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Adverse selection and seasoned security offerings

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To Hanne, Peder and Vera

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Chapter 1

Introduction

1.1 The intellectual foundation

In order to facilitate trade between two parties, the seller must be able to convince the buyer that the good up for sale has no hidden defects that will reduce the value of the good when they surface. However, in many situations the seller is unable to give a guarantee about the quality of the good, and the only way to create a trade is to reduce the price dramatically. A well known example of this problem is the puzzling fact that used cars sell at a huge discount relative to new cars. In one of the first formal discussions of this problem, Akerlof (1970) illustrates his results with the following intuitive explanation of the used car discount: The seller of a used car is likely to have more information about the true quality of the car than potential buyers. To capture this, assume that buyers only know the distribution of used car quality and that sellers know the quality with certainty. As soon as a buyer is offered a car, he should infer that the quality is below average, otherwise it would not have been offered in the market. Thus, it is only the cars with below the average quality ("lemons") that is offered in the market, and the price is updated to reflect this fact. Repeating this story, Akerlof came to the strong conclusion that asymmetric information between seller and buyer, creating an effect also known as adverse selection, may cause markets to shrink severely.[1]

Adverse selection obviously imposes a serious problem on the sellers with good cars. However, the nature of this problem is very general, and there exists several classic articles that explore the implications of adverse selection in different situations. Looking at

[1]The reason for why markets do not break down entirely, is that there will always be someone (with a fairly good car) who derives a very low utility from owning the car—and therefore finds it optimal to incur the cost of selling it below the price that would have prevailed with symmetric information.

workers trying to signal their abilities to the jobmarket as an example, Spence (1974) derives conditions that must be satisfied in order to make high-quality workers able to send a trustworthy signal of their quality to the jobmarket. In a related study, Rothschild and Stiglitz (1976) look at how contracts can be designed to have insurance clients voluntarily and truthfully reveal their "quality". Stiglitz and Weiss (1981) show how adverse selection may cause credit to be rationed, and Myers and Majluf (1984) show how a company may forgo a project even if it has positive net present value.

The "underinvestment" result in Myers and Majluf (1984) rests on the same intuition that explains the used car discount. If company insiders know more about the true value of the firm than outsiders, then insiders will be reluctant to issue risky securities when the firm is undervalued. For the same reason, they will have an incentive to issue securities when the firm is overvalued. The security market is aware of these incentives, and will regard an issue as a signal of overvaluation. For the same reason as good cars drop out of the market, leaving relatively more lemons for sale—the high-value firms in Myers and Majluf's model do not issue securities.

Capital market participants react to security issue announcements by revaluing the issuer's stock price. This revaluation depends in part on the market's perception of the issuing firm's objectives and in part on the nature of the information asymmetry between investors and the firm concerning the true value of its securities. Holding fixed the perception of issuing firm's objectives, the arguments in Myers and Majluf (1984) imply that the stock price should drop on the announcement of an equity issue. Based on U.S. data, Asquith and Mullins (1986), Masulis and Korwar (1986), and Eckbo and Masulis (1992) find that the average abnormal (market model) return to firm commitment underwritten seasoned equity offerings (SEOs) over the two-day period ending with the first Wall Street Journal announcement of the issue is about −3%, a value-reduction equal to approximately 20% of the proceeds of the average issue.[2] However, Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) report that common stock returns of industrial firms making SEOs tend to substantially underperform those of a control group of non-issuing firms over the five-year period following the offering date. More specifically, Loughran and Ritter find that the average five-year stock return following SEOs is 60 percentage points below that of non-issuing firms of the same size. Their evidence suggests that the market reaction to

[2]See Eckbo and Masulis (1995) for a review. International evidence is more mixed. The announcement of equity offerings often produce a positive announcement period abnormal return. However, the international evidence is frequently based on the announcement of rights offerings, and Eckbo and Masulis (1992) provide an equilibrium argument (based on adverse selection) that is consistent with a positive announcement period effect for rights offerings.

SEO announcements is informationally inefficient: "... if the market fully reacted to the information implied by an equity issue announcement, the average announcement effect would be −33%, not −3%." (Loughran and Ritter, 1995, p.48).[3] If true, this "new issues puzzle" constitutes a serious challenge to the presumption of rational pricing in security markets.

The main goal of this study is to explore whether or not short- and long-horizon market reactions to seasoned security issue announcements are consistent with rational pricing in security markets.

1.2 Organization and main results

Chapter 2 reviews some of the studies that have looked at long-horizon abnormal stock returns. The main conclusion from the review is that the techniqe of measuring long-horizon abnormal returns using a control group of non-issuing firms, matched on for example size, suffers from serious methodological deficiencies.

The extant literature on long-run performance following corporate events assumes that equilibrium expected returns are constant over the portfolio holding period. However, as surveyed by Ferson (1995), and consistent with dynamic asset pricing models, there is empirical evidence that changes in security risk levels and factor risk premiums (and therefore expected returns) have predictable components related to publicly available information on economic fundamentals. For example, if current corporate yield spreads indicate that a certain stock will have a relatively high expected return over the next period, failure to condition on this information will lead the econometrician to falsely identify "abnormal performance". Chapter 3 provides a discussion of methodological issues, and show how to avoid confusing true abnormal performance with the effect of time-varying expected returns.

Chapter 4 employs some of the methods discussed in chapter 3 on a large number of seasoned security offerings on the New York Stock Exchange (NYSE) and the American Stock Exchange (Amex).[4] As noted in the previous section, Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) find that firms making seasoned equity offerings systematically underperform a control group of non-issuing firms. This finding is part of what has become known as the "new issues puzzle". In addition to the use of time-varying expected return benchmarks in the estimation of long-run performance, chapter 4 provides

[3] If correct, this argument implies that equity issues are characterized by severe adverse selection, requiring a value-reduction of the average issuer by an amount more than twice the size of the equity issue!

[4] Chapter 4 is joint work with B. Espen Eckbo and Ronald W. Masulis.

two additional contributions to the debate on this puzzle. First, contrary to the literature, our long-run performance estimates are linked to the initial issue announcement effect. As a result, it is possible to make in-sample inferences concerning whether the −3% market reaction to equity offerings represents an unbiased estimate of the future. Second, the chapter examines a range of issue and issuer characteristics not presented in earlier papers, including convertible and straight debt in addition to equity, rights versus underwriting as the flotation method, and whether the issuer is an industrial firm or a public utility. Replication of the conventional matching procedures yields similar evidence of long term stock return underperformance. However, this evidence of underperformance disappears when employing time-varying multifactor expected return benchmarks. This finding is robust across type of security (equity or debt), and across issuer types (industrial firms and public utilities). The main finding is that it cannot be rejected that the −3% average equity issue announcement effect is unbiased and consistent with informationally efficient markets.

Eckbo and Masulis (1992) develop a model framework for firm's equity flotation method choice under adverse selection which captures a number of empirical regularities. Chapter 5 formalizes and extends the Eckbo-Masulis model to include private placements as an alternative to rights and underwritten offerings, and to allow the possibility of issue rejection by either the private placement investor or the underwriter. The model is specifically designed to explain the choice among private placements, standby rights offerings, and uninsured rights offerings—which represent the complete range of flotation methods used on the Oslo Stock Exchange, the empirical laboratory for the model. In the model, private investors and investment banks (underwriters) perform a private inspection of the companies that seek project financing. The knowledge about firm quality acquired during the inspection allows a credible quality certification of high-value firms. Thus, these firms can raise funds at more favorable conditions than would have been possible without the inspection. Firms that are rejected by a private investor, may withdraw the issue or use uninsured rights or rights with standby underwriting. Since the rejection by the private investor is based on an informative inspection, the expected issuer-value for private placements exceeds that for uninsured rights and standby rights. An event study of the market reactions to announcements of private placements, standby rights, and uninsured rights largely produces results that are consistent with the model. As in and Bøhren, Eckbo and Michalsen (1997) announcing a standby rights offering gives a negative market reaction, while announcements of uninsured rights are regarded as good news. Consistent with the findings of Wruck (1989) and Hertzel and Smith (1993) on U.S. private equity offerings, and Kang and Stulz (1996) on Japanese offerings, the announcement of private placements on the OSE produces positive abnormal announcement period returns.

While chapter 4 looks at insufficient control for risk as a potential explanation for the "new issues puzzle", chapter 6 explores the timing hypothesis and the underreaction hypothesis as alternative explanations. The timing-hypothesis builds on the notion that investors are are overly optimistic about the prospects of issuing firms, and as a consequence do not fully incorporate into prices managers incentive to time an equity issue. This results in initial overpricing of issuing firms and a subsequent long-run underperformance when investors correct this initial mispricing over time. The overconfidence hypothesis of Daniel, Hirshleifer and Subrahmanyam (1998) is closely related, but is derived in a formal model and carries some explicit empirical predictions. The overconfidence hypothesis is based on the assumption that investors are overconfident about the precision of their private information, but not about the precision of public information. Overweighting private information relative to public information causes underreaction to new *public* information. Thus, the theory predicts that discretionary corporate events (such as equity issues) associated with abnormal announcement period returns, on average should be followed by long-run abnormal performance of the same sign as the average announcement period abnormal return. The empirical predictions of these hypotheses are tested using data on Norwegian private and public equity offerings. Neither the timing-hypothesis nor the underreaction hypothesis receive convincing support by the reported evidence.

Chapter 2

Long-run stock return performance: A review

This chapter reviews some of the studies that have looked at long-horizon abnormal stock returns. In most cases, the horizon is defined to be from three to five years. When estimating long-run abnormal stock returns, one needs to make three major methodological choices: How to model stocks' expected return, how to cumulate returns over the long-run period, and how to obtain a test statistic that is unbiased and that exhibits good power properties. Section 2.1 looks at the different choices that has been made in the literature. Section 2.2 reviews and discusses the empirical evidence on long-run abnormal performance for a wide range of corporate events—however, with emphasis on long-run abnormal stock returns after security offerings. Section 2.3 concludes the chapter.

2.1 Methodology

This section reviews the wide range of different methods used to estimate and statistically evaluate abnormal long-run stock returns. The difference between the methods lies in how expected returns are modeled, in how the returns are cumulated over time, and finally in what test statistic that is used to evaluate the statistical significance of the abnormal performance.

2.1.1 Expected return

Let abnormal performance over period t be defined as

$$AR_{it} = R_{it} - E(R_{it}), \tag{2.1}$$

where R_{it} is the realized return on asset i during period t, and $E(R_{it})$ is the expected return on asset i formed at the beginning of period t. In order to measure AR_{it} empirically, one needs an empirical model of $E(R_{it})$. Table 2.1 lists four classes of expected return measures that subsume the majority of specific measures (benchmarks) used in the literature.

Table 2.1
Empirical measures of expected stock return

(a) Market index	$E(R_{it}) \equiv R_{mt}$ where R_{mt} is a market index. Among the indices used are the equally and value weighted NYSE/Amex/Nasdaq CRSP indices.
(b) Reference portfolio	$E(R_{it}) \equiv R_{pt}$ where R_{pt} is a reference portfolio chosen based on some characteristic of the event firm. A widely used procedure is to partition the universe of listed firms into groups based on market capitalization (size) and book-to-market ratio. With, say 10 size groups and 5 book-to-market group, one gets 50 reference portfolios. The reference portfolio containing a given event firm is then used to measure the expected return.
(c) Matching firm	$E(R_{it}) \equiv R_{jt}$ where R_{jt} is the return for a matching firm. The matching firm is chosen based on one or several firm characteristics such as market capitalization, book-to-market ratio, industry, pre-event stock return performance, stock price, and dividend yield. The firm in the set of eligible matching firms closest to the event-firm is chosen as the matching firm.
(d) Factor model	$E(R_{it})$ is measured using an empirical asset pricing model. Models that have been used include the CAPM, the Fama and French (1993) three-factor model, and various multifactor models.

The benchmarks in table 2.1 are different when it comes to explicitly accounting for the idiosyncratic risk of the event firm's stock return. When using a market index, a benchmark portfolio, or a matching firm as the return benchmark, one implicitly assumes that event firms on average have the same risk as the market, as the group of firms in the benchmark portfolio, or the class of eligible matching firms, respectively. A factor model, on the other hand, makes an explicit attempt to control for the degree of risk through event-firm specific factor betas.

2.1.2 Long-run abnormal returns

To assess the average long-run abnormal performance in a sample of event firms, the literature has used cumulative abnormal return (CAR), average excess buy-and-hold return (BHAR), and various versions of Jensen's alpha (Jensen, 1968). Let τ_i and T_i denote the starting and ending months of the return cumulation (holding period), respectively, and let T be the number of months between τ_i and T_i (including both the start and end month). The cumulative abnormal return for event-firm i is:

$$CAR_{iT} = \sum_{t=\tau_i}^{T_i} AR_{it}. \tag{2.2}$$

With N event-firms in sample, the average cumulative abnormal return is:

$$CAR_T = \frac{1}{N}\sum_{i}^{N} CAR_{iT}, \tag{2.3}$$

Abnormal buy-and-hold return for event firm i is defined as:

$$BHAR_{iT} = \prod_{t=\tau_i}^{T_i}(1 + R_{it}) - \prod_{t=\tau_i}^{T_i}(1 + E(R_{it})). \tag{2.4}$$

The average abnormal buy-and-hold return is then:

$$BHAR_T = \frac{1}{N}\sum_{i=1}^{N} BHAR_{iT}. \tag{2.5}$$

A factor model procedure assumes that the expected returns are generated by a set of K risk factors. Following the idea originally developed by Jensen (1968), the abnormal return on a portfolio p is estimated by regressing the returns on portfolio p on a constant α_p and the K risk factors:

$$R_{pt} - R_{ft} = \alpha_p + \sum_{k=1}^{K} \beta_{pk}\Lambda_{pt} + \epsilon_{pt} \tag{2.6}$$

where R_{pt} is the return on a portfolio constructed using the sample of event-firms, R_{ft} is the riskfree rate of return, β_{pk} is the event-firm portfolio's sensitivity to risk factor k, and Λ_{pt} the excess return on risk factor k. The estimate for the constant term is the "Jensen's alpha" and represents the average monthly abnormal return for the event-firms over the estimation period. An alternative factor-model approach is also commonly used. Instead of regressing the factors on a portfolio of event-firms, the factors are regressed on the excess return of individual event-firms. This gives N firm specific alphas over which we can compute the average abnormal performance. This approach is the same as computing CAR_T using a factor model to measure expected returns. Frequently used factor models are the CAPM:

$$R_{pt} - R_{ft} = \alpha_p + \beta_{p1} \left[R_{mt} - R_{ft} \right] + \epsilon_{pt}, \tag{2.7}$$

and the Fama and French (1993) three-factor model:

$$R_{pt} - R_{ft} = \alpha_p + \beta_{p1} \left[R_{mt} - R_{ft} \right] + \beta_{p2} SMB_t + \beta_{p3} HML_t + \epsilon_{pt}, \tag{2.8}$$

where R_{mt} is the return on the market portfolio, SMB_t is the difference in return between small and large firms, and HML_t is the difference in return between high and low book-to-market firms.

Since abnormal long-term stock return is the difference between realized and expected returns, the natural choice of a test statistic is the conventional t-statistic. In most applications this statistic is computed using the cross-sectional sample standard deviation. This statistic is applicable for CAR_T and $BHAR_T$. For CAR_T, an alternative test statistic may be formed by computing a t-statistic based on the time-series standard deviation of abnormal returns from a pre-event testing period. This approach is followed by Kothari and Warner (1997).

2.1.3 Evidence from simulations

Since all the expected return benchmarks can be paired with any of the methods to compute the long-run abnormal returns. the array of abnormal long-run performance measures is quite large. Several recent papers use simulations in order to explore the the statistical properties of the different measures. In sum, the studies of Kothari and Warner (1997), Barber and Lyon (1997), Lyon, Barber and Tsai (1999), and Cowan and Sergeant (1997) conclude that most measures of abnormal long-run stock returns are biased or have low power.

