)	(3.19)	(3.92)
NASDUM		−0.810		2.818		−1.136
		(−0.28)		(1.55)		(−0.37)
NADVOL		0.354		0.617		0.330
		(0.43)		(1.46)		(0.37)
NYDVOL		−0.020		0.460		−0.063
		(−0.33)		(1.55)		(−1.06)
SIZE	−0.189***	−0.125	−1.692***	−2.168***	−0.054	0.058
	(−3.60)	(−1.28)	(−8.19)	(−5.55)	(−1.10)	(0.61)
BM	0.324***	0.285***	0.357	0.467*	0.321***	0.268***
	(4.66)	(4.93)	(1.40)	(1.99)	(4.52)	(4.57)
RET2-3	0.387	0.870***	−6.100***	−5.733***	0.969***	1.462***
	(1.29)	(3.22)	(−4.57)	(−4.54)	(3.11)	(5.31)
RET4-6	0.656***	0.908***	−4.724***	−4.254***	1.139***	1.371***
	(2.71)	(3.80)	(−3.24)	(−2.86)	(5.01)	(6.38)
RET7-12	0.839***	0.912***	−2.172***	−2.047**	1.109***	1.177***
	(4.60)	(5.04)	(−2.86)	(−2.65)	(5.99)	(6.48)

Coefficient estimates are time-series averages of cross-sectional OLS regressions. The dependent variable is the excess return. The independent variables are defined as follows. NASDUM is a dummy variable for NASDAQ-listed stocks, which equals one if the security trades on the NASDAQ, and zero otherwise. SIZE represents logarithm of the market capitalization of firms in millions of dollars. BM is the logarithm of the ratio of book value of equity plus deferred taxes to market capitalization. NADVOL is the value of dollar volume if the stock trades on the NASDAQ, and zero otherwise; NYDVOL is the value of dollar volume if the stock trades on the NYSE/AMEX; and zero otherwise. RET2-3, RET4-6, RET7-12 equal the logarithms of the cumulative returns over the second through third, fourth through sixth, and seventh through 12th months prior to the current month, respectively. All coefficients are multiplied by 100. The *t*-statistics are in parentheses. * denotes significance at the 10% level, ** denotes significance at the 5% level, and *** denotes significance at the 1% level.

To calculate FS's (2003) D/E and Z portfolios, we retrieve the required accounting variables from Compustat, including net working capital, total book assets, retained earnings, book debt (all from balance sheets), earnings before interest and taxes, and total sales revenue (both from income statements). The D/E and Z portfolios are constructed over the period from July 1964 to December 2004. The details of the portfolio formation process are presented in the Appendix.

Table 2 presents the results of the Fama–MacBeth cross-sectional regression by regressing excess returns on various firm characteristics outlined in the previous section. The first two columns of Table 2 present the results based on different sets of explanatory variables. The estimation is also implemented on January and non-January observations to see how January seasonality affects the cross-sectional patterns of firm characteristics.

Overall, the results in Table 2 indicate a significant small-firm effect, which is significant only in January. The BM premium and the momentum premium are both significantly positive, but mostly significant in non-January months. Thus, the results confirm the momentum literature that there is an intermediate-term momentum effect in the U.S. stock market. Also, there is a strong return reversal in January months because the coefficients of the lagged returns are all significantly negative in January.

6. Empirical results

We first present the results of the FS model, along with the FF model, in explaining various asset pricing anomalies. The results on the augmented five-factor model are then presented.

6.1. Do relative distress and relative leverage really explain asset-pricing anomalies?

To examine whether relative distress and relative leverage fully explain the asset-pricing anomalies, we first perform a Fama–MacBeth regression of excess returns on the factor loadings on the FS factors, factor loadings on the FF factors, and various firm characteristics. We follow Fama and French's (1992) procedure by assigning the post-ranking factor loadings of a portfolio to its component stocks each year. The results are

Table 3
Fama–MacBeth regression results.