Measurement biases

The sources of measurement biases vary between models, but the following five sources seem to be important:

New listing/survival ship bias When the long-run stock returns of an event firm are compared with the returns on an index or benchmark portfolio that includes new firms as they are listed, one effectively compares a firm which typically has a long post-event history of returns with firms that begin trading after the event date. Based on the finding that newly listed firms (IPOs) underperform relative to the average firm (Ritter, 1991), Barber and Lyon (1997) conjecture that the new listing bias is positive. Moreover, if the event-firm selection procedure requires event-firms to have pre-event return history, the new listing bias is exacerbated. Simulation studies confirm the conjecture that the new listing bias is positive. Barber and Lyon (1996) find that adding a 24 month pre-event return requirement to a simulation of three-year $BHAR_T$ using the NYSE/Amex/Nasdaq equally weighted index as the return benchmark, increases the average $BHAR_T$ from -0.1% without the pre-event return requirement to 1.18% with this requirement. The new listing bias will only affect long-run performance measures that compute expected returns using an index or benchmark portfolio. Thus, expected returns computed using matching-firms or a factor model with macro economic factors will not contain a new listing bias.

Rebalancing bias The monthly rebalancing implicitly assumed when compounding the return on an equally weighted index or benchmark portfolio will create a negative bias in long-run performance measures. The source of this bias is the negative autocorrelation in periodic returns, created by bid-ask jumps and non-synchronous trading.[1] In order to maintain equal portfolio weights, one have to sell stocks that have high prior-month returns and buy stocks that have low prior month returns. Due to negatively autocorrelated returns, this means purchasing firms that subsequently perform well and selling firms that subsequently perform poorly. Consequently, the monthly returns on equally weighted indices or benchmark portfolios are inflated relative to the returns of event firms, resulting in a negative rebalancing bias. The bias is more pronounced for daily returns than for monthly returns, and the bias becomes more severe when monthly returns are compounded (as with $BHAR_T$) rather than added (as with CAR_T). Canina et al. (1998) conduct an interesting simulation that document this bias. They compare the returns on the CRSP equally weight-

[1]The effect of bid ask-jumps on returns are analyzed in Blume and Stambaugh (1983), while Roll (1983) analyzes the effect of non-synchronous trading.

ed monthly index (the equally weighted average of monthly individual stock buy-and-hold returns) with a monthly index created by compounding the returns of the equally weighted CRSP daily index. Over the 1964–1993 period, the index using compounded daily returns had an average monthly return of 1.72% compared with a monthly return of 1.23% for the CRSP equally weighted monthly index. On average this amounts to a 6.04% difference in annual returns.

Skewness bias The skewness bias arises when the return of a single event-firm is compared to the return on an index or a benchmark portfolio. The bias is caused by the fact that single-firm stock returns are positively skewed, while the return on a portfolio of several firms are not. Using a χ^2 distribution with one degree of freedom as an example distribution, Barber and Lyon (1997) show that a positively skewed distribution tends to create a negative bias in long-run abnormal returns. Like the rebalancing bias, the skewness bias is inflated by compounding, hence, it is a more serious problem for $BHAR_T$ than for CAR_T. Moreover, the sample variance underestimates the true variance, such that t-statistics become too large—causing overrejection of the null hypothesis of no abnormal long-run performance.

Cross-sectional dependence bias When using conventional t-statistics to evaluate the statistical significance of abnormal long-term returns, one assumes that abnormal returns are independently distributed. This assumption does not hold in most studies since corporate events, for example equity issues or stock splits, tend to cluster in certain time periods, and it is common to see a corporate event repeatedly during the period of return calculation. Lyon, Barber and Tsai (1999) and Brav (1998) study the importance of this cross-sectional dependence bias. For example, Lyon, Barber and Tsai (1999) find pervasive test misspecifications in non-random samples using both holding buy-and-hold returns and calendar time portfolios of event firms.

Bad model bias The bad model bias occur when event-firms have systematically different return generating characteristics than what is captured by the expected return model. Kothari and Warner (1997) focus on this type of bias. Based on simulation evidence using CAR_T and several different benchmark models,[2] they find that all models are severely

[2]Their study includes the following models of abnormal returns: a market adjusted model ($R_{it} - R_{mt}$), a market model ($R_{it} - \alpha_i - \beta_i R_{mt}$), the CAPM ($R_{it} - R_{ft} - \beta_i[R_{mt} - R_{ft}]$), and the Fama and French (1993) three factor model ($R_{it} - R_{ft} - \beta_{i1}[R_{mt} - R_{ft}] - \beta_{i2}HML_t - \beta_{i3}SMB_t$)). Where R_{it} it the period t return of event-firm i, R_{ft} is the riskfree rate of return, R_{mt} is the return on the market portfolio, SMB_t is the

misspecified, and that the degree of misspecification is not sensitive to the choice of model. Using a 36-month return cumulation period, they find rejection frequencies ranging from 18.4% to 34.8% when the nominal rejection level is 5%. That is, the models show abnormal performance too often. Other authors have found that different benchmark models can lead to different conclusions about the long-run performance in specific samples. In a study of postmerger performance of acquiring firms, using four different expected return benchmarks, Franks, Harris and Titman (1991) find significantly negative, non-significant, and significantly positive abnormal 36-month returns—using the same set of event-firms.

Table 2.2 contains some of the simulation based evidence reported by Kothari and Warner (1997) and Barber and Lyon (1997). The table reveals that the results for $BHAR_T$ reported by Barber and Lyon (1997) and Kothari and Warner (1997) are inconsistent. While Barber and Lyon find that $BHAR_T$ tend to give negatively biased test-statistics, Kothari and Warner find positively biased test-statistics for $BHAR_T$. However, the buy-and-hold returns computed in these two papers are not directly comparable. Barber and Lyon use the model in (2.5) while Kothari and Warner use the model:

$$\prod_{t=\tau_i}^{T_i} (1 + AR_{it}) - 1.$$

Nevertheless, Barber and Lyon (1996) find that the simulation results are sensitive to the period from which event-months are drawn, whether or not Nasdaq firms are included in the simulations (dropping Nasdaq firms tends to reduce the negative bias in $BHAR_T$), and whether or not sample selection and abnormal return computation require pre-event return history for either event-firms or benchmark firms (the new listing bias).

Another interesting result to note from table 2.2 is that using the equally weighted market index or reference portfolios as expected return benchmarks gives highly misspecified tests for abnormal long-run performance. The reason is that these benchmarks suffer from the new listing bias, the rebalancing bias, the skewness bias, and the cross-sectional dependence bias—and most likely also from the bad model bias. The CAPM and the Fama and French (1993) three-factor model seem to do a little better than the equally weighted market index or reference portfolios, but these factor models also overreject the null hypothesis of no abnormal long-run performance estimates. The only test statistic that performs reasonably well is the buy-and-hold return using a size-/and book-to-market matched control firm as the expected return benchmark. At a 5% theoretical level of significance, Barber and

difference in return between small and large firms, and HML_t is the difference in return between high and low book-to-market firms.

Table 2.2
Empirical rejection levels of three-year abnormal returns in random samples under the null of no abnormal performance

Barber and Lyon (1997) draw 1000 random samples of 200 event months without replacements. Each event-month is associated with a firm drawn at random with replacement. Kothari and Warner (1997) draw firms first, then they draw a random event month. Kothari and Warner use 250 samples of 200 firms. Denote the number of simulations by S, then a well specified one-sided test with a nominal level of significance given by α should reject the null in $S\alpha$ samples. Column "5.0" contains the empirical rejection rates of the hypothesis that abnormal returns are negative, when the nominal rejection level is set to 5%. Column "95.0" contains the empirical rejection rates of the hypothesis that abnormal returns are positive, when the nominal rejection level is set to 5%.

	Barber and Lyon (1997)[a]			Kothari and Warner (1997[b])		
	5.0	95.0	Mean	5.0	95.0	Mean
(a) CAR_T						
Equally-weighted market index	1.4	15.8	3.5	8.4	35.2	3.7
Fifty size/book-to-market portfolios	4.6	6.9	0.7	–	–	–
Size-matched control firm	6.0	6.1	−0.6	–	–	–
Size/book-to-market matched control firm	6.0	4.4	−0.6	–	–	–
CAPM	–	–	–	2.8	28.4	3.3
Fama and French (1993) three-factor model	6.9	4.1	−0.9	6.4	34.0	3.9
(b) $BHAR_T$						
Equally-weighted market index	10.0	2.4	−0.1	0.0	91.2	27.8
Fifty size/book-to-market portfolios	20.1	0.5	−5.2	–	–	
Size-matched control firm	5.4	5.7	−0.2	–	–	
Size/book-to-market matched control firm	5.0	5.1	−0.9	–	–	
CAPM	–	–	–	1.2	30.8	5.48
Fama and French (1993) three-factor model	–	–	–	0.8	34.0	6.13

[a]Barber and Lyon (1997) use the cross-sectional sample standard deviation of abnormal returns to compute the t-statistics for CAR_T and $BHAR_T$.

[b]Kothari and Warner (1997) compute the t-statistic for CAR_T based on the time-series standard deviation of abnormal returns from a 24-month pre-event testing period. They define buy-and hold returns in a different way than the one normally used. Instead of cumulating returns and then take the difference, they cumulate the difference:

$$\prod_{t=\tau_i}^{T_i} (1 + AR_{it}) - 1.$$

Lyon (1997) document empirical rejection levels of 5.0% and 5.1% for the null of negative abnormal performance and positive abnormal performance respectively. However, the usefulness of these simulations is highly questionable. Remember that the event-firm sample is drawn at random. Thus, even if size and book-to-market ratio are totaly unrelated to stock returns, matching on these characteristics should produce a sample with the same mean buy-and-hold return—producing a zero bias. The only useful information that this type of simulation might provide is on the variability of the mean bias when the simulation is repeated several times. Interestingly, simulation error (variability in the simulations) seems to be non-neglectable. Lyon, Barber and Tsai (1999) use the same simulation setup and find empirical rejection rates of 5.5% and 3.0% respectively.

Alternative test statistics

Lyon, Barber and Tsai (1999) identify two approaches to long-run performance measurement that yield well specified test statistics in randomly drawn samples of event-firms. The first approach is based on buy-and-hold returns together with reference portfolios carefully constructed to avoid the rebalancing bias and new-listing bias. The universe of NYSE/Amex/Nasdaq listed firms are partitioned into 70 reference-containers based on firm size and book-to-market ratio.[3] The buy-and-hold return on the reference portfolio for event-firm i with event-month s is calculated as:

$$\frac{1}{n_s}\sum_{j=1}^{n_s}\left[\prod_{t=s}^{s+\tau}(1+R_{jt})\right]-1,$$

where n_s is the number of firms in event-firm i's reference-container in month s, τ is the number of months in the cumulation period, and R_{jt} is the monthly return on reference firm j. Since no new firms are added to the reference portfolio after month s, the new-listing bias is avoided. Moreover, since the buy-and-hold returns are first computed for each reference firm and then averaged, the rebalancing bias is avoided. In the second approach, a calendar time portfolio of event-firms is compared to the expected return benchmark.[4] If a factor model is used as the benchmark, inference is based on the t-statistic of the added constant term (the Jensen's alpha). The advantage of the latter approach is that a factor model easily can be set up to avoid the new-listing and rebalancing bias. Moreover, the skewness-bias is eliminated since event-firm returns are measured as a time-series of portfolio returns.

[3]To be more precise, they use all firms on CRSP with share codes 10 and 11.

[4]This approach was first employed by Loughran and Ritter (1995), using the Fama and French (1993) three-factor model as the benchmark.

Returning to the reference-portfolio approach, the skewness-bias still remains, and will cause a bias in test statistics if not taken care of. Lyon, Barber and Tsai (1999) explore several methods that handle the inherent skewness in the abnormal performance measure. They identify two methods that both yield tests that are well specified in random samples and have good power properties:[5] a bootstrapped skewness adjusted t-statistic and bootstrapped empirical p-values.[6] Of these methods, the latter is the most widely used in the existing long-run performance literature. The idea is to create the empirical counterpart of the distribution of event-firm abnormal returns. The steps needed to create this empirical distribution, is first to draw a random "pseudo-sample" of event-firms: For each firm in the original event-sample another firm from the same reference-container is drawn at random. This gives one "pseudo-sample", for which abnormal returns are computed using matching firms or reference-portfolios. The average abnormal return for the "pseudo-sample" gives one point on the empirical distribution of abnormal returns. The process is repeated S times to map out the whole empirical distribution of abnormal returns. Denote the average abnormal return in the original event-sample by $\overline{ar}$. The null hypothesis of no abnormal performance against the alternative of negative abnormal performance is rejected at the α percent level if the number of abnormal returns in the empirical distribution that is less than $\overline{ar}$ does not exceed αS.

Test statistics in non-random samples

The typical long-run performance study selects event-firms based on the occurrence of a certain corporate event. Consequently, event-firms share an important common characteristic. If this characteristic is correlated with expected returns, either directly or through other firm characteristics that are correlated with expected returns, simulations that draw event-samples at random may understate test misspecification. Lyon, Barber and Tsai (1999) recognize this problem and report the behavior of test-statistics in non-random samples. Table 2.3 provides some of the evidence reported by Lyon, Barber and Tsai (1999) on biases in non-random samples.

The evidence in table 2.3 is rather discouraging. It shows that test-statistics that were well specified in random samples, turn out to be severely misspecified when used to evaluate long-run performance in samples that are not drawn at random. This raises a serious concern

[5] The power of a test is the ability to reject an incorrect null hypothesis. The power is evaluated by adding a known level of abnormal return to the calculated abnormal return in the randomly drawn samples.

[6] Bootstrapped empirical p-values were first employed in the long-run performance literature by Ikenberry, Lakonishok and Vermaelen (1995)

Table 2.3
Empirical rejection levels of five-year buy-and-hold abnormal returns in non-random samples under the null of no abnormal performance using reference portfolios constructed to avoid new listing and rebalancing bias (Lyon, Barber and Tsai (1999))

Lyon, Barber and Tsai (1999) draw 1000 non-random samples of 200 event months without replacements. Each event-month is associated with a firm drawn at random with replacement. Large and small firms are firms from the largest and smallest market capitalization deciles. Firms with high and low book-to-market ratios are firms from the highest and lowest book-to-market ratio deciles. Firms with high and low pre-event six month returns are firms from the highest and lowest pre-event six month return deciles. Industry clustered samples are drawn by making sure that the 200 firms in each of the simulation is drawn from the same randomly drawn two-digit SIC code. Calendar clustered samples are drawn by making sure that the 200 firms in each of the simulation has the same randomly drawn event-month. Overlapping returns are created by a first draw of 100 randomly selected event firm-months, then for each of these 100 firms, a new event-month is selected within τ months of the first event-month, where τ is the long-run horizon. Denote the number of simulations by S, then a well specified one-sided test with a nominal level of significance given by α should reject the null in $S\alpha$ samples. Column "2.5" contains the empirical rejection rates of the hypothesis that abnormal returns are negative, when the nominal rejection level is set to 2.5%. Column "97.5" contains the empirical rejection rates of the hypothesis that abnormal returns are positive, when the nominal rejection level is set to 2.5%. The calendar time portfolio of event-firms used to compute Jensen's alpha is equally weighted. Results are similar with value weighting (See Lyon, Barber and Tsai, 1999).

	Bootstrapped skewness-adjusted t-statistic		Bootstrapped empirical p-values		Fama and French (1993) Jensen's alpha	
Non-random sample	2.5	97.5	2.5	97.5	2.5	97.5
Large firms	2.3	2.6	3.6	2.1	0.7	2.9
Small firms	2.8	3.3	3.4	3.5	0.2	3.1
Firms with high book-to-market ratio	2.5	2.7	2.9	2.6	0.0	15.0
Firms with Low book-to-market ratio	4.0	1.2	6.7	1.3	16.8	0.0
Firms with high six-month pre-event returns	9.1	0.4	14.8	0.3	1.0	1.0
Firms with low six-month pre-event returns	1.6	3.0	3.4	3.8	1.2	0.2
Industry clustering	10.5	15.9	14.2	12.9	3.9	7.5
Calendar clustering	3.1	2.5	2.8	2.8	3.1	7.8
Overlapping returns	6.1	6.7	5.1	4.8	0.2	2.2

about the validity of the evidence we have so far on long-run abnormal performance after corporate events. The problem may even be exacerbated by the fact that firms selected based on certain corporate events are likely to be similar on several of the characteristics listed in table 2.3. For example, we know that firms performing a seasoned equity offering (SEO) typically have a significant pre-announcement run-up in stock prices, that SEOs tend to cluster in time ("hot issue markets"), and that several SEOs by the same firm during a relatively short period are not unusual.

One potential explanation for the misspecifications in table 2.3 is that self-selection by event-firms (most corporate events studied in the literature are voluntary actions taken by the company) results in cross-sectional dependence in abnormal returns. Brav (1998) suggests a testing methodology that is tailored to handle this potentially important problem. Moreover, it is robust with respect to the non-normalities observed in abnormal return estimates. The suggested method goes under the general heading of Markov Chain Monte Carlo techniques. The basic idea is to take a Bayesian approach by specifying prior distributions for all random variables in the model, then use an iterative simulation procedure to arrive at the posterior distribution, *given the data*, for the variables that we are interested in. In short, the methodology involves the following steps: First, a null-model for asset returns is specified. In general, any type of model is applicable—and Brav (1998) uses the Fama and French (1993) three-factor model. Since the main goal is to take care of potential cross-sectional correlation, all the individual firm Fama-French regressions are stacked in a Seemingly Unrelated Regression setup:

$$R = FB + E,$$

where R is the vector of stacked event-firm returns, F is a block diagonal matrix of the factors in the Fama and French (1993) three-factor model, B is the corresponding matrix of factor loadings, and E is a vector of residuals. Denote the variance-covariance matrix of E by Σ. Second, the posterior distribution of $\{B, \Sigma\}$ given $\{R,F\}$ is constructed using a method of iterative simulations known as the Gibbs sampler. When convergence is attained, the Gibbs sampler deliver draws of B and Σ from the posterior distribution. Third, using S draws on B and Σ and the actual factors F, one can obtain a set $\{R_1, R_2, \ldots, R_S\}$ of simulated event-firm returns. Finally, the simulated event-firm returns are used to compute the distribution of average abnormal returns. As opposed to the empirical p-values suggested by Ikenberry, Lakonishok and Vermaelen (1995), this distribution is generated under the restriction that event-firm abnormal returns are cross sectionally related. The hypothesis of no abnormal performance is rejected at the α level of significance if the average abnormal return of the original event-firms sample is below the $\alpha/2$-percentile or above $(1 - \alpha/2)$-percentile.[7]

In a sample of initial public offerings (IPOs) Brav (1998) finds that accounting for the cross-sectional correlation tends to stretch out the empirical distribution of abnormal returns. In other words, tests that ignores the cross-sectional correlation will reject the null hypothesis of no abnormal return too often. This result is consistent with the findings reported in table 2.3. Although the testing methodology advanced by Brav (1998) suggests a solution to the cross-sectional dependence bias, the size (degree of misspecification) and power properties of the test-statistic is not known. Thus, it is not possible to say how good the test is relative to the other tests discussed in this chapter.

2.2 Empirical evidence

Since Ritter's (1991) study of the long-run stock returns after initial public offerings (I-POs), there has been a large number of other studies looking at the long-run stock return performance of listed firms after specific corporate events such as seasoned equity offerings (SEOs), debt offerings, share repurchases, spin-offs, proxy contests, exchange listings, analyst recommendations, and earnings announcements. The majority of these studies find evidence of either abnormally high or abnormally low long-run stock returns. However, the long-run performance is in general sensitive to the methodology used. The next sections review some of these studies in more detail.

2.2.1 Security offerings

Long-run performance has been studied for unlisted firms that go public, for listed firms that season equity, and for listed firms that issue straight or convertible debt.

Initial public offerings

Ritter (1991) was the first large sample study of long-run post event stock returns after security offerings. In a sample of 1526 IPOs on New York Stock Exchange (NYSE), American Stock Exchange (AMEX), and Nasdaq over the period 1975–1984 he found a remarkable underperformance over the three years following the IPO. Using a size/industry matched company as the benchmark, the average abnormal buy-and-hold return is −27.4% over the 10 sample years. However, it is the bad performance of IPOs during the period 1980–1984

[7] The distribution is by construction centered at zero.

that drives the negative average abnormal buy-and-hold returns. Almost 73% of the firms in the sample went public during this "hot-issue" period. The underperformance also varies considerably between industries. Some industries, like oil and gas, underperform substantially compared to their benchmark firms, while other industries, such as financial institutions, show large overperformance. Ritter views this evidence as support for the hypothesis that firms are timing IPOs to peaks in industry specific fads. In other words, to periods where firms in a specific industry are overvalued by the market. However, he recognizes that the results also are consistent with chance—i.e., just bad luck.

Loughran and Ritter (1995) revisit the long-run performance of IPOs using an extended sample of 4,753 NYSE/AMEX/Nasdaq companies going public during 1970 to 1993. Using several of the methods described in the methodology section, they confirm the findings of Ritter (1991). On average IPOs underperform relative to a size matched control firm by −27% over a three-year period after the IPO. Extending the three-year post IPO period to a five-year period, they find that IPOs continue to perform worse than the benchmark for another two years, resulting in an abnormal five-year buy-and-hold return of approximately −51%. Also in this larger sample, the underperformance is concentrated in "hot-issue" periods during the beginning of the 1970's and 1980's.

Motivated by a large number of empirical studies showing that that size and book-to-market ratio are important determinants of stock returns, Loughran and Ritter (1995) run monthly cross-sectional regressions over the period 1973–1992 to control for this. For two of the model specifications used, the average coefficients for 240 cross-sectional regressions are:

$$\begin{aligned} r_{it} &= 1.42 && && -0.49 \times \text{ISS}_{it} && +\epsilon_{it}, \\ r_{it} &= 1.70 && -0.05 \times \ln MV_{it} \quad +0.30 \times \ln(BV/MV)_{it} && -0.38 \times \text{ISS}_{it} && +\epsilon_{it} \end{aligned} \tag{2.9}$$

where $\ln MV$ is the natural logarithm of market capitalization (size), $\ln(BV/MV)$ is the natural logarithm of the book-to-market ratio, and ISS is a dummy variable that takes on the value one if company i has issued equity during the five years prior to t. The authors interpret these regressions as evidence of a size/book-to-market effect in the long-run stock return performance of issuing companies, but that it only accounts for about 22% (the drop in the coefficient on ISS). Loughran and Ritter (1995) interpret their evidence as a reinforcement of the conclusion from Ritter (1991): firms go public when they are substantially overvalued.

Brav and Gompers (1997) perform a more detailed analysis of the underperformance of IPOs based on different firm characteristics, with main focus on whether or not the firm is taken public with the help of a venture capitalist, and using size/book-to-market based benchmarks. Compared to the Loughran and Ritter (1995) sample, Brav and Gomper's sample excludes unit offerings,[8] and is restricted to the period 1975–1992. This gives a total sample of 4,341 IPOs, of which 934 is backed by a venture capitalist. Brav and Gompers (1997) also point out that these IPOs typically are made by small firms with low book-to-market ratio (growth firms). About 43% of the venture-backed and 32% of the non-venture backed companies are found in the intersection between the lowest size decile and the lowest book-to-market quintile.

Using a wide range of different indices as benchmarks, Brav and Gompers (1997) find IPO underperformance in the nonventure-backed IPO sample similar to Ritter (1991) and Loughran and Ritter (1995). However, the venture-backed IPOs does not underperform the benchmarks to the same extent. Using industry portfolios as the benchmark, nonventure-backed IPOs underperform the benchmark by −33.8% while venture-backed IPOs only underperform by −4.4%. Brav and Gompers (1997) also document two other interesting results. First, using a set of 25 size/book-to-market reference portfolios purged of firms conducting IPOs or SEOs, they do not find long-run abnormal performance in neither of the two IPO sub-samples. They interpret this as evidence that the underperformance is not an IPO effect, but a size/book-to-market effect—small firms with low book-to-market ratio tend to underperform relative to market wide or industry wide indices. Second, they show that underperformance is sensitive to the weighting scheme used to aggregate individual firm performance. If value weights are used instead of the equal weights implied by the buy-and-hold abnormal performance used by for example Loughran and Ritter (1995), the abnormal performance becomes much less pronounced. This is, of course, related to the finding that underperformance is concentrated among small firms.

The results from the cross-sectional regressions in Loughran and Ritter (1995) may seem to contradict the results of Brav and Gompers (1997). This is not necessarily true. Consider the regressions reported in table 2.4. The table report results from an experiment similar to the one used by Loughran and Ritter (1995), but with the difference that in some model specifications the issue dummy is replaced by a size/book-to-market dummy.

In table 2.4 each monthly regression uses all firms listed on NYSE, AMEX, or Nasdaq with common stocks that have returns and market values on CRSP, and book values on COMPUSTAT. For months January through June, book-to-market ratios (BV/MV) are

[8] A unit offering is a combined offering containing one share and one warrant. Unit offerings are excluded because only the share trade publicly such that calculating the return to investors becomes difficult. Moreover, unit offerings tend to be issued by very small and risky companies.

Table 2.4

Monthly cross-sectional regressions of stock return on size, book-to-market ratio, new issue dummy and size dummy, 1970–1993

The population is all firms listed with common stocks on NYSE, AMEX, or Nasdaq. For each monthly regression, the sample firms are required to have returns and market values on CRSP, and book values on COMPUSTAT. For months January through June, book-to-market ratios (BV/MV) are from the fiscal year two years back, for July through December book-to-market ratios are from the last fiscal year. Market value of equity (MV) is from the last June before the estimation month. The issue dummy takes on the value of one if a company issued equity during the five year period preceding the estimation month, and zero otherwise. The size/book-to-market dummy takes on a value of one if a company's size is in the first or second size quintile and in the first or second book-to-market quintile, and zero otherwise. The quintiles are created using NYSE breakpoints. The coefficients reported are the average values of all the cross-sectional regressions. Parentheses contain t-values computed using the coefficient time-series and the number of positive coefficients.

	Model specifications				
Independent variables	(1)	(2)	(3)	(4)	(5)
Intercept	1.37 (3.80; 61)	1.52 (4.36; 62)	1.72 (3.69; 59)	1.83 (4.07; 60)	1.88 (4.20; 60)
$\ln(MV)$			−0.07 (−1.24; 48)	−0.10 (−1.76; 47)	−0.10 (−1.82; 48)
$\ln(BV/MV)$			0.47 (5.75; 66)	0.41 (5.42; 65)	0.44 (5.89; 68)
Issue dummy	−0.50 (−4.02; 42)		−0.27 (−2.90; 43)		
Size/book-to-market dummy		−0.65 (−4.84; 34)		−0.25 (−3.02; 40)	−0.20 (−2.40; 40)

computed using book value of equity and market value of equity from the fiscal year two years back, for months July through December book-to-market ratios are computed using information from the last fiscal year. Market value of equity (MV) is from the last June before the estimation month. The issue dummy takes on the value of one if a company issued equity during the five year period preceding the estimation month, and zero otherwise. The IPO data of Ritter (1991) and the SEO data of Eckbo, Masulis and Norli (1999) is used to determine if a company issued equity. The size/book-to-market dummy is created by first assigning all firms in the sample to 25 size/book-to-market portfolios based on the intersection of size and book-to-market quintile portfolios. The size/book-to-market dummy takes on a value of one if a company is in any of the four portfolios created by intersecting the two smallest size portfolio with the two smallest book-to-market portfolios, and zero otherwise.

Model specification (1) and (3) in table 2.4 replicate the cross-sectional regressions reported in equation 2.9 above. As should be expected, the results are close to the ones reported by Loughran and Ritter (1995). However, comparing specification (1) with specification (2) and (3) with (4), we see that the results using the size/book-to-market dummy is indistinguishable from the results using the issue dummy. In order to handle the close correspondence between the issue dummy and the size/book-to-market dummy, model specification (5) excludes all issuers from the sample. The results are virtually unchanged with this modification. Thus, one cannot conclude (as do Loughran and Ritter (1995)) that the issue dummy picks up a separate issue effect. It could be part of a more general size/book-to-market effect.

Explaining IPO underperformance

Ritter (1991) and Loughran and Ritter (1995) argue that firms are timing their decision to go public in order to take advantage of temporary overvaluations. Timing and subsequent underperformance implies that markets are inefficient. Loughran and Ritter (1995) suggest that investors are betting on long-shots, and that they are overestimating the probability of picking a big winner. Brav and Gompers (1997) view the documented underperformance, not as a IPO phenomenon, but as more general size/book-to-market phenomenon. They suggest three possible explanations for this more general view. First, the "bad luck" explanation mentioned in Ritter (1991). Since underperformance is concentrated in certain periods, the underperformance may be due to unexpected shocks hitting small firms in these periods. If this is the case, the cross-sectional dependence in abnormal returns can cause inflated standard test statistics, and consequently supporting the hypothesis of abnormal performance. However, using the cross-sectional-dependence robust Markov Chain Monte Carlo techniques described in section 2.1.3, Brav (1998) cannot support the hypothesis that the observed abnormal returns for IPOs are consistent with the Fama and French (1993) three factor model. Second, underperformance of small, low-book-to-market firms could be explained by investor sentiment. Since large institutional investors tend to avoid investing in very small firms, small firm's equity are primarily held by individuals.[9] If individual investors are less sophisticated than institutional investors, small firms are more easily subjected to fads. Third, individual investors may buy small, low book-to-market firms because they value them as lottery tickets—which are not bought at their expected value, but rather for their huge upside potential.

A few studies have explored other potential explanation for the IPO underperformance. Teoh, Welch and Wong (1998*a*) look at earnings management in the year before and in the year of the IPO. If investors are unable to fully understand the extent to which IPO firms engage in earnings management, reporting high earnings before the IPO will transfer into an increased offering price. One way to inflate earnings is to aggressively report current discretionary accounting accruals. Using a wide range of specifications, they find that aggressive reporting of current accruals is followed by a larger post IPO underperformance. Using cross-sectional regressions, they report the following average coefficients:

$$\begin{aligned} r_{it} = +0.86 &+0.16 \times \ln(BV_{it}/MV_{it}) &+0.07 \times \ln MV_{it} \\ &-0.03 \times \ln MV_{it} \times D_{it} &-0.85 \times \ln DCS_{it} \times IPO_{it} \\ &-0.77 \times \ln DCS_{it} &+X_{it}\beta + \epsilon_{it}, \end{aligned} \quad (2.10)$$

where r_{it} is a monthly return from the 6th to the 17th month after the IPO, DSC is discretionary current accruals, IPO is a dummy variable that takes on the value of one if the relevant financial statement is the one immediately following the IPO and zero otherwise, $D = 1$ if $MV_{it} >$\$100 millions and zero otherwise, and X is other accrual variables with corresponding coefficients β. The other variables are defined above. The regressions show that increasing discretionary current accruals tend to lower the monthly return for all firms. However, the interaction variable $\ln DCS_{it} \times IPO_{it}$ shows that managing earnings before an IPO has an additional strong effect. They view this as evidence that part of the IPO underperformance is explained by investors inability to fully understand managers incentive

[9]The costs of gathering information on small firms cannot be recouped without buying a large proportion of the firm. Institutional investors avoid this for two reasons: First, buying a large fraction of the firm without bidding up the price is difficult. Second, owning more than 5% is avoided by many institutions due to regulatory reasons.

to manage earnings in order to obtain a higher offering price.

Carter, Dark and Singh (1998) look at IPO underperformance conditional on the reputation of the underwriter that assists in taking the firm public. Using three different measures of underwriter reputation, but focusing on the Carter-Manaster index, they find that the better the reputation of the underwriter the less the long-run underperformance.[10]

Teoh, Welch and Wong (1998*a*) and Carter, Dark and Singh (1998) seem to document an IPO effect in their data. However, neither of the papers make appropriate control for the size/book-to-market effect documented in Brav and Gompers (1997) and in table 2.4. We know that most IPOs are concentrated in the lowest size decile and the lowest book-to-market quintile measured using NYSE breakpoints. Thus the IPO dummy variable used by Teoh, Welch and Wong (1998*a*) proxy well as a small-firm dummy variable. Consequently, their results are consistent with the view that discretionary accruals by small firms have an additional negative impact on returns. The Carter, Dark and Singh (1998) result could also be a size effect. They report a correlation between the Carter-Manaster index and the size of the offering of about 0.6. Hence, grouping firms according to this variable and computing group averages should give similar results as grouping on size.