	Model (1)	Model (2)	Model (3)	Model (4)	Model (5)	Model (6)
Intercept	0.409*	0.457*	3.490***	2.381***	3.602***	4.237***
	(1.66)	(1.67)	(5.76)	(4.57)	(5.69)	(7.87)
β_{SMB}	0.423**		−0.780***		−1.221***	−1.676 ***
	(2.44)		(−3.75)		(−3.92)	(−5.45)
β_{HML}	0.893***		0.073		−0.175	−0.167
	(4.92)		(0.45)		(−0.68)	(−0.69)
β_Z		0.495		0.618***	−0.318	−0.672**
		(1.39)		(2.64)	(−1.16)	(−2.42)
$\beta_{D/E}$		1.768***		0.706**	0.129	−0.276
		(3.44)		(1.98)	(0.27)	(−0.61)
SIZE			−0.365***	−0.253***	−0.368***	−0.479***
			(−5.05)	(−4.50)	(−4.98)	(−7.66)
BM			0.191***	0.220***	0.175***	0.164***
			(2.99)	(3.67)	(2.89)	(2.92)
RET2-3						0.671**
						(2.28)
RET4-6						0.871***
						(3.62)
RET7-12						0.950***
						(5.29)

Coefficient estimates are time-series averages of cross-sectional OLS regressions. The dependent variable is the excess return. The independent variables are defined as follows. β_{SMB} and β_{HML} are the post-ranking estimators of factor loadings based on the Fama–French three-factor model. β_Z and $\beta_{D/E}$ are the post-ranking estimators of factor loadings based on the Ferguson–Shockley model. The factor loadings are estimated following Dimson's (1979) procedure with one period lag using 100 size-BM-sorted equally-weighted portfolios. SIZE represents logarithm of the market capitalization of firms in millions of dollars. BM is the logarithm of the ratio of book value of equity plus deferred taxes to market capitalization. RET2-3, RET4-6, RET7-12 equal the logarithms of the cumulative returns over the second through third, fourth through sixth, and seventh through 12th months prior to the current month, respectively. All coefficients are multiplied by 100. The *t*-statistics are in parentheses. * denotes significance at the 10% level, ** denotes significance at the 5% level, and *** denotes significance at the 1% level.

presented in Table 3. Since the asset-pricing anomalies, especially the small-firm effect, appear to exhibit seasonal patterns, we examine the seasonality by splitting the sample into January and non-January observations.

Table 3 indicates that both FS and the FF factors carry significant factor risk premiums when they are estimated without the firm characteristics (see Models (1) and (2) in Table 3). However, the coefficients of size and BM remain significant throughout all scenarios. In the last column of Table 3 where all factor loadings are jointly estimated with the inclusion of all firm characteristics, the result indicates that the premiums on the FF and FS factors become negative, whereas all asset-pricing anomalies retain their significance. Thus, the results indicate that neither the FS model nor the FF model explains the asset-pricing anomalies, at least unconditionally.

The results based on the BCS regression are reported in Tables 4 and 5. In Table 4, we report the results on the unconditional tests in which the factor loadings on a security are

Table 4
Regression results based on unconditional tests.