In sum, the current evidence seems to indicate that small firms with low book-to-market ratios underperform a wide range of different indices. Whether the data also shows a separate IPO effect is less clear. If the IPO effect is just a sub-effect of a more general size/book-to-market effect, the IPO evidence discussed in this section cannot support the view that firms go public when they are temporarily overvalued.

Seasoned equity offerings

The remarkable stock return underperformance for IPOs found by Ritter (1991), lead to the question of whether this is a particular IPO phenomenon, or whether it is a more general equity issue phenomenon. Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) explore this question by looking at the long-run stock return performance of companies after SEOs. They find underperformance of SEOs similar to that of IPOs. Using size-matched control firms as the benchmark and a sample of 3,702 SEOs during 1970–1990, Loughran and Ritter find a five-year average underperformance of about 60 percentage points. In a shorter sample period, 1975–1989, excluding all offerings with a secondary component, and using a

[10]Carter and Manaster (1990) develop and index of underwriter reputation based on the relative placements in stock offering "tombstone" advertisements. By comparing the relative positions of investment banks in the advertisements, they place underwriters in one of ten categories, where nine is the most prestigious and zero is the least prestigious.

size/industry matched control firm as the benchmark, Spiess and Affleck-Graves (1995) find an average underperformance of about 42%. As for IPOs, the underperformance of SEOs varies with the year in which the stocks were issued—with small or no underperformance in years with few offerings and larger underperformance in years with many offerings. Spiess and Affleck-Graves (1995) find that all firms, independent of firm characteristics such as age, size, book-to-market ratio, and trading system—seems to underperform relative to similar non-issuing firms. However, Nasdaq listed, young, and small firms with low book-to-market ratio seems to underperform more than other issuers.

Several papers have examined the robustness of the SEO underperformance using other benchmarks. Brav, Geczy and Gompers (1995) use a sample of 3,931 NYSE/AMEX/Nasdaq listed SEOs from 1975–1992. Using buy-and-hold returns and a wide range of different indices as benchmarks they find results similar to Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995). However, when the average buy-and-hold return are computed weighting each issuer by its market capitalization, the average long-run underperformance is significantly reduced. The reason for this finding is that the underperformance is concentrated among small firms. Similar results are found using a portfolio of SEOs and the Fama and French (1993) three-factor model as the benchmark. Regressing the Fama-French factors on the returns on an equally weighted SEO portfolio produces a significantly negative Jensen's alpha. When the portfolio includes SEOs with value weights, the Jensen's alpha is closer to zero and becomes statistically insignificant. However, Jensen's alpha for a portfolio of small issuers are significantly negative regardless of the weighting scheme. In a sample of more than 6,000 SEOs during the period 1961–1993, Mitchell and Stafford (1997) largely confirm the findings of Brav, Geczy and Gompers (1995).

Loughran and Ritter (1999) criticize the use of the Fama-French three-factor model as benchmark when evaluating long-run performance. According to Loughran and Ritter (1999), the reduction in abnormal performance typically obtained with value-weighted event-firm portfolios and the Fama-French model as benchmark, is caused by the inability of this setup to detect abnormal performance when events occur as a result of behavioral timing. Their main arguments are based on the notion that small firms typically are missvalued by more than large firms, that missvaluations are time-varying, and that there are periods where a lot of firms take some action to exploit this (for example, issue equity in periods of overvaluation). If we accept these assumption, it is clear that regressing a value-weighted portfolio of event firms on the Fama-French factors, or any other set of factors for that matter, will have less power to detect the abnormal performance following missvaluation. Moreover, they point out that the factors in the Fama-French model include issuers, and is therefore a "contaminated" benchmark with less ability to detect abnormal performance. The latter is potentially an important concern, however, in practice it turns out not to be important. The studies of Eckbo, Masulis and Norli (1999) and Brav, Geczy and Gompers (1998) show that excluding issuers from the benchmarks makes little difference in the performance of SEOs.

The use of size and book-to-market ratio as matching criteria when selecting a control firm is motivated by empirical evidence showing a relationship between these firm characteristics and stock returns. Jegadeesh (1997) takes this idea further and selects matching firms based on other firm characteristics such as 6-month and 36-month pre-issue returns, stock price, and dividend yield. One of the main findings is that as long as one match on book-to-market ratio, the long-run underperformance of SEOs are relatively insensitive to matching on other firm characteristics. Using a sample of 3,174 NYSE/AMEX/Nasdaq listed SEOs over the period 1977–1994, Jegadeesh (1997) finds long-run buy-and-hold underperformance between -16.5 and -21.7 percent. Jegadeesh also documents that the size effect found by Brav, Geczy and Gompers (1995) disappears when matching firms are selected based on size and book-to-market ratios.

Explaining SEO underperformance

Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) suggest that the long-run underperformance of SEOs can be explained by overoptimistic investors and timing of the SEO to a period where the issuing firm is overvalued. Several papers have explored different aspects of this view. Rangan (1995) and Teoh, Welch and Wong (1998*b*) look at discretionary accruals in the years around the offering. The idea is that if investors are overly optimistic about the prospect of firms that issue equity, the optimism could potentially have been induced by inflated earnings. Both papers find evidence of earnings management prior to the SEO. Teoh, Welch and Wong (1998*b*) find that although cash flows from operations are declining prior to the issue, the reporting of discretionary accruals cause earnings to peak around the issue. Moreover, discretionary accruals prior to the SEO is negatively related to the post-issue long-run stock return performance. The authors view this as evidence in favor of timing and overly optimistic investors. However, the method used is the same as in Teoh, Welch and Wong (1998*a*) (see equation (2.10)), thus the results are subject to the same criticism. That is, to the extent that an SEO dummy in the cross-sectional regressions proxy as a size-dummy, the results are also consistent with the view that all firms that manage their earnings experience subsequent declining stock returns, and that this effect is stronger for small firms.

Cornett, Mehran and Tehranian (1998) employ a more direct test of the timing hypothesis. They study a sample of 150 SEOs by commercial banks over the period 1983–1991. Capital regulations in the banking industry state that a bank is not allowed to have a total capital ratio below a certain level. If the total capital ratio falls below the regulated lower bound, the bank may need to issue new equity in order to meet the requirement. Cornett, Mehran and Tehranian (1998) define an involuntary SEO as an issue performed by a bank with capital ratio close to or below the required minimum ratio. Their sample contains 80 involuntary issues, with the remaining 70 issues defined to be voluntary. If timing is driving the long-run underperformance of SEOs, we should expect to see less or no underperformance for involuntary issues. The results support the timing hypothesis, showing no abnormal three-year post issue stock return performance for the involuntary issues, while the voluntary issues show significant underperformance.

Another direct test of the timing hypothesis is done by Brous, Datar and Kini (1998). They argue that if the long-run underperformance by SEOs are due to timing by managers and the inability of investors to fully understand the implications of managers incentive to manage earning prior to the issue, we should expect to see that investors get disappointed when firms convey their post-issue earnings. That is, post-issue earnings announcement should on average be associated with negative stock price reactions. Using a sample of 1486 SEOs over the period 1977–1990, they study about 21,000 earnings announcement in the five-year period subsequent to the issue. The results show no evidence of abnormal stock price reactions to the earnings announcements.

The major competing hypothesis in explaining the long-run underperformance of SEOs is methodology based. Recall that SEO underperformance measured using a size-matched benchmark or the Fama and French (1993) three-factor model mainly is driven by small firms. Fama and French (1993) document that small firm stock returns are low during the post-1963 period, and Fama (1998) use this to argue that the SEO underperformance is explained by the non-existence of an asset pricing model that properly describe the pricing of small stocks during this period. In particular, the Fama and French (1993) three-factor model overestimate the returns on small firms during this period.

In sum. SEO underperformance measured using size-matching or the Fama and French (1993) three-factor model seem to be concentrated among small firms. However, when a control firm is selected based on both size and book-to-market ratio, the underperformance seems to exist in all size and book-to-market quintiles. As for the causes of the SEO underperformance, the timing hypothesis of Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) has received a lot of attention. However, the evidence is mixed and it would be premature to conclude whether or not timing and investor overoptimism are the real reasons behind the week post SEO performance. On the other hand, there exists no convincing evidence that the performance is just a matter of econometrics. As stated in the introductory chapter, the main purpose of this study is to look further into this as a possible explanation.

Debt offerings

Compared to the number of studies that have looked at long-run stock return performance after equity issues, there are only a few studies that have explored the long-run performance after debt issues. This is probably related to the fact that the performance of debt issuing companies is normal in the post issue period. In a sample of 1,533 straight bond offerings over the period 1971 to 1991, Hansen and Singal (1997) find no evidence of post-issue abnormal five-year stock returns. This result is confirmed by Cheng (1994). Using a sample of 662 bond issues over the period 1977 to 1988, Cheng (1994) finds an abnormal post-issue three-year abnormal return of 3.1 percent. In a sample of 1557 straight debt issues over the period 1975–1989, Spiess and Affleck-Graves (1996) report a five-year post-issue abnormal performance of −7.94 percent (using size/book-to-market matched control firms). This is significantly different from zero using a Wilcoxon signed rank test. In a subsample of 292 straight debt issuers that have not issue debt or equity in the five years before or five years after the debt issue (called an "independent" sample), they find a five-year abnormal stock return of about −21 percent. From these results they conclude that firms issuing bonds also underperform in the five-year after the issue. Given the results reviewed in the methodology section, the −7.94 abnormal performance in the full sample is hardly statistically significant using any methods that takes cross-sectional dependence among issuers into account. Thus, their conclusion must rest on the fact that the independent subsample shows a reasonably large underperformance. However, this subsample is selected using hindsight (no issues in the five years *after* the bond issue), and as such cannot be part of an implementable investment strategy. Thus, choosing an appropriate benchmark, using information that is only available at the time of the issue, is not possible for this subsample. In sum, it seems safe to conclude that there are no long-run abnormal performance after straight bond issues. For convertible bond issues the results are similar to the results after equity issues. In a sample of 672 convertible bond offerings, and using size/book-to-market matched control firms as the benchmark, Spiess and Affleck-Graves (1996) find a five-year abnormal stock return performance of −39 percent.

2.2.2 Other corporate events

An extensive review of the evidence on long-run performance after other corporate events than security issues is beyond the scope of this chapter. However, table 2.5 lists some of the studies in this area.

One thing to keep in mind when looking at the table, is that results tend to be very sensitive to the applied benchmark. Abnormal returns tend to increase when a broadly based market index is used to measure abnormal performance. Lakonishok and Vermaelen (1990) find almost three times larger overperformance in their repurchase tender offer sample when using the value weighted market index as benchmark instead of the size based reference portfolio. Michaely, Thaler and Womack (1995) report about twice as large overperformance for the dividend initiation sample using the equally weighted market index as the benchmark. Finally, Franks, Harris and Titman (1991) report significant positive, non-significant, and significant negative long-run abnormal returns after mergers using different benchmarks[11]

In the two stock repurchase samples studied by Mitchell and Stafford (1997), they find insignificant underperformance when value-weighting. Thus, the small firm effect found in the IPO and SEO samples seems to exist in these samples as well. For the open market share repurchase sample, the size-effect is confirmed by Ikenberry, Lakonishok and Vermaelen (1995) who find no abnormal return for value stocks (high book-to-market) and large overperformance for glamour stocks (an average monthly abnormal return of 0.43 over a 4-year post-event period.)

A concern that applies to all studies that use buy-and-hold returns, is that one will find long-run abnormal performance even if the abnormal performance only is concentrated in, say, the first few months of the holding period.[12] Two examples of this is Ikenberry, Rankine and Stice (1996) and Michaely, Thaler and Womack (1995) who both report a three-year 12.1 percent abnormal performance after stock splits and dividend initiations respectively. However, for the stock split sample 7.92% was generated the first year while the same period for the dividend initiation sample showed 6.0% abnormal return. Compounding these numbers over two years (at normal returns) brings us close to the abnormal three-year performance for these samples.

A final point to consider is the use of bootstrapped p-values to evaluate statistical significance. Several studies generate an empirical distribution of abnormal return using

[11] Agrawal, Jaffe and Mandelker (1992) argue that this is an artifact of the sample period used by Franks, Harris and Titman (1991).

[12] This is also pointed out by Mitchell and Stafford (1997) and Fama (1998).

Table 2.5
Monthly post-event abnormal return for corporate events

The table reports benchmarks used in the studied cited in the following way: Upper case letters 'S', 'B', 'M', and 'I' refer to Size, Book-to-market, Momentum, and Industry matching respectively. Lower case letters 'p' and 'f' refer to the use of matching portfolios and matching firms respectively. For example, 'SBp' means that the study used a size/book-to-market reference portfolio as benchmark. 'FF93' refer to the Fama-French three-factor model, 'DM86' refer to the Dimson and Marsh (1986) benchmark (a market model with a size factor appended), and 'MF' refer to a multifactor benchmark. A * indicates that the results are statistically significant at conventional levels.

	Sample Size	Sample Period	Return Horizon (Months)	Bench-mark	Monthly Average Abnormal Return
Repurchase tender offer					
Lakonishok and Vermaelen (1990)	221	1962–86	24	Sp	0.36*
Mitchell and Stafford (1997)	448	1959–93	36	FF93	0.15
Open market share repurchase					
Ikenberry, Lakonishok and Vermaelen (1995)	1,239	1980–90	48	SBf	0.25*
Mitchell and Stafford (1997)	2,432	1959–93	36	FF93	0.17*
Dividend initiations (increases)					
Michaely, Thaler and Womack (1995)	561	1964–88	36	Sp	0.34
Chemmanur and Liu (1995)	2,230	1965–92	36	Sp	0.34*
Dividend omissions (decreases)					
Michaely, Thaler and Womack (1995)	887	1964–88	36	Sp	−0.54*
Chemmanur and Liu (1995)	1,001	1965–92	36	Sp	−0.17*
Stock splits					
Desai and Jain (1997)	5,596	1976–91	36	SBMp	0.33*
Ikenberry, Rankine and Stice (1996)	1,275	1975–90	36	SBp	0.34*
Proxy Contests					
Ikenberry and Lakonishok (1993)	97	1968–88	56	DM86	−0.24
Wahal (1996)	119	1987–93	12	Ip	−0.37
Exchange Listing					
Dharan and Ikenberry (1995)	2,889	1962–90	36	SBp	−0.20
Announcement of positive unexpected earnings					
Bernard and Thomas (1989)	1,630	1974–86	2	Sp	1.0*
Announcement of negative unexpected earnings					
Bernard and Thomas (1989)	1,630	1974–86	2	Sp	−1.0*
New Analyst Buy Recommendation					
Womack (1996)	694	1989–90	6	Sp	0.3
New Analyst Sell Recommendation					
Womack (1996)	209	1989–91	6	Sp	1.52*
Mergers and Acquisitions (performance for buyers)					
Rau and Vermaelen (1998)	3,169	1980–91	36	SBp	−0.42*
Agrawal, Jaffe and Mandelker (1992)	937	1955–87	60	DM86	−0.17*
Franks, Harris and Titman (1991)	399	1975–84	36	MF	−0.08
Loughran and Vijh (1997)	947	1970–89	60	Sf	−0.11
Mitchell and Stafford (1997)	2,906	1959–93	36	FF93	−0.22*

the procedure described in section 2.1. However, when computing the p-value, several studies count the number of observations that are of the same sign but are more extreme than the observed abnormal performance in the actual event-sample under study. In other words, a one-sided statistical test. Regardless of the null hypothesis lined up for the study, choosing a one-sided test based on the sign of the sample mean, results in inflated statistical significance. If the null hypothesis is that there are no abnormal performance (as it would be under the market efficiency hypothesis) the test should be two sided.

2.3 Conclusion

Under the efficient market hypothesis, security prices reflect the information that is available to market participants, implying that it should be impossible to obtain abnormal returns from an investment strategy that only exploits public information. Thus, the results reported in table 2.5 seem to be a serious blow to the efficient market hypothesis. Fama (1998) makes two arguments against this interpretation of the results from the long-run performance literature. First, he points out that abnormal performance is sensitive to the method used to evaluate performance. In particular, abnormal performance tends to be reduced when applying the Fama-French three-factor model as benchmark for value-weighted event-firm portfolios. Second, abnormal performance after specific events may exist even if markets are efficient—a random split between underperformance and overperformance is consistent with market efficiency. Counting the number of events that is followed by stock price underperformance and the number of events that is followed by stock price overperformance, Fama (1998) points out that overperformance is about as common as underperformance.

Although the above argument obviously is correct, market efficiency also requires that any abnormal performance is not persistent over time. However, it is difficult to assess the persistence of, say five-year abnormal returns, when the time-series used to estimate the performance is less than 20 years (as it is in most studies). Thus, as is apparent from the review in this chapter, the focus of the debate on long-run abnormal performance have been on the quality of the benchmark used to estimate abnormal returns. This is also the approach taken in the rest of this study. In particular, the main analysis of long-run performance is going to rest on a benchmark constructed using an approach similar to the one that uses the Fama-French model, however, the factors are constructed using prespecified macroeconomic variables known from previous studies to be related to security returns.

Chapter 3

Conditional long-run stock return performance

3.1 Introduction

The extant literature on long-run performance following corporate events assumes that equilibrium expected returns are constant over the portfolio holding period. However, as surveyed by Ferson (1995), and consistent with dynamic asset pricing models, there is empirical evidence that changes in security risk levels and factor risk premiums (and therefore expected returns) have predictable components related to publicly available information on economic fundamentals. For example, if current corporate yield spreads indicate that a certain stock will have a relatively high expected return over the next period, failure to condition on this information will lead the econometrician to falsely identify "abnormal performance". This chapter discuses how to avoid confusing true abnormal performance with the effect of time-varying expected returns by explicitly conditioning abnormal performance on a set of prespecified information variables.

3.2 Asset returns with time-varying expectations

Assume that the asset return generating process is:

$$R_{it} = E_{t-1}(R_{it}) + \beta'_{it-1} f_t + e_{it}, \tag{3.1}$$

where R_{it} is return on asset i between time $t-1$ and time t ($t = 1, 2, \ldots, T$), $E_{t-1}(\cdot)$ is a conditional expectation using information that is available at time $t-1$, the $(K \times 1)$ vector β_{it-1} contains conditional factor sensitivities which measure the systematic risk of security i relative to the factors, f_t is a $(K \times 1)$ vector of random risk factors with $E_{t-1}(f_t) = 0$, and e_{it} is idiosyncratic risk with $E_{t-1}(e_{it}) = 0$.

Given that factor risk is pervasive and that idiosyncratic risk can be eliminated by holding well diversified portfolios, the arbitrage pricing theory (APT) first developed by Ross (1976) and later extended to an intertemporal version (that allows time-varying betas and time-varying risk-premiums) by Connor and Korajczyk (1990), implies that the conditional expected return on asset i is approximately linear in their sensitivities to factor risk. Assume that conditional expected return on asset i is exactly linear in the sensitivities:

$$E_{t-1}(R_{it}) = \lambda_{0t-1} + \beta'_{it-1}\lambda_{t-1}, \tag{3.2}$$

where λ_{0t-1} is the conditional expected risk-premium on a "zero-beta" portfolio,[1] and λ_{t-1} is a $(K \times 1)$ vector containing expected factor risk-premiums. The "$t-1$" subscript indicates that expectations about risk premiums for period t (the period starting at $t-1$ and ending at t) are formed using information available at time $t-1$. Substituting equation (3.2) into (3.1) gives the asset return generating process:

$$R_{it} = \lambda_{0t-1} + \beta'_{it-1}(f_t + \lambda_{t-1}) + e_{it}. \tag{3.3}$$

The maintained hypothesis throughout this chapter is that equation (3.3) represents the true asset return generating process.

However, in order to facilitate the econometric specification of the asset return generation process we need to put structure on the time-variation in factor sensitivities and risk-premiums and define risk factors. When risk factors are defined so that $(f_t + \lambda_{t-1})$ represents portfolio returns, we implicitly impose further restrictions on the asset return generating process. We follow Shanken (1992) and define the K-vector f_t as composed of two subvectors $f_{1t} = R_{mt} - E_{t-1}(R_{mt})$ and $f_{2t} = M_t - E_{t-1}(M_t)$ with dimension k_1 and k_2 respectively. The k_1-vector R_{mt} represent risk factors as factor mimicking portfolio returns,[2] while the k_2-vector M_t represents risk factors as macroeconomic variables.

The factor mimicking portfolios must satisfy the expected return relation (3.2), thus,

[1] The zero-beta portfolio has zero sensitivity to all risk factors

[2] The j'th factor mimicking portfolio has unit sensitivity to risk-factor j and zero sensitivity to all other risk factors.

the following restriction apply:[3]

$$\lambda_{1t-1} = E_{t-1}(R_{mt}) - \lambda_{0t-1} 1_{k_1}, \tag{3.4}$$

where 1_{k_1} is a $(k_1 \times 1)$ vector of ones and λ_{1t-1} is the first k_1 elements of λ_{t-1}.[4] Using this restriction and the definition of f_2 together with (3.1) and (3.2) gives the asset return generating process:

$$R_{it} = \lambda_{0t-1} + \beta'_{1it-1}(R_{mt} - \lambda_{0t-1} 1_{k_1}) + \beta'_{2it-1}(f_2 + \lambda_{2t-1}) + e_{it}, \tag{3.5}$$

where β_{jit-1} $(j = 1, 2)$ are the two k_1 and k_2 subvectors of β_{it-1}.

3.3 Conditional Jensen's alpha

When factor mimicking portfolios are the sole representation of risk factors $(k_2 = 0)$, taking the conditional expectation of both sides of equation (3.5) generates equation (3.2). Thus, monthly conditional expected returns may be estimated by regressing R_{mt} on R_{it}. However, when some factors are not portfolio returns $(k_2 > 0)$, the monthly conditional expected returns are not identified unless we specify a model for the conditional expectations of the macro economic variables used to represent risk factors. This has often been solved by subtracting out the time-series mean or taking the first difference of the macro economic variable. This approach is clearly not in the spirit of a model of conditional expected returns. Instead, and in order to avoid setting up a parametric model of the time-series behavior of macro economic variables, we follow an approach used by for example Ferson and Korajczyk (1995), and construct stock portfolios that mimic the macro economic variables. The precise nature of the construction of these portfolios will be made clear in the empirical chapters. However, the implication of the approach is that k_2 is zero, and the asset return generating

[3] This follows from:

$$E_{t-1}(R_{kt}) = \lambda_{0t-1} + \beta_{kt-1}\lambda_{1kt-1}, \quad \forall\ k = 1, \ldots k_1.$$

which, using the fact that element j in β_{kt-1} equals 1 for $j = k$ and zero otherwise, reduces to

$$\lambda_{1kt-1} = E_{t-1}(R_{kt}) - \lambda_{0t-1}, \quad \forall\ k = 1, \ldots k_1.$$

[4] Throughout the chapter a "1" with a subscript is used to denote a column vector of ones with number of rows equal to the subscript.

process becomes:

$$R_{it} = (1 - \beta'_{1it-1} 1_K)\lambda_{0t-1} + \beta'_{1it-1} R_{mt} + e_{it}. \tag{3.6}$$

This restricted model is analogous to the restriction implied by the "zero-beta" version of the Sharp-Lintner-Mossin Capital Asset Pricing Model (CAPM) tested by Gibbons (1982).

In order to estimate the model in (3.6) using a time-series of observed returns, we must specify a model for the time-variation in the factor loadings and make further assumptions about λ_{0t-1}. Following for example Ferson and Schadt (1996), it is assumed that factor loadings are linearly related to a set of L information variables Z_{t-1}:

$$\beta_{1it-1} = b_{i0} + B_{i1} Z_{t-1} \tag{3.7}$$

where b_{i0} is a K-vector of "average" factor loadings that are time-invariant and B_{i1} is a $(K \times L)$ coefficient matrix. The product $B_{i1} Z_{t-1}$ captures the predictable time variation in the factor loadings.

The expected return on the zero-beta portfolio can either be assumed to be a constant that must be estimated (i.e., $\lambda_{0t-1} = \lambda_0 \quad \forall\ t$), or made observable by assuming that a risk free asset exist, in which case λ_{0t-1} equals the risk free rate of return over the period starting on $t-1$ and ending on t (the risk free rate of return is known at the beginning of this period.) In the rest of this chapter, it is assumed that a risk free asset exists. Under this assumption, and using (3.7), the return generating processes in (3.6) becomes:

$$r_{it} = b'_{i0} r_{mt} + b'_{i1}(Z_{t-1} \otimes r_{mt}) + e \tag{3.8}$$

where r_{it} and r_{mt} are asset return and returns on factor mimicking portfolios in excess of the return on the risk free asset, the KL-vector b_{i1} is $\text{vec}(B_{i1})$, and the symbol $\otimes$ denotes the Kronecker product.[5] To test for abnormal performance, we follow the tradition started by Jensen (1968), and append a constant term α_i to the model. Taking the conditional expectation of both sides in (3.8) shows that the model implies that α_i should be zero. Moreover, the constant term is a measure of the average monthly abnormal return.[6] If we assume that b_{i1} is zero, the model reduces to a constant beta model such as the CAPM or

[5]The operator vec(·) vectorizes the matrix argument by stacking each column starting with the first column of the matrix. The Kronecker product of two matrices ($A \otimes B$) multiplies B with every element in A.

[6]Ferson and Schadt (1996) use the model in (3.8) to evaluate mutual fund performance.

the Fama and French (1993) three-factor model.[7]

To estimate the conditional Jensen's alpha, the returns on a portfolio of issuing firms are used as the dependent variable in a time-series regression. Monthly return on this "issuer portfolio" is computed as the weighted sum of the returns on the stocks of firms that have issued equity within the last five years. The time-series series regression can be written as:

$$r_p = [1_T \;\; r_m \;\; Z] \begin{bmatrix} \alpha_p \\ b_{p0} \\ b_{p1} \end{bmatrix} + e_p, \tag{3.9}$$

where r_p is a $(T \times 1)$ vector of issuer portfolio excess returns, 1_T is a $(T \times 1)$ vector of ones, r_m is a $(T \times K)$ matrix of excess returns on factor mimicking portfolios with typical row r'_{mt}, Z is a $(T \times LK)$ matrix of information variables multiplied by excess returns on factor mimicking portfolios with typical row $(Z_{t-1} \otimes r_{mt})'$, and e_p is a $(T \times 1)$ vector of idiosyncratic risk. The components of the coefficient vector have been defined above except that the previously used subscript i is replaces by p to indicate that the returns in r_p are portfolio returns. Since the firms in the issuer portfolio change frequently, the assumption that betas are time-varying seems particularly appropriate. Moreover, since the number of firms in the portfolio also varies considerably over the sample period, the covariance matrix of the error term e_p is likely not to satisfy the assumption $\text{cov}(e_p) = \sigma^2 I_T$ which is required to make OLS the Best Linear Unbiased Estimator (BLUE). It is more likely that the variance of the error term is different between months due to the changing number and identity of firms in the issuer portfolio. To allow for this, and still use OLS to estimate the coefficients, we estimate the covariance matrix of the coefficient vector using the heteroskedasticity consistent estimator of White (1980).

3.3.1 The Fama-French three factor model

Under the assumption that expected returns are time-varying in the sense defined by (3.2), it becomes interesting to explore the implications for the long-run stock return performance measures discussed in chapter 2. A widely used abnormal performance measure is based on the Fama-French three-factor model which assumes that expected returns are constant through time. The abnormal return of a portfolio p is estimated by regressing the portfolio-returns on a constant term α_p and the three Fama-French risk-factors. Given the model

[7] In the CAPM r_{mt} is the excess return on the portfolio of aggregated wealth. In the Fama-French three-factor model r_{mt} contains the excess return on the value-weighted market portfolio, and two zero-investment portfolios constructed based on market capitalization and book-to-market sorted portfolios.

assumptions underlying the Fama-French model, the estimated constant term measures the average abnormal return of portfolio p over the estimation period. Assume that the true model of asset returns is (3.9). Estimating abnormal performance using the Fama-French model implies running the following regression:

$$r_p = \left[\begin{array}{cc} 1_T & r_m \end{array} \right] \left[\begin{array}{c} \alpha_p \\ b_{p0} \end{array} \right] + u_p.$$

The OLS estimate for the coefficient vector is:

$$\left[\begin{array}{c} \hat{\alpha}_p \\ \hat{b}_{p0} \end{array} \right] = \left(\left[\begin{array}{cc} 1_T & r_m \end{array} \right]' \left[\begin{array}{cc} 1_T & r_m \end{array} \right] \right)^{-1} \left[\begin{array}{cc} 1_T & r_m \end{array} \right]' r_p.$$

Replacing r_p with the *true* model and taking the expected value of the OLS estimate gives:

$$E \left[\begin{array}{c} \hat{\alpha}_p \\ \hat{b}_{p0} \end{array} \right] = \left[\begin{array}{c} \alpha_p \\ b_{p0} \end{array} \right] + E \left(\left[\begin{array}{cc} 1_T & r_m \end{array} \right]' \left[\begin{array}{cc} 1_T & r_m \end{array} \right] \right)^{-1} \left[\begin{array}{cc} 1_T & r_m \end{array} \right]' Z b_{p1}.$$

The only case in which this OLS estimator is unbiased is when all the elements in the vector b_{p1} is zero. Equation (3.7) shows that this implies an asset return generating model with constant betas. This should not be surprising since the Fama-French regression is based on this assumption, and OLS estimates will be unbiased when the model estimated is the true model.

The derivation used to compute the bias in the OLS estimates of the coefficients in the Fama-French model also highlights the isomorphic relationship between a model with time-varying betas and K risk factors and a model with constant betas and more that K risk factors. The model in equation (3.8) was derived using time-varying betas. However, a non-zero b_{i1} is also consistent with the existence of additional risk-factors that are correlated with the predetermined information variables in Z_{t-1}. Thus, even if the Fama-French model is estimated using time-varying betas, but the Fama-French factors only capture a subset of the K risk factors, an omitted variable bias similar to the above bias will occur in the OLS estimates. This is, of course, not just a potential problem for the Fama-French model. Any empirical specification of the return generating process will suffer from the same problem if the specification only includes a subset of the relevant risk factors.

Chapter 4

SEO performance: The U.S. evidence[†]

There is substantial evidence that news of seasoned equity offerings are associated with an average two-day announcement period abnormal stock return of −3%. However, recent studies of the long-run performance of issuing firms question whether this market reaction to issue announcements is indeed rational. In particular, Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) find that firms making seasoned equity offerings systematically underperform a control group of non-issuing firms. This finding is coined "The new issues puzzle" by Loughran and Ritter (1995). This chapter provides three primary contributions to the debate on this puzzle. First, contrary to the literature, our long-run performance estimates are linked to the initial issue announcement effect. As a result, we are in a position to make in-sample inferences concerning whether the −3% market reaction to equity offerings represents an unbiased estimate of the future. Second, we examine a range of issue and issuer characteristics not presented in earlier papers, including convertible and straight debt in addition to equity, rights versus underwriting as the flotation method, and whether the issuer is a industrial firm or a public utility. Finally, we draw on recent asset pricing econometrics to generate time-varying expected return benchmarks in the estimation of long-run performance. Replication of the conventional matching procedures

[†]This chapter is joint work with B. Espen Eckbo at the Amos Tuck School of Business Administration, Dartmouth College, and Ronald W. Masulis at the Owen Graduate School of Management, Vanderbilt University.

yields similar evidence of long term stock return underperformance. However, this evidence of underperformance disappears when employing time-varying multifactor expected return benchmarks. This finding is robust across type of security (equity or debt), and across issuer types (industrial firms and public utilities). We cannot reject the hypothesis that the −3% average equity issue announcement effect is unbiased and consistent with informationally efficient markets.

4.1 Introduction

Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) report that common stock returns of industrial firms making seasoned equity offerings (SEOs) tend to substantially underperform those of a control group of non-issuing firms over the 2-5 years following the offering date. More specifically, Loughran and Ritter find that the average five-year stock return following SEOs is 60 percentage points below that of non-issuing firms of the same size. Their evidence suggests that the market reaction to SEO announcements, which averages −3% for industrial firm-commitment offerings by firms listed on the New York- (NYSE) and the American (Amex) stock exchanges,[1] is informationally inefficient: "... if the market fully reacted to the information implied by an equity issue announcement, the average announcement effect would be −33%, not −3%." (Loughran and Ritter (1995), p.48). If true, this "new issues puzzle" constitutes a serious challenge to the presumption of rational pricing in security markets. Not surprisingly, these studies have generated substantial interest among empirical researchers in corporate finance.