	Fama–French factor-adjusted returns			Ferguson–Shockley factor-adjusted returns		
	Full period (1)	Jan. only (2)	Non-Jan. (3)	Full period (4)	Jan. only (5)	Non-Jan. (6)
Intercept	0.467*	3.775***	0.176	0.162	0.993	0.089
	(1.84)	(4.43)	(0.64)	(0.63)	(0.82)	(0.33)
NASDUM	−1.415	4.580**	−1.941	−1.158	5.226**	−1.719
	(−0.33)	(2.09)	(−0.42)	(−0.29)	(2.28)	(−0.39)
NADVOL	0.472	0.019	0.512	0.386	−0.502	0.464
	(0.39)	(0.04)	(0.39)	(0.33)	(−0.86)	(0.37)
NYDVOL	−0.021	0.493**	−0.067	−0.033	0.269	−0.059
	(−0.50)	(2.12)	(−1.57)	(−0.72)	(1.25)	(−1.25)
SIZE	−0.042	−1.705***	0.104*	0.017	−0.901**	0.098
	(−0.69)	(−4.82)	(1.80)	(0.24)	(−2.63)	(1.33)
BM	0.125***	−0.153	0.150***	0.158***	0.132	0.160***
	(4.25)	(−1.30)	(5.15)	(4.61)	(1.25)	(4.65)
RET2-3	0.143	−4.829***	0.580**	0.140	−3.870***	0.492*
	(0.51)	(−4.38)	(2.15)	(0.56)	(−3.54)	(1.92)
RET4-6	0.628***	−2.775**	0.927***	0.458**	−2.791**	0.743***
	(2.82)	(−2.40)	(4.69)	(1.97)	(−2.39)	(3.42)
RET7-12	0.559***	−2.684***	0.844***	0.552***	−3.077***	0.871***
	(3.23)	(−3.64)	(5.14)	(2.89)	(−4.45)	(4.75)

Coefficient estimates are time-series averages of cross-sectional OLS regressions. The dependent variable in scenarios (1)–(3) is the risk-adjusted returns using the Fama–French factors, while in scenarios (4)–(6) it is the risk-adjusted excess returns using the Ferguson-Shockley factors (factor loadings are estimated following Dimson's (1979) procedure with one period lag in each case). For the unconditional case, we estimate the factor loadings using the whole sample period. The independent variables are defined as follows. NASDUM is a dummy variable for NASDAQ-listed stocks, which equals one if the security trades on the NASDAQ, and zero otherwise. SIZE represents logarithm of the market capitalization of firms in millions of dollars. BM is the logarithm of the ratio of book value of equity plus deferred taxes to market capitalization. NADVOL is the value of dollar volume if the stock trades on the NASDAQ, and zero otherwise; NYDVOL is the value of dollar volume if the stock trades on the NYSE/AMEX; and zero otherwise. RET2-3, RET4-6, RET7-12 equal the logarithms of the cumulative returns over the second through third, fourth through sixth, and seventh through 12th months prior to the current month, respectively. All coefficients are multiplied by 100. The *t*-statistics are in parentheses. * denotes significance at the 10% level, ** denotes significance at the 5% level, and *** denotes significance at the 1% level.

Table 5
Regression results based on conditional tests.

	Fama–French factor-adjusted returns			Ferguson–Shockley factor-adjusted returns		
	Full period (1)	Jan. only (2)	Non-Jan. (3)	Full period (4)	Jan. only (5)	Non-Jan. (6)
Intercept	0.359	2.103**	0.205	−0.069	−0.315	−0.047
	(1.32)	(2.25)	(0.70)	(−0.26)	(−0.29)	(−0.16)
NASDUM	−1.363	5.688**	−1.983	−1.649	4.737**	−2.210
	(−0.36)	(2.53)	(−0.49)	(−0.47)	(2.21)	(−0.58)
NADVOL	0.499	0.246	0.521	0.502	0.101	0.538
	(0.47)	(0.45)	(0.45)	(0.50)	(0.18)	(0.50)
NYDVOL	0.015	0.688**	−0.045	−0.026	0.494**	−0.071
	(0.29)	(2.71)	(−0.88)	(−0.52)	(2.08)	(−1.36)
SIZE	−0.078	−1.706***	0.065	0.045	−0.960**	0.134
	(−1.07)	(−4.56)	(0.90)	(0.54)	(−2.69)	(1.52)
BM	0.169***	0.078	0.177***	0.209***	0.517***	0.182***
	(3.28)	(0.43)	(3.33)	(3.57)	(3.41)	(2.92)
RET2-3	−0.084	−5.986***	0.434	0.082	−4.964***	0.526*
	(−0.24)	(−3.54)	(1.35)	(0.26)	(−3.19)	(1.81)
RET4-6	0.854***	−3.611**	1.247***	0.613**	−3.623**	0.985***
	(2.87)	(−2.43)	(4.40)	(2.09)	(−2.44)	(3.50)
RET7-12	0.067	−4.770***	0.492**	−0.070	−5.694***	0.425
	(0.26)	(−5.11)	(2.08)	(−0.24)	(−5.97)	(1.60)