However, there is growing evidence that long-run performance estimates are highly sensitive to the econometric methodology used. For example, Spiess and Affleck-Graves (1995), Jegadeesh (1997), as well as our results, indicate that stock price underperformance following security offerings is substantially reduced when both size and book-to-market factors are used to select matching firms. Furthermore, the matching procedure is ineffective at creating portfolios of like risk or expected returns to the event portfolios of security issuers. Moreover, when employing the Fama and French (1993) three-factor model and value-weighting portfolio returns, Brav, Geczy and Gompers (1995) and Mitchell and Stafford (1997) report statistically insignificant long-run abnormal performance following SEOs.[2] This study

[1]See Asquith and Mullins (1986), Masulis and Korwar (1986), and Eckbo and Masulis (1992). The −3% represents the average abnormal (market model) return to firm commitment underwritten SEOs over the two-day period ending with the first Wall Street Journal announcement of the issue.

[2]Fama (1998) presents a discussion of these and other long-run performance studies in light of their implications (or lack thereof) for market efficiency.

draws a similar conclusion when estimating long-run performance using either the Fama-French model or a model where factors are constructed from macroeconomic variables.

The extant literature on long run performance following corporate events assumes that equilibrium expected returns are constant over the portfolio holding period. In this study, we show that the use of conditional (time-varying) expected return benchmarks also drives estimated long-run underperformance to statistical insignificance. As surveyed by Ferson (1995), and consistent with dynamic asset pricing models, there is empirical evidence that changes in security risk levels and factor risk premiums (and therefore expected returns) have predictable components related to publicly available information on economic fundamentals. For example, if current corporate yield spreads indicate that a certain stock will have a relatively high expected return over the next period, failure to condition on this information will lead the econometrician to falsely identify "abnormal performance". Thus, our conditional long-run performance estimates reduce the risk of confusing true abnormal performance with the effect of time-varying expected returns.

The evidence derived from our multifactor model is also of particular interest in light of the ongoing debate over whether risk factors based on predetermined attributes of common stocks, such as those in the Fama-French model, are themselves proxies for market mispricing. For example, the covariance between stock returns and a book-to-market factor may be driven either by exposure to underlying macroeconomic risk factors or by mispricing. As pointed out by Ferson and Harvey (1998), a factor model containing macroeconomic risk factors avoids this ambiguity.[3] A somewhat related argument is made by Loughran and Ritter (1999) who warn that the Fama-French factors may be "contaminated" by a disproportionate number of equity issuing firms.[4] With these econometric issues in mind, it is interesting that our conditional multifactor model, which uses macroeconomic risk factors and is free of factor contamination, yields SEO long-run abnormal return estimates close to zero.

We estimate long-run performance using a sample of seasoned security offering announcements consisting of 2327 security offerings (1620 equity and 707 debt) on the NYSE and Amex over the 1963-1983 period. Our sample excludes security offerings whose announcements go unreported in the Wall Street Journal Index and Dow Jones New Retrieval System over this period, by NASDAQ issuers, as well as issues occurring during the "hot issue markets" in the mid to late 1980s. We also provide long-run performance estimates across

[3] See also the discussion in Berk (1995) on the role of equity market capitalization (size) as a risk factor.

[4] Specifically, they report that the small-firm- (S) and low-book-to-market (L) portfolios used to generate the Fama-French size (SMB) and book-to-market (HML) factors contain a disproportionally high number of issuing firms. See Fama and French (1993) for a further description of the SMB and HML risk factors.

subsamples classified by issuer type (industrial vs. public utility), by security type (common stock, straight debt, and convertible debt), and by equity flotation method (right vs. underwritten offerings).

Initially, when using the Loughran and Ritter (1995) average holding period procedure, and matching stocks on size (equity capitalization) only, we document substantial SEO underperformance much like that reported in the extant literature from different sample periods.[5] However, this long-run underperformance disappears economically and statistically when using time-varying expected return benchmarks in the context of a portfolio investment strategy.[6] This conclusion is robust with respect to the SEO flotation method and whether the issuer is an industrial firm or public utility. Moreover, we also report statistically insignificant long-run abnormal performance following debt issues using several alternative estimation procedures. Our evidence on debt issues, as well as on issuer type and flotation method, further supports the view that the "new issues puzzle" is about proper risk adjustment rather than about market underreaction to the news released in security issue announcements.

The rest of the chapter is organized as follows. Section 2 discusses the econometrics of long-run performance estimation, and it details how our methodology differs from the earlier literature. Section 3 describes the data selection and main sample characteristics. Section 4 discusses the empirical results using matching-sample techniques, while section 5 shows empirical estimates using factor model procedures. Section 6 concludes the paper.

4.2 Long-run abnormal performance estimators

The empirical analysis in section 4 and 5 implements two procedures for estimating long-run abnormal performance. First, under the *matched control firm procedure* (discussed in section 2.1 in chapter 2), abnormal returns are estimated using the realized stock return to a non-issuing firm matched to have similar risk characteristics as that of the issuer. Second, under the *factor model procedure* (see chapter 3), a multi-factor return generating process is used to form expected one-period returns for both issuing firms and non-issuing matched-firms.

[5] In their procedure, common stocks that are listed on a major exchange and have not undertaken an SEO or an IPO within the last five years become eligible to be a matching firm.

[6] The strategy consists of continually taking long positions in non-issuing matched firms, and offsetting short positions in issuers, using either equal- or value-weighting of returns.

4.2.1 Matched control firm procedures

Ritter (1991), Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) all estimate the long-run (up to five year) return to issuers and to a matched control sample of non-issuers following the offering date. Under the assumption that the issuer and the non-issuing match have identical risk characteristics, the expected value of the long-run buy-and-hold return difference between issuers and matching firms equals zero.[7] This null hypothesis is rejected by both Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995). With a sample of 3,702 SEOs on the NYSE/Amex/NASDAQ exchanges over the period 1970-1990, Loughran and Ritter (1995) estimate this average difference in holding period returns to be −60% for five-year holding periods. Similarly, Spiess and Affleck-Graves (1995) estimate the average difference to be −42% in a sample of 1,247 SEOs from 1975–1989 (627 NYSE/Amex issues and 620 NASDAQ issues). Both results raise the question of whether the matching procedure adequately controls for risk.

We implement a number of modifications to the standard matching procedure in this literature for selecting non-issuing control firms and for computing the relative performance of portfolios of issuers and matched firms. Loughran and Ritter (1995) select a non-issuing match based on equity capitalization at the prior year-end,[8] and Spiess and Affleck-Graves (1995) control for size and the two-digit industry SIC code of the issuer. In contrast, we match on size and book-to-market ratio at the prior month-end, i.e., two of the risk factors in the three-factor model of Fama and French (1993), and we separate industrial firms from public utilities.[9] In each month t, size and book-to-market matching is accomplished by first identifying all companies with an equity capitalization in the interval $[(v_i/1.3), 1.3v_i]$, where v_i is the market value of the issuer's common stock at the end of month t.[10] The company within this size-interval with the closest book-to-market ratio to the issuer is then selected as the matching firm for month $t+1$.[11] Moreover, as discussed in the empirical section below, we also examine the effect of the prior listing requirements, as well as the effect of updating the list of matching firms monthly in response to delistings and security

[7] We are unaware of any long-run performance studies using the matched control firm procedure that reports evidence to support this critical assumption.

[8] They select the non-issuing firm with the closest but greater size than the SEO firm.

[9] Spiess and Affleck-Graves (1995) state that their industry matching procedure leads to results that are indistinguishable from a match based on book-to-market. Brav, Geczy and Gompers (1995) and Jegadeesh (1997) also analyze the effect of size vs. size/book-to-market matching. Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) exclude utilities from their analysis.)

[10] A similar sorting rule is used by Barber and Lyon (1997).

[11] Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) match on the closest but *larger* firm measured by equity capitalization when employing the size-matching procedure.

issues by these firms.

4.2.2 Factor model procedures

Factor model procedures assume that expected returns are generated by a set of K pre-specified risk factors. Following a tradition started by Jensen (1968), the abnormal return of portfolio p is estimated by regressing the returns of portfolio p on a constant term α_p and the K risk factors. The estimated constant term is "Jensen's alpha" and represents the average abnormal return of portfolio p over the estimation period. The expectation of Jensen's alpha equals zero for passively held portfolios provided the specified factor model adequately captures the pervasive risk factors underlying the economy.

We employ two alternative factor model specifications in this investigation. The first model we use is the Fama and French (1993) three-factor model:

$$r_{pt} = \alpha_p^{\text{ff}} + \beta_{p1}\text{MR}_t + \beta_{p2}\text{SMB}_t + \beta_{p3}\text{HML}_t + e_{pt}^{\text{ff}}, \tag{4.1}$$

where r_{pt} is the one-period (monthly) return on a portfolio of issuing firms in excess of the one-month Treasury bill, MR_t is the return on the value-weighted market portfolio of all NYSE/Amex/NASDAQ stocks, SMB_t is the return on a portfolio of small firms minus the return on a portfolio of large firms, HML_t is the return on a portfolio of firms with high book-to-market values minus the return on a portfolio of firms with low book-to-market values, and the slope coefficients β_{pj} represent portfolio p's exposures to each of the three risk factors (p's systematic risks). The constant term α_p^{ff} is portfolio p's Jensen's alpha, i.e., the monthly average abnormal performance over the estimation period. Equation (4.1) is estimated over the 1963–1988 period.

Second, we estimate Jensen's alpha using the K-factor model (from equation (3.8) in chapter 3):

$$r_{it} = \alpha_p + b_{i0}' r_{mt} + b_{i1}'(Z_{t-1} \otimes r_{mt}) + e_{it}, \tag{4.2}$$

where r_{it} and r_{mt} are asset return and returns on factor mimicking portfolios in excess of the return on the risk free asset, the K-vector b_{i0} measures average factor loadings, the KL-vector b_{i1} is designed to pick up predictable time-variation in factor loadings. As indicated in chapter 3 and in the introduction to this chapter, the motivation for the conditional model framework is the growing evidence that expected returns are predictable using publicly available information.[12] In the presence of time-varying expected returns, an estimate of Jensen's alpha derived from an unconditional model is a biased measure of the true abnormal performance, and our conditional factor model estimation represents an attempt to correct for this bias.

4.3 Sample characteristics and two-day announcement effects

This section explains the sample selection procedure and provides descriptive statistics of the issuer sample. We also provide estimates of the two-day announcement effect for the various security offering categories. At the heart of the "new issues puzzle" is the question of whether the two-day announcement effect represents an unbiased estimate of the subsequent long-run performance of the issuing firms' shares.

4.3.1 Sample characteristics

The sample of SEOs used in this paper is drawn from Eckbo and Masulis (1992) with the addition of the years 1982 and 1983 following a similar sampling process. The sample of debt offerings are from Eckbo (1986). The two samples reflect the following restrictions:

(1) The common stock of security issuers is listed on either the NYSE or the Amex at the time of the offering.[13] Moreover, all issuer stocks are found on the University of Chicago CRSP monthly stock return file, and all debt issuers are in the 1982 COMPUSTAT Annual Industrial (Primary, Secondary, Tertiary, and Historical Research) files as well. When book-to-market ratios are used, all stock issuers need to have book equity figures reported by COMPUSTAT.

(2) The issues are publicly announced prior to the offering date. The debt issues are all announced in the Wall Street Journal Index, while the source of the SEO announcements also includes the Investment Dealer's Digest, Moody's Industrial and Utilities

[12]Ferson and Harvey (1991) and Evans (1994) argue that time-variation in conditional betas for passive portfolios is economically and statistically small in the U.S. However, Ferson and Schadt (1996) find that time-varying betas are important in their measurement of the performance of managed U.S. mutual funds. Moreover, it is commonly accepted that conditional expected risk premiums tend to vary with economy-wide factors such as the business cycle.

[13]More specifically, the stock must be listed between the initial announcement date and the public offering date of the security issue. Approximately half of the industrial offers are by NYSE-listed firms and the other half by Amex-listed companies. Almost all the utilities are NYSE-listed reflecting their typically large size.

Manuals, Dow Jones News Retrieval Service, LEXIS, and the SEC Registered Offerings Statistics (ROS) tape.

(3) For debt issuers, all issues are for cash, there are no simultaneous offers of equity instruments, and overseas issues are excluded. For SEOs, there are no simultaneous offer of debt, preferred stock or warrants. Moreover, combination primary/secondary stock offers, canceled or postponed offers, and non-U.S. offers and non-U.S. issuers are excluded.

(4) All offers permit classification as to the flotation method. Information on the flotation method is found in offering prospectuses, in the Investment Dealer's Digest's Corporate Financing Directory, in the "Rights Distributed" section of Moody's Dividend Record, as well as other Moody's manuals.

The data base further reflects a minimum size restriction on debt issues as well as on the debt issuer's leverage change. The size restriction eliminates the possibility that the firm's overall leverage ratio decreases over the year of the debt issue(s):

(5) A debt issuer is included in the sample only in years where the firm increased its book value of long-term debt by a minimum amount (with no offsetting reduction in short-term liabilities). The minimum leverage increase is $50 million in 1964 and increases by 5% a year to $114 million in 1981. Moreover, a debt offering in year t is included in the sample only if its size is greater than or equal to the minimum leverage increase for year t.[14]

This selection procedure produces a total of 1,620 SEOs and 707 corporate debt offers over the period 1963 through 1983. The 1,620 SEOs are by 761 separate firms, i.e., an average of 2.1 SEO per issuer over the observation period. The bond offerings are made by 208 different companies, with an average of 3.4 offerings per firm.

Table 4.1 shows the annual distribution of seasoned public offerings classified by: issuer type (industrial firm vs. public utility), equity flotation method (preemptive rights vs. firm commitment underwriting), and debt type (straight vs. convertible). Public utility issuers represent 51% (830 cases) of the equity issue sample and 35% (249 cases) of the debt issue sample.[15] Utility issuers are examined separately as their investment and financing policies

[14] Eckbo (1986) reports that, in a typical year, the ratio of the number of annual debt issues to the number of debt *retirements* by the sample firms is approximately 45 to 2. Overall, one can confidently argue that the debt offerings in our sample represent leverage-increasing events.

[15] Utilities are defined as firms with CRSP SIC codes in the interval [4910, 4939]. This classification differ slightly from the one used originally by Eckbo and Masulis (1992).

Table 4.1
Number of seasoned equity and debt issues by NYSE- and Amex-listed firms in the sample, 1963–1983

The sample period for Seasoned Equity Offerings is 1963–1983, which is slightly different from the sample period for debt offerings 1964–1981. The sample only contains issuers with common stock listed on NYSE or Amex (CRSP share code 10 or 11). Utilities (Utl) are defined as firms with CRSP SIC codes in the interval [4910, 4939].

	Seasoned Equity Issues							Bond Issues[a]						
		Firm-commitment offers			Rights offers[b]				Straight debt offers[c]			Convertible debt offers		
Year	Tot	Tot	Ind	Utl	Tot	Ind	Utl	Tot	Tot	Ind	Utl	Tot	Ind	Utl
1963	12	2	2	0	10	7	3	–	–	–	–	–	–	–
1964	17	8	4	4	9	6	3	5	4	4	0	1	1	0
1965	20	5	5	0	15	11	4	16	10	3	7	6	4	2
1966	25	12	9	3	13	12	1	29	23	15	8	6	6	0
1967	26	12	9	3	14	11	3	46	31	17	14	15	13	2
1968	44	26	20	6	18	12	6	24	18	9	9	6	5	1
1969	39	24	13	11	15	7	8	38	33	11	22	5	4	1
1970	49	36	17	19	13	4	9	61	55	34	21	6	6	0
1971	82	63	41	22	19	6	13	50	40	18	22	10	10	0
1972	79	66	27	39	13	3	10	29	28	13	15	1	1	0
1973	56	48	9	39	8	3	5	20	18	7	11	2	2	0
1974	53	47	8	39	6	2	4	54	53	37	16	1	1	0
1975	88	79	22	57	9	1	8	47	46	33	13	1	1	0
1976	92	88	31	57	4	1	3	31	31	24	7	0	0	0
1977	65	62	4	58	3	0	3	28	28	18	10	0	0	0
1978	90	86	24	62	4	3	1	35	35	26	9	0	0	0
1979	85	81	24	57	4	2	2	48	44	29	15	4	4	0
1980	160	155	85	70	5	3	2	69	65	44	21	4	3	1
1981	151	148	68	80	3	2	1	77	70	48	22	7	7	0
1982	130	130	64	66	0	0	0	–	–	–	–	–	–	–
1983	257	257	208	49	0	0	0	–	–	–	–	–	–	–
Sum	1620	1435	694	741	185	96	89	707	632	390	242	75	68	7

[a]This category represents firm-commitment underwritten offerings.
[b]This category represents uninsured rights and rights with standby underwriting.
[c]Straight debt includes non-convertible bonds, notes and debentures.

are highly regulated. The regulatory policy is public knowledge and thus makes it less likely that a utility announcing a stock offer is attempting to take advantage of temporary market overpricing. For example, stock offers often require state utility commission approval or SEC approval for utility holding companies. In other instances, an equity offer is mandated if a utility rate increase is to be approved. This further lowers the probability that the utility is "timing" the equity market.

In the SEO sample, 11% (185 cases) are floated using preemptive rights, while the remaining issues are sold using firm commitment underwriting. As pointed out by Eckbo and Masulis (1992), rights offers virtually disappeared during the seventies for large U.S. corporations. In light of the theoretical and empirical results in that paper, equity rights offers are examined separately from firm commitment underwritten offerings. In an equity rights offering, the issuer employs short-term warrants which give current shareholders the right to purchase the new issue on a pro rata basis. There are two types of rights offers in our sample: "uninsured rights" where the issuer bears the risk that shareholders do not fully subscribe to the issue, and "rights with standby underwriting" (or simply "standby rights") where the issuer insures that any unsubscribed portion of the rights offer is take up or purchased by an investment bank. In contrast, a firm commitment underwritten offer does not employ rights, and the investment bank guarantees the sale of the entire issue at a fixed price.[16]

In the debt sample, 11% (75 cases) are convertible debt instruments, while the remaining are straight debt (non-convertible bonds, notes and debentures).[17] Convertible debt issues are singled out due to their hybrid debt-equity nature. If the probability of market mispricing increases with the idiosyncratic risk of the security, then convertibles rank somewhere between straight debt and equity in terms of potential adverse selection. The convertibles are almost exclusively issued by industrial firms (91% or 68 of the 75 cases). In addition to the 75 convertibles, there are 632 issues of straight debt, of which 189 are mortgage bonds (173 issued by public utilities). As reported by Eckbo (1986), the sample captures on average 68% of *all* announced offerings made by the sample firms during the 18-year sample period and, due to the minimum offer size requirement, a much greater proportion of the total dollar amount raised through new public debt offers.

[16]Eckbo and Masulis (1992) report that direct flotation costs for industrial issuers average 6%, 4%, and 1% of the offering proceeds for firm commitment offers, standbys and uninsured rights offers, respectively. See also Eckbo and Masulis (1995).

[17]Lewis and Seward (1997) observe an increased use of convertibles in the 1980s, following our sample period.

Table 4.2
Descriptive statistics of U.S. equity and debt issuers, 1963–1983

	Seasoned Equity Offerings					Bond Offerings[a]				
		Firm-commitment		Rights			Straight		Convertible	
	Tot	Ind	Utl	Ind	Utl	Tot	Ind	Utl	Ind	Utl
Number of observations[b]	1609	687	739	94	89	707	390	242	68	7
Amount offered ($ millions)	48	44	48	80	49	158	194	113	118	85
Market value of equity ($ millions)	664	713	554	1203	625	3285	4867	1237	1720	1090
Amount offered divided by market value of equity	0.16	0.21	0.12	0.18	0.08	0.13	0.13	0.12	0.14	0.10
Proportion of offers intended for capital investments	0.42	0.53	0.11	0.56	0.59	0.38	0.27	0.48	0.41	0.29
Book-to-market ratio[c]	0.81 (1333)	0.62 (564)	0.99 (638)	0.75 (42)	0.74 (89)	0.81 (636)	0.86 (354)	0.82 (219)	0.50 (57)	0.55 (6)

[a]Public offers of debt are sold by firm commitment underwriting contracts.

[b]The stated purpose of the equity offerings was to reduce debt in 426 cases, to fund capital expenditures in 219 cases, and both to reduce debt and fund capital expenditures in 587 cases. The remaining equity offerings could not be classified. The stated purpose of the debt offering was to reduce debt in 390 cases and to fund capital expenditures in 216 cases. The remaining debt offerings could not be classified.

[c]The source of this information is COMPUSTAT. The numbers in parentheses are sample sizes for the book-to-market ratio, which are lower than the total number of observations due to missing COMPUSTAT book values.

Table 4.2 lists sample statistics describing issue frequencies, issuer equity capitalization (size), the frequency with which the issues are intended to finance investments (as reported in the Wall Street Journal), and the book-to-market ratio. The average debt issue is typically three times the average SEO, causing approximately a 16% increase in equity and a 13% increase in debt. The average book-to-market ratio ranges from 0.50 for convertible debt issuers to 0.86 for industrial issuers of straight debt. For SEOs, the average book-to-market ratio ranges from 0.62 for firm commitment offers by industrials to 0.99 for firm commitment offerings by utilities. In the debt sample, the Wall Street Journal announcements indicate that the proceeds of 216 issues are earmarked for capital expenditure (investment) programs, another 390 issues are made to refund old debt, while the remaining 101 offerings lack information to classify them. In the SEO sample, the proceeds of 806 issues are used to finance capital expenditures and investments, 426 are used to finance debt reductions, while 388 offers cannot be classified.

4.3.2 Average two-day announcement effects

Eckbo (1986) and Eckbo and Masulis (1992) report two-day announcement effects for the debt and most of the equity issues in our sample. Both papers use a market model regression of the type

$$r_{it} = \alpha_i + \beta_i r_{mt} + \gamma_i d_t + \epsilon_{it}, \tag{4.3}$$

where r_{it} and r_{mt} denote the continuously compounded daily rates of return to issuer i and the value weighted market portfolio of all NYSE/Amex/NASDAQ stocks, ϵ_{it} is a mean zero error term, and d_t is a dummy variable which takes on a value of one during the two-day event period [day -1, day 0] and zero otherwise. Thus, the parameter γ_i measures the average daily announcement-induced abnormal return for stock i, so that the two-day abnormal return equals $2\gamma_i$.

Table 4.3 lists the average values of the two-day announcement effect, $(2/N)\sum_{i=1}^{N}\gamma_i$, reestimated for various subsamples of SEOs and debt offerings of size N. The estimation period is one year (252 trading days) on each side of the two day announcement period, starting on trading day -253. The table also reports the test statistic

$$z = (1/\sqrt{N})\sum_{i=1}^{N}(\hat{\gamma}_j/\hat{\sigma}_{\gamma_i}), \tag{4.4}$$

where the "hat" denotes an OLS estimate and $\hat{\sigma}_{\gamma_i}$ is the estimated standard error of $\hat{\gamma}_i$.

Table 4.3
Average pre-announcement date run-up and announcement-day abnormal returns to NYSE- and Amex-listed issuers of seasoned equity and debt, using a market model as the returns benchmark, 1963–1983

One year run-up and announcement day abnormal returns for issuer i are computed using the following market model:

$$r_{it} = \alpha_i + \beta_i r_{mt} + \sum_{j=1}^{2} \gamma_{ij} d_{jt} + \epsilon_{it},$$

where r_{it} is daily excess return on issuer i, r_{mt} is daily excess return on a value weighted market portfolio of all NYSE-, Amex-, and NASDAQ-listed firms. The estimation period is a total of two years, starting on trading day −253 relative to the announcement date and ending on trading day +253 relative to the announcement date. The dummy variable d_{1t} takes on a value of one on trading days −253 through −2. The dummy variable d_{2t} takes on a value of one in a two-day window starting the day before the announcement day. The percentage abnormal return for the 252-day run-up period is $252 \times \gamma_{i1} \times 100$, and the abnormal return over the two-day announcement window is $2 \times \gamma_{i2} \times 100$. Under the null hypothesis of zero abnormal return, the following test statistic converge in distribution to the standard normal

$$z_j = \frac{1}{\sqrt{N}} \sum_{i=1}^{N} \frac{\hat{\gamma}_{ij}}{\hat{\sigma}_{ij}}.$$

Where $\hat{\gamma}_{ij}$ is the OLS estimate of γ_{ij}, and $\hat{\sigma}_{ij}$ is the standard error of this estimate. Returns are continuously compounded. The parentheses contain number of observations and p-values. The p-values are for two-sided tests.

	Seasoned Equity Offerings			Bond Offerings	
	Firm-commitment	Standby rights	Uninsured rights	Straight	Convertible
(a) Industrial issuers					
One-year runup	37.6 (687; .000)	20.4 (46; .005)	7.3 (50; .133)	5.3 (390; .003)	20.6 (68; .000)
Two-day announcement return	−2.3 (687; .000)	−1.9 (46; .000)	−0.8 (50; .238)	−0.0 (390; .810)	−1.1 (68; .000)
(b) Utility issuers					
One-year runup	2.0 (729; .006)	2.2 (84; .285)	−6.3 (5; .595)	−1.2 (242; .674)	17.2 (7; .055)
Two-day announcement return	−0.5 (729; .000)	−1.1 (84; .000)	−0.2 (5; .790)	−0.0 (242; .960)	−0.1 (7; .640)

Under the null hypothesis of zero abnormal return, z is approximately standard normal for large samples.

Table 4.3 summarizes several stylized facts about the market reaction to seasoned security offerings. First, the average market reaction to firm commitment SEOs is significantly negative (−2.3% for industrial issuers and −0.5% for utility issuers). Second, the market reaction is significantly negative but smaller for rights offers with standby underwriting (−1.9% for industrials and −1.1% for utilities), and insignificantly different from zero for uninsured rights. The relative magnitude of the market reaction across flotation methods is predicted by the adverse selection model of Eckbo and Masulis (1992).[18] Moreover, the generally smaller reaction to utility issuers is also consistent with the lower adverse selection risk for this issuer category.

Third, the average market reaction to convertible debt offerings is significantly negative (−1.1%) while the reaction to straight debt offerings is insignificantly different from zero. Thus, there is an inverse relationship between the idiosyncratic risk of the security issued and the market reaction to the issue-announcement. As pointed out by earlier studies, this result conforms to the predictions of the Myers and Majluf (1984) adverse selection model and the Myers (1984) pecking order hypothesis.

Alternatively, the Loughran-Ritter 'new issues puzzle' suggests that the market systematically underreacts to security offering announcements. The underreaction hypothesis does not predict the differences in market reactions across security types, issuer types and flotation methods shown in Table 4.3. The rest of this paper is devoted to the question of whether the security offering categories listed in Table 4.1 are systematically followed by negative long-run abnormal performance, as suggested by the underreaction hypothesis.

4.4 Performance estimates using the matching procedure

4.4.1 Sample-wide averages

We begin our estimation using the particular matching firm procedure in Loughran and Ritter (1995). For the sample period 1970 to 1990, Loughran-Ritter report an average five-year holding period return of 33.4% for seasoned equity offerings by industrial firms. The average size-matched control firm yields 92.8% over the same five-year holding period, indicating a 59.4% underperformance by industrial SEOs. The first row of panel (a) of Table 4.4 reports our corresponding numbers. The average five-year holding period return for

[18] See also Bøhren, Eckbo and Michalsen (1997) for further evidence on this model.

industrial issuers is 52.1%, and for the size-matched control sample it is 84.3%, representing 32.2% underperformance by issuing firms. Thus, using the size-matching procedure, we obtain results qualitatively similar to Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) in our observation period.

In panel (a) of Table 4.4 where we replicate the Loughran-Ritter procedure, a matching firm is selected from a list of size-ranked firms generated in December prior to the year of the issue. If a matched firm delists or itself becomes an issuing firm some time during the five-year holding period, the firm is replaced by another match *from the original list.* To illustrate, suppose that a matching firm delists four years into the return cumulation, and that the initial equity capitalization of the issuer was $100 million. According to this procedure, a new match with a size of $100 million is selected from the original list. This, of course, does not ensure that the issuing firm and the new match have the same size four years later, at the time of the rematch. In panel (b) of Table 4.4, we ensure a similar size match by the following modification to the sampling procedure: Whenever there is a need to replace a matching firm, the new match is selected using information in the month prior to the replacement month, constraining the size of the matching firm to equal (or be as close as possible to) that of the issuer in that month. Furthermore, the initial matching firm is selected in the month prior to the month of the issue. As shown in panel (b), these changes reduce the average long-run buy-and-hold return to matching firms by 6.4% (from 84.3% to 77.9%), with a corresponding reduction in the negative long-run abnormal return to issuing firms.

Comparing the industrial issues in the first rows of panels (a) and (b) of 4.4, only 6.7% of the matching firm samples used in the two panels overlap. This lack of overlap is primarily driven by changing the month of the initial matching process from the calendar year-end prior to the issue in panel (a) to the month-end prior to the issue in panel (b). Of the 782 industrial equity issuers in the first row of panel (a), 576 require no rematches over the five-year holding period, 176 have a single rematch, 34 have two rematches, and 5 have three rematches over the period.[19] Alternatively, one could perform a systematic monthly rematching of *all* issuing firms for the entire five-year holding period. Such a procedure, which is likely to generate substantially greater differences than those implied by panels (a) and (b) is close in spirit to our conditional factor model approach described below and is therefore not included here.

Table 4.4 also shows the effect of using a five-year pre-event return requirement in the

[19]There is a similar frequency of rematches for the issues in panel (b).

Table 4.4
Five-year buy-and-hold returns (%) to firms undertaking seasoned public offerings on the NYSE and Amex, and their matched control sample, 1963–1983

The matches are chosen using size- and book-to-market matching. The size-matching is done using the equity market value of the issuer. Book-to-market matching is done by first selecting all companies that have an equity market value within 30% of that of the issuer. Then the company with the closest book-to-market value is chosen as the matching firm. Monthly book-to-market rankings in year t are created by dividing the end-of-year book-value from year $t-1$ with monthly market capitalizations for year t. The pre-event return requirement means that a match is required to have a five-year return history on CRSP prior to the announcement date. Numbers in the columns marked "issuer" and "match" are computed using:

$$\frac{1}{N}\sum_{i=1}^{N}\left[\prod_{t=\tau_i}^{T_i}(1+R_{it})-1\right]\times 100$$

The p-values in the column marked $p(t)$ are p-values of the t-statistic using a two-sided test of no difference in average five-year buy-and-hold return for issuer and matching firms. The p-values in the column marked $p(N)$ are bootstrapped $p-$values of a two-sided test.

Matching procedure	Industry	Seasoned Equity Offerings N	Issuer	Match	$p(t)$	$p(N)$	Bond Offerings N	Issuer	Match	$p(t)$	$p(N)$
(a) Match drawn from year-end-ranking prior to the first year of return cumulation											
Size and pre-event return requirement	Ind	782	52.1	84.3	.000	.000	453	36.5	38.8	.652	.511
	Utl	805	33.4	41.7	.017	.038	220	26.0	26.2	.968	.868
Size but no pre-event return requirement	Ind	782	52.1	82.5	.000	.000	453	36.5	39.3	.601	.816
	Utl	827	32.5	37.6	.131	.118	247	23.7	22.9	.883	.939
Size, book-to-market, and pre-event return requirement	Ind	595	49.0	56.7	.287	.520	371	38.4	43.5	.382	.592
	Utl	697	33.9	48.8	.000	.003	194	24.4	28.5	.475	.908
Size, book-to-market, but no pre-event return requirement	Ind	605	48.2	60.7	.088	.251	395	36.6	41.8	.342	.677
	Utl	718	32.9	46.0	.000	.010	216	21.7	24.8	.557	.968
(b) Match drawn from month-end-ranking prior to the first month of return cumulation											
Size and pre-event return requirement	Ind	790	51.3	77.9	.000	.005	453	36.5	40.5	.452	.483
	Utl	806	33.3	46.2	.000	.017	220	25.9	24.5	.799	.991
Size but no pre-event return requirement	Ind	790	51.3	77.4	.000	.012	453	36.5	38.6	.692	.903
	Utl	828	32.4	42.8	.003	.043	247	23.6	21.8	.705	.947
Size, book-to-market, and pre-event return requirement	Ind	609	47.2	55.1	.237	.578	391	36.8	39.3	.662	.849
	Utl	703	33.8	51.0	.000	.002	195	25.6	31.6	.353	.295
Size, book-to-market, but no pre-event return requirement	Ind	610	47.3	56.0	.203	.504	404	35.7	39.5	.482	.862
	Utl	724	32.8	49.1	.000	.000	217	22.8	27.9	.382	.389

selection of matching firms.[20] The five-year listing requirement (which is implemented by both Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995)) builds in a survivorship bias in the matched sample which, according to the simulations in Kothari and Warner (1997) and Barber and Lyon (1997), can induce an underperformance effect in the issuer returns of approximately 4% for a 3-year holding period. However, the effect appears to be much smaller in our sample. For example, comparing rows one and three in column five of panel (b), the effect of the five-year prelisting requirements for matching firms is to reduce the buy-and-hold return by only 0.5% for industrial firms.