Coefficient estimates are time-series averages of cross-sectional OLS regressions. The dependent variable in scenarios (1)–(3) is the risk-adjusted returns using the Fama–French factors, while in scenarios (4)–(6) it is the risk-adjusted excess returns using the Ferguson–Shockley factors (factor loadings are estimated following Dimson's (1979) procedure with one period lag in each case). The independent variables are defined as follows. NASDUM is a dummy variable for NASDAQ-listed stocks, which equals one if the security trades on the NASDAQ, and zero otherwise. SIZE represents logarithm of the market capitalization of firms in millions of dollars. BM is the logarithm of the ratio of book value of equity plus deferred taxes to market capitalization. NADVOL is the value of dollar volume if the stock trades on the NASDAQ, and zero otherwise; NYDVOL is the value of dollar volume if the stock trades on the NYSE/AMEX; and zero otherwise. RET2-3, RET4-6, RET7-12 equal the logarithms of the cumulative returns over the second through third, fourth through sixth, and seventh through 12th months prior to the current month, respectively. All coefficients are multiplied by 100. The *t*-statistics are in parentheses. * denotes significance at the 10% level, ** denotes significance at the 5% level, and *** denotes significance at the 1% level.

estimated over the whole sample. In Table 5, the results are based on the standard BCS test, which is a test of conditional asset-pricing models.

The empirical results for the unconditional asset-pricing models are summarized as follows.

1. Size effect. The results in Table 4 indicate that the coefficient of firm size becomes insignificantly different from zero when returns are adjusted by either the FF model or the FS model.[8] However, the small-size premium remains significant in January months,

[8]While not reported here, we find that the small-firm premium is still significant when the independent variables exclude the dummy variable and the turnover of NASDAQ firms (i.e., NASDUM and NADVOL), which is consistent with BCS (1998), suggesting that the small-firm premium might be related to trading volume or a NASDAQ dummy.

regardless of the models for return adjustment. As we shall show later, the small-firm premium remains as a January phenomenon that cannot be explained by any of the three asset-pricing models we discuss here.

2. BM effect. The BM premium is significantly positive, and cannot be explained by either of the two models. In comparison, the coefficient of BM under the Ferguson–Shockley adjustment is even larger than that under the Fama–French adjustment. For example, with the Fama–French adjusted returns as the dependent variable, the BM coefficient is 0.125 for Model 1. In comparison, the coefficient is 0.158 with the Ferguson–Shockley adjustment (see Model 4). Also, the results show that the BM premiums mainly come from non-January months.
3. Momentum effect. Table 4 indicates that the momentum effect remains mostly intact regardless of the models for return adjustment. The only difference is that the coefficient of the short-term return (RET2-3) becomes insignificant, and it is still significant in non-January months. In comparison with the results in BCS (1998) where the sample only covers the period 1966–1995, our empirical results, which expands the sample period to 1966–2004, indicate that the pattern of momentum effect remains mostly the same, but appears to be weaker. Overall, the results indicate that neither the FS model nor the FF model fully explains the momentum effect. The results are consistent with the findings of Avramov and Chordia (2006).
4. The zero-intercept hypothesis. An interesting result from Table 4 is that the intercept is significantly different from zero under the FF adjustment, but is insignificant under the FS adjustment. This means that although the FS model fails to fully explain the asset-pricing anomalies, there is no "systematic" departure of returns from the pricing of the FS model that bears significant abnormal returns. In other words, the results somehow supportFS's (2003) findings that the FS model provides a good account of the cross-sectional variation in average returns.

Table 5 reports the results for tests of the conditional models. The patterns are mostly the same as the unconditional tests in Table 4 with only a few exceptions. First, the BM premium remains significant, but the coefficient is slightly larger than that in Table 4. Second, the momentum effect is also weakened; for both the FF and FS return adjustments, only the coefficient of RET4-6 remains significant. The results suggest that the two asset-pricing models capture the momentum effects better in a conditional setting. This also means that part of the momentum effect is related to time-varying risk. Third, the coefficients of NASDUM and NYDVOL become significant in January, indicating that the January seasonality is yet another stubborn anomaly that deserves further investigation.

Overall, the results indicate that the FS model subsumes the FF model as an asset-pricing model as there is no systematic departure from the pricing of FS model; the zero-intercept hypothesis cannot be rejected with the FS-model adjustment. Nevertheless, both models fail to account for the BM and the momentum effects, either conditionally or unconditionally. The empirical evidence on the FS model echoes Lewellen and Nagel's (2006) findings that the conditional CAPM cannot fully explain the asset-pricing anomalies.

6.2. The performance of the augmented five-factor model

So far we have completed the examination of the competing explanatory ability of the FS model and the FF model in explaining the existing asset-pricing anomalies. The results are somehow disappointing because both models fail to fully explain most of the anomalies. Thus, although factors on relative distress and relative leverage provide some explanatory power on stock returns, the FS model, like the FF model, remains as an incomplete asset-pricing model in explaining the cross-sectional variation in stock returns.

Why do the FF model and the FS model fail to account for the asset-pricing anomalies? As discussed in Section 3, our conjecture is that the FS factors in combination with the equity-only market index proxy the "true" market portfolio, while the FF factors proxy the hedging portfolio against the changes in the investment opportunity set over time. To test this conjecture, we perform the BCS regression on the augmented five-factor model, and examine how the model explains the existing anomalies.

Table 6 presents the results for the unconditional and conditional versions of the augmented five-factor model. Empirical results are presented on the full sample, and on two subsamples, i.e., January and non-January observations.

Let us first focus on the results of the full sample. The left panel of Table 6 indicates that with the return adjustment based on the unconditional five-factor model, the BM effect remains strongly significant. The momentum effect is also significant; both the coefficients of RET4-6 and RET7-12 are significantly positive. The results are not materially different from those of the unconditional FS model reported in Table 4. This means that incorporating SMB and HML factors into the FS model does not provide significant incremental explanatory ability of the pricing anomalies.

The right panel of Table 6 presents the results for the conditional model. The results for the conditional augmented five-factor model are fascinating. The results indicate that the significance of *all* coefficients, including the intercept, disappears. In comparison with the conditional FS model presented in Table 5, the conditional augmented five-factor model provides additional explanatory power on the BM effect and the momentum effect. In addition, the intercept is also closer to zero; the intercept is -0.027 with a t-statistic of -0.08.

Next, let us look at the results on the January and non-January subsamples. For January observations, the results in Table 6 indicate that several firm characteristics remain statistically significant. Clearly the January small-firm effect cannot be explained by the augmented five-factor model, either conditionally or unconditionally. The January return reversals cannot be explained by the augmented five-factor model either; the coefficients of RET2-3 and RET7-12 are significantly negative in January months, regardless of the models for return adjustment. For non-January months, Table 8 indicates that the unconditional augmented model still fails to explain several pricing anomalies, but the conditional version of the augmented model performs very well in non-January months.

To further evaluate the performance of the augmented five-factor model over time, we conduct subperiod analysis by dividing the original 30-year sample into three 10-year subsamples. Table 7 presents the results on subperiods.

For the unconditional five-factor model, the empirical results reveal that the size effect disappears, but the BM premium is significant for the last two subperiods, 1985–1994 and 1995–2004. For the momentum effect, the coefficient of RET7-12 is significant for the first two subperiods and that of RET4-6 is significant for the first subperiod. The momentum effect disappears in the last subperiod, 1995–2004. The results are consistent with the results in Table 6, suggesting that the momentum effect does not survive as an asset-pricing anomaly.

Table 6
Regression results of the augmented five-factor model.

	Unconditional tests			Conditional tests		
	Full period (1)	Jan. only (2)	Non-Jan. (3)	Full period (4)	Jan. only (5)	Non-Jan. (6)
Intercept	0.211	2.088**	0.046	−0.027	0.070	−0.036
	(0.86)	(2.64)	(0.17)	(−0.08)	(0.06)	(−0.10)
NASDUM	−3.277	3.435	−3.867	−2.713	5.628**	−3.446
	(−0.91)	(1.47)	(−0.99)	(−0.53)	(2.11)	(−0.62)
NADVOL	0.926	−0.380	1.041	0.832	−0.250	0.927
	(0.90)	(−0.66)	(0.93)	(0.57)	(−0.40)	(0.58)
NYDVOL	−0.052	0.209	−0.075	−0.007	0.489*	−0.050
	(−1.12)	(1.11)	(−1.55)	(−0.12)	(2.03)	(−0.89)
SIZE	0.041	−0.865***	0.121*	−0.004	−0.996***	0.083
	(0.64)	(−3.01)	(1.86)	(−0.05)	(−2.81)	(1.06)
BM	0.127***	0.111	0.128***	0.104	0.297*	0.087
	(4.14)	(1.10)	(4.18)	(1.31)	(1.81)	(1.03)
RET2-3	−0.051	−3.609***	0.262	−0.386	−6.074***	0.114
	(−0.12)	(−4.60)	(0.63)	(−0.79)	(−3.12)	(0.25)
RET4-6	0.518**	−1.629	0.706***	0.283	−2.379	0.517
	(2.00)	(−1.49)	(2.77)	(0.69)	(−1.36)	(1.26)
RET7-12	0.615***	−3.267***	0.956***	−0.541	−7.212***	0.046
	(2.92)	(−4.20)	(4.68)	(−1.37)	(−4.82)	(0.12)

Coefficient estimates are time-series averages of cross-sectional OLS regressions. The dependent variable is the risk-adjusted returns using the augmented five-factor model. Scenarios (1)–(3) are based on unconditional tests, while scenarios (4)–(6) are based on conditional tests. The independent variables are defined as follows. NASDUM is a dummy variable for NASDAQ-listed stocks, which equals one if the security trades on the NASDAQ, and zero otherwise. SIZE represents logarithm of the market capitalization of firms in millions of dollars. BM is the logarithm of the ratio of book value of equity plus deferred taxes to market capitalization. NADVOL is the value of dollar volume if the stock trades on the NASDAQ, and zero otherwise; NYDVOL is the value of dollar volume if the stock trades on the NYSE/AMEX; and zero otherwise. RET2-3, RET4-6, RET7-12 equal the logarithms of the cumulative returns over the second through third, fourth through sixth, and seventh through 12th months prior to the current month, respectively. All coefficients are multiplied by 100. The *t*-statistics are in parentheses. * denotes significance at the 10% level, ** denotes significance at the 5% level, and *** denotes significance at the 1% level.

For the conditional five-factor model, the right panel of Table 7 indicates that most firm characteristics lose their explanatory ability in subperiods. The BM effect is only marginally significant for 1975–1984 and the coefficient of RET7-12 is negative for 1995–2004. Overall, the conditional model also performs reasonably well in subperiods.

An interesting phenomenon is that the coefficient of the NASDAQ dummy (NASDUM) remains significant for two subperiods, 1985–1994 and 1995–2004. This suggests that NASDAQ stocks on average earn higher risk-adjusted returns than those listed on the NYSE and AMEX for 1985–2004.

Overall, the results suggest that the augmented five-factor model is capable of explaining most well-known anomalies. This also means that the pricing anomalies associated with size, BM, and momentum are consistent with the ICAPM, with the factor loadings and prices of risks being allowed to vary in a sample manner. In comparison, Avramov and Chordia (2006) find the size and value effects are explained by a conditional Fama–French

Table 7
Regression results of the augmented five-factor model in subperiods.

	Unconditional tests			Conditional tests		
	1975–1984 (1)	1985–1994 (2)	1995–2004 (3)	1975–1984 (4)	1985–1994 (5)	1995–2004 (6)
Intercept	−0.052	0.404	−0.175	0.054	0.336	−1.034
	(−0.14)	(1.22)	(−0.31)	(0.12)	(0.84)	(−1.40)
NASDUM	−1.220	1.931***	1.204**	−1.839	1.981***	2.204***
	(−1.02)	(3.50)	(2.30)	(−1.30)	(3.52)	(3.17)
NADVOL		−0.446***	−0.134		−0.395**	−0.229
		(−3.46)	(−0.52)		(−2.56)	(−0.92)
NYDVOL	−0.046	−0.027	0.030	0.037	−0.009	0.084
	(−0.79)	(−0.49)	(0.26)	(0.47)	(−0.12)	(0.62)
SIZE	0.060	−0.023	−0.063	−0.101	−0.042	−0.044
	(0.76)	(−0.31)	(−0.40)	(−0.93)	(−0.42)	(−0.23)
BM	0.097	0.157***	0.157***	0.223*	0.020	0.123
	(1.49)	(3.65)	(3.97)	(1.71)	(0.19)	(0.77)
RET2-3	−0.023	−0.139	−0.244	−0.712	0.094	−0.971
	(−0.04)	(−0.34)	(−0.25)	(−0.64)	(0.17)	(−1.29)
RET4-6	0.961*	0.097	0.322	−0.308	0.498	0.434
	(1.76)	(0.32)	(0.65)	(−0.35)	(1.39)	(0.56)
RET7-12	1.015***	0.621**	0.135	0.220	−0.190	−1.724**
	(2.93)	(2.53)	(0.31)	(0.34)	(−0.32)	(−2.51)

We separate our sample into two subperiods: 1975–1984, 1985–1994, and 1995–2004. Coefficient estimates are time-series averages of cross-sectional OLS regressions. Scenarios (1)–(3) are based on unconditional tests, while scenarios (4)–(6) are based on conditional tests. The independent variables are defined as follows. NASDUM is a dummy variable for NASDAQ-listed stocks, which equals one if the security trades on the NASDAQ, and zero otherwise. SIZE represents logarithm of the market capitalization of firms in millions of dollars. BM is the logarithm of the ratio of book value of equity plus deferred taxes to market capitalization. NADVOL is the value of dollar volume if the stock trades on the NASDAQ, and zero otherwise; NYDVOL is the value of dollar volume if the stock trades on the NYSE/AMEX; and zero otherwise. RET2-3, RET4-6, RET7-12 equal the logarithms of the cumulative returns over the second through third, fourth through sixth, and seventh through 12th months prior to the current month, respectively. All coefficients are multiplied by 100. The *t*-statistics are in parentheses. * denotes significance at the 10% level, ** denotes significance at the 5% level, and *** denotes significance at the 1% level.

three-factor model, but the explanatory power of past return remains robust. They attribute the past return effect to model mispricing that varies with macroeconomic variables. We find that the momentum effect can be consistent with a risk-based explanation.

To sum up, we find that neither the Fama–French factor model nor the Ferguson-Shockley factor model succeeds in explaining the common asset-pricing anomalies, especially the BM and momentum anomalies. Further analysis reveals that those asset-pricing anomalies can be explained by a conditional augmented five-factor model.

7. Conclusion

Among the competing rational and behavioral arguments on the size and BM anomalies, FS (2003) argue that the two anomalies are induced because of the measurement error of the equity-only market betas. By incorporating firms' relative leverage and relative distress to capture the economy's total debt claims, they show that their model provides more information than Fama and French's (1993) three-factor model in explaining the size and BM anomalies. In this paper, we conduct various tests to examine the competing explanatory power of the FS model and the Fama–French three-factor model in explaining the existing anomalies.

Our empirical results show that, when individual stocks' excess returns are adjusted by Fama and French's (1993) model or FS's (2003) model, the explanatory power of firm size and BM remains significant. As a further exploration, we find that a conditional version of the augmented five-factor model, which combines Fama and French's (1993) and FS's (2003) factors, is able to explain almost all anomalies, including the BM and the momentum effects, neither of which is captured by either of the two models alone.

Our paper has some important implications. First, asset-pricing anomalies such as size and BM have two facets: one related to the cross section and the other related to time series. We find that FS's three-factor model accounts for the cross-sectional variation in stock returns, but the time-series variation in stock returns is better captured by the Fama–French factors, i.e., SMB and HML. Second, most asset-pricing anomalies can be explained in a conditional multifactor framework; there is no need to attribute the pricing anomalies to market mispricing. Of course, the single-factor CAPM alone does not suffice to explain the anomalies. Third, our model conforms more in spirit to Merton's (1973) ICAPM, rather than to Ross's (1976) APT.

Finally, it is worthwhile to note that NASDAQ firms, small-firms, and loser firms experience positive risk-adjusted returns in January. Hence, the January seasonality remains a topic that deserves further investigation.

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Appendix: The formation of D/E and Z portfolios

Following Ferguson and Shockley's (2003) methodology, we construct two portfolios to mimic the part of common return associated with relative leverage (based on the ratio of debt-to-equity) and the part of return associated with relative distress (based on Altman's (1968) Z). The way to generate the leverage and distress return time series $R^{D/E}$ and R^{Z} is the same with the construction of Fama and French's (1993) SMB and HML factors. In June of each year t, firms are assigned to one of three book debt-to-market equity (BD/ME) portfolios based on the one-third and two-third percentile cuts determined only from the NYSE firms in the sample. Independently and simultaneously, firms are assigned to one of two Altman's Z portfolios: $Z \leq 2.675$ and $Z > 2.675$. The Altman's Z is defined as follows:

$$Z = 1.2\frac{WC}{TA} + 1.4\frac{RE}{TA} + 3.3\frac{EBIT}{TA} + 0.6\frac{ME}{BD} + 1.0\frac{S}{TA},$$

where WC is net working capital, TA is total book assets, RE is retained earnings, BD is book debt (all from balance sheets), $EBIT$ is earnings before interest and taxes, and S is total sales revenue (both from income statements), and ME is market value of equity.

To be included in portfolio construction, a firm must have CRSP market equity for December of $(t-1)$, and data available for all the relevant Compustat items for Z in $(t-1)$. Only firms with ordinary common equity, as defined by CRSP, are used to form the leverage and distress portfolios.

The intersection of the two sorts based on BD/ME and Z results in six debt-to-equity/Z portfolios as of June of each year. For July of t through June of $(t+1)$, the return on each portfolio is calculated as the value-weighted average return of the stocks in the portfolio (where value is the market value of equity at the beginning of the return month). In the end, we have six return series that cover the 486 months from July 1964 through December 2004. In each month t, $R_t^{D/E}$ is calculated as the simple average return of the two Z portfolios within D/E portfolio 3 (the highly levered firms) minus the simple average return of the two Z portfolios within D/E portfolio 1 (the least levered firms). Similarly, R_t^Z is the simple average return of the three D/E portfolios within Z portfolio 2 (high-Z firms) minus the simple average return of the three D/E portfolios within Z portfolio 1 (the low-Z firms).

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