We next modify the matching procedure by selecting matching firms based on both size and book-to-market ratio. Without the five-year pre-event return requirement and using size and book to market ratios from the month prior to the first month of return cumulation, the average difference in size between industrial SEO issuers and their matching firms is 1.37%. The average difference in book-to-market ratios is 2.93%.[21] When using only size matching the corresponding numbers are 0.5% and 227%.[22] These numbers indicate that it is the book-to-market ratio that is the binding constraint in the size and book-to-market matching procedure. Comparing the third and fourth rows of panel (b) in Table 4.4 with the last two rows of the same panel, we see that 180 industrials and 104 utilities are excluded from the sample due to missing COMPUSTAT equity book values. To ensure that our bootstrapped p-values are unbiased when we match on size- and book-to-market ratios, all stock returns on the CRSP tape with missing book-values are excluded.

Without the five-year pre-event return requirement, the average five-year buy-and-hold return for industrial SEOs is 47.3%, while the average return for the size- and book-to-market matched control firms is 56.0%. The difference, 8.7 percentage points, is not significantly different from zero using either a standard t-test or the bootstrapped p-values (neither using a one-sided nor a two-sided test).[23] This result differs from the results of

[20]Loughran and Ritter (1995) require issuing firms to have at least a five-year history of returns on CRSP. There is no such prior return history requirement in our sample of issuers.

[21]The issuers have larger market capitalization than their matching firms in 52% of the cases. Issuers have larger book-to-market ratios than the matching firms in 49% of the cases.

[22]The huge average percentage difference in book-to-market ratio is caused by a few outliers. The most extreme case is an issuer with a book-to-market ratio of 0.01 while the matching firm has a book-to-market ratio of 1.47. The average book-to-market ratio for issuers when matching firms are selected based on size only is 0.52 and for matching firms 0.88.

[23]The bootstrapped p-values reported in Table 4.4 are for the hypothesis that the sample mean is equal to the mean of the empirical distribution, $\hat{\mu}$, against the alternative hypothesis that sample mean is different from $\hat{\mu}$. That is, we employ a two-sided test. Given an observed sample mean of BH, the two-sided p-value is one minus the probability mass between BH and $2\hat{\mu} - BH$. One-sided p-values for the alternative hypothesis that five-year buy-and-hold returns for issuers are less than for matching firms are constructed as the probability mass below BH. The empirical distributions are roughly symmetric, and for distributions

the existing literature. Loughran and Ritter (1995) control for book-to-market effects using cross sectional- and time series regressions. They identify a book-to-market effect in returns. However, it does not eliminate their estimates of SEO underperformance. Spiess and Affleck-Graves (1995) control for the book-to-market effect using a method somewhat similar to the one used here.[24] Although they do not report the buy-and-hold returns based on size- and book-to-market matching, they state that the results are similar to size- and industry based returns which they report to be on average 55.7% for SEOs and 98.1% for the average matching firm. Using a standard t-test, the difference of −42.4% is significantly different from zero.

The long-run performance literature employ samples which exclude regulated utilities. Interestingly, as shown in the last row of Table 4.4, utility SEOs underperform size- and book-to-market matched firms by a statistically significant 16.3 percentage points, which is almost twice the 8.7% underperformance estimated for industrial issuers. This result should be evaluated in light of the fact that there is a difference in the quality of the matching procedure for utilities and industrials. The difference arises because utilities are both relatively high frequency equity issuers and they represent a small proportion of all NYSE-/Amex listed firms. As a result, a small number of utilities are sampled frequently as matching firms, which exposes the portfolio of matching firms to excessive non-systematic risk.[25] Moreover, utility issuers and their matches tend to exhibit a greater disparity (in term of size and book-to-market ratio) relative to the industrial issuers and their matches.[26]

Table 4.4 also reports the long-run performance of firms that issue debt. Again using size- and book-to-market matching and no pre-event return requirements, the average industrial debt issuer has a five-year buy-and-hold return of 35.7%, which is 3.8% lower than the 39.5% return to the average matching control firm. The corresponding numbers for utilities are 22.8% and 27.9%, respectively, i.e., an underperformance of 5.1%. These levels of underperformance are significantly different from zero at 10% significance level using a t-test, but are insignificant at conventional levels when we use the more appropriate boot-

using many observations, centered close to zero. Thus, one-sided p-values are about half their two-sided counterparts. The one-sided p-value for the hypothesis that −8.7% is less than the mean of the empirical distribution is 0.257.

[24]Spiess and Affleck-Graves (1995) choose the matching firm that minimize the sum of the absolute percentage difference between sizes and book-to-market values.

[25]To illustrate, in order to generate the set of matching firms for the 724 utility issues in the last row of Table 4.4, we need a total of 929 matching firms (due to delistings or equity issues by matching firms). This set of 929 matching firms turns out to contain only 118 different utilities, or 13%. The corresponding percentage of different matching firms used for industrial issuers is 77%.

[26]This concern favors the use of a factor model as return benchmark for the estimation of long-run abnormal returns for the sample of utility issuers.

strapped p-values. Note that Table 4.4 shows that p-values based on the t-statistic tend to overstate the level of statistical significance compared to the bootstrapped p-values. Thus, our inferences are primarily based on the bootstrapped p-values.[27]

Table 4.5 shows the means of the five-year abnormal returns for SEO and bond offer samples, grouped on an annual basis. The annual p-values are constructed by simulating the empirical distribution of average long-run abnormal performance using the same number of stocks as contained in our issuer samples for a given year. For example, there are 181 industrial SEOs in our sample for 1983. To construct the empirical distribution for SEOs in 1983, we draw (with replacement) 1000 samples of 181 stocks. For each firm in each of the 1000 samples, we select a size- and book-to-market matched control firm and compute the post-1983 five year abnormal stock performance level. Finally, an average abnormal performance level is computed over the 181 stocks in each sample. This generate the empirical distribution of five-year abnormal buy-and-hold returns.

The annual p-values for SEOs show that the performance of industrial issuers are indistinguishable from the performance of the control firms in all but two years (1969 and 1971) and we expect on average to find one significant case in twenty by chance. For utilities, we see that the underperformance documented in Table 4.4 is concentrated in the second half of the sample period. The average abnormal performance of utility SEOs is −23.8% over the 1975–1983 period. The pattern for bond offerings is much the same as that for industrial SEOs. With the exception of utilities in 1981, neither industrial nor utility debt issuers exhibit abnormal performance in any of the years.

4.4.2 Subsample results

Tables 4.6 and 4.7 show long-run abnormal returns for SEOs and bond offerings classified by flotation method and security type. Table 4.6 reports the long-run abnormal returns for firm-commitment underwritten offerings and rights offerings of common stock for industrials and utilities. Interestingly, although the abnormal returns in Table 4.6 are computed relative to size- and book-to-market matched control firms, panel (a) shows a mean one-year pre-announcement stock price run-up that is similar to the run-up documented in Table 4.3. The mean one-year stock price run-up for industrial issuers is 28.1% for firm-commitment underwritten offerings and and 8.8% for pure rights and standby offerings, where the latter is indistinguishable from zero based on conventional t-values.

[27]In the context of buy-and-hold abnormal returns, Barber and Lyon (1997) and Lyon, Barber and Tsai (1999) provide substantial evidence that bootstrapped p-values have better power than p-values based on conventional t-statistics.

Table 4.5
Annual means of five-year abnormal returns (%) to firms undertaking seasoned public offerings on the NYSE and Amex, using matching firm returns as the returns benchmark, 1963–1983

Abnormal returns for SEOs and public debt offerings classified as industrials and utilities. The abnormal returns are the difference between the five-year buy-and-hold returns for issuers and matching firms, denoted BH. The matching firms are chosen using size and book-to-market matching and without the five-year pre-event return requirement. The size and book-to-market rankings are from the month prior to the start date of return cumulation. The figures reported below are for $BH \times 100$, where

$$BH = \frac{1}{N}\sum_{i=1}^{N} BH_i = \frac{1}{N}\sum_{i=1}^{N}\left[\prod_{t=\tau_i}^{T_i}(1+R_{it}) - \prod_{t=\tau_i}^{T_i}(1+R_{it}^{m})\right]$$

Numbers in parentheses are bootstrapped p-values for a two-sided test of the hypothesis that BH equals the mean of the empirical distribution.

	Seasoned Equity Offerings				Bond Offerings			
Year	N	Industrials	N	Utilities	N	Industrials	N	Utilities
1963	3	46.2 (.774)	3	−13.6 (.622)	–	–	–	–
1964	3	4.9 (.906)	6	7.0 (.460)	3	10.9 (.925)	0	–
1965	4	−23.2 (.695)	4	−2.4 (.841)	7	15.1 (.991)	7	1.7 (.868)
1966	9	−31.8 (.552)	4	−4.0 (.840)	7	−0.3 (.937)	7	−0.9 (.811)
1967	8	−31.0 (.521)	5	2.8 (.754)	20	−31.9 (.398)	13	1.0 (.599)
1968	21	−7.9 (.730)	12	−12.9 (.162)	14	2.7 (.963)	8	−1.3 (.930)
1969	16	−44.9 (.040)	18	11.5 (.193)	15	29.2 (.184)	21	−4.0 (.497)
1970	18	−12.8 (.524)	28	−3.8 (.446)	40	−6.8 (.624)	20	1.1 (.653)
1971	38	−59.4 (.000)	32	7.1 (.246)	26	−4.2 (.982)	21	−4.6 (.770)
1972	25	10.5 (.601)	46	−2.8 (.601)	12	−11.7 (.804)	13	−5.8 (.978)
1973	11	14.8 (.730)	41	−8.0 (.280)	9	−15.9 (.694)	11	7.3 (.465)
1974	7	27.1 (.711)	39	0.7 (.940)	37	14.1 (.735)	16	10.5 (.462)
1975	20	14.7 (.805)	63	−13.6 (.201)	32	3.6 (.807)	12	2.5 (.857)
1976	27	9.0 (.865)	53	−22.2 (.042)	22	−6.6 (.773)	5	18.9 (.426)
1977	3	−30.9 (.659)	51	−11.4 (.134)	17	−17.4 (.777)	9	−7.1 (.589)
1978	23	−44.8 (.647)	56	−21.1 (.078)	24	−45.1 (.566)	8	−7.7 (.371)
1979	19	−31.1 (.585)	47	−31.2 (.007)	29	−16.5 (.800)	12	−17.5 (.320)
1980	71	−18.5 (.604)	59	−29.5 (.006)	42	3.2 (.861)	17	4.4 (.938)
1981	54	−0.2 (.991)	63	−24.0 (.168)	48	8.0 (.957)	17	−58.6 (.125)
1982	49	31.0 (.276)	54	−22.2 (.196)	–	–	–	–
1983	181	−6.1 (.648)	40	−46.8 (.031)	–	–	–	–
1963 -74	163	−19.8 (.127)	238	−1.1 (.719)	190	−0.9 (.972)	137	0.2 (.663)
1975 -83	447	−4.7 (.876)	486	−23.8 (.000)	214	−6.4 (.828)	80	−14.1 (.222)
1963 -83	610	−8.7 (.534)	724	−16.4 (.003)	404	−3.8 (.862)	217	−5.1 (.389)

Table 4.6

Mean long-run abnormal returns (%) for seasoned equity issuers classified by issuer type (industrials/utilities) and flotation method (firm commitment vs. rights offers), using matching firm returns as the returns benchmark, 1963–1983

The matches are chosen using size- and book-to-market matching and without the five-year pre-event return requirement. The size-matching is done using the equity market value of the issuer. Book-to-market matching is done by first selecting all companies that have an equity market value within 30% of that of the issuer. The company with the closest book-to-market value is chosen as the matching firm. The size rankings are from the month-end prior to the start date of return cumulation. The figures reported below are for $BH \times 100$, where

$$BH = \frac{1}{N}\sum_{i=1}^{N} BH_i = \frac{1}{N}\sum_{i=1}^{N}\left[\prod_{t=\tau_i}^{T_i}(1+R_{it}) - \prod_{t=\tau_i}^{T_i}(1+R_{it}^{m})\right]$$

Numbers in parentheses are number of observations and p-values derived from t-statistics. The p-values are for a two-sided test of the hypothesis that BH equals zero.

	Industrials		Utilities	
Period [$a \pm$ months][a]	Firm-commitment	Rights	Firm-commitment	Rights
(a) 12-months performance intervals during five years prior to issue announcement				
$[a-60, a-49]$	−1.37 (459; .697)	1.48 (26; .890)	−4.47 (578; .000)	−3.33 (75; .026)
$[a-48, a-37]$	−4.19 (486; .233)	−1.22 (30; .953)	−7.96 (597; .000)	−3.70 (76; .039)
$[a-36, a-25]$	−1.10 (513; .738)	0.55 (36; .934)	−6.48 (620; .000)	−4.32 (80; .009)
$[a-24, a-13]$	2.29 (539; .332)	−2.38 (36; .688)	−3.00 (629; .001)	−2.39 (85; .138)
$[a-12, a-1]$	28.11 (569; .000)	8.77 (41; .289)	−1.79 (636; .082)	−2.56 (88; .193)
(b) 12-months performance intervals during five years after issue announcement				
$[a, a+11]$	1.08 (569; .668)	−2.54 (41; .670)	−5.69 (636; .000)	−0.20 (88; .890)
$[a+12, a+23]$	−4.61 (522; .047)	4.77 (40; .414)	−5.36 (356; .000)	−4.03 (68; .067)
$[a+24, a+35]$	−2.67 (474; .353)	−13.67 (38; .044)	−2.67 (213; .146)	−3.60 (34; .484)
$[a+36, a+47]$	−1.40 (430; .597)	−3.40 (33; .615)	−3.34 (152; .133)	−0.94 (25; .884)
$[a+48, a+59]$	1.62 (404; .547)	11.45 (29; .362)	−1.23 (132; .585)	7.25 (16; .072)
(c) One-year to five-year performance intervals				
$[a, a+11]$	1.08 (569; .668)	−2.54 (41; .670)	−5.69 (636; .000)	−0.20 (88; .890)
$[a, a+23]$	−2.23 (569; .505)	−0.26 (41; .978)	−9.76 (636; .000)	−3.67 (88; .081)
$[a, a+35]$	−8.45 (569; .104)	−14.14 (41; .286)	−12.54 (636; .000)	−5.19 (88; .101)
$[a, a+47]$	−9.71 (569; .112)	−12.18 (41; .322)	−14.85 (636; .000)	−7.48 (88; .168)
$[a, a+59]$	−8.70 (569; .192)	−9.13 (41; .611)	−17.50 (636; .000)	−8.20 (88; .266)

[a] a is the announcement month of an issue.

Table 4.7
Mean long-run abnormal returns (%) to public debt issuers classified by issuer type (industrials/utilities) and flotation method (straight/convertible), using matching firm returns as the returns benchmark, 1964–1981

The matches are chosen using size- and book-to-market matching and without the five-year pre-event return requirement. The size-matching is done using the equity market value of the issuer. Book-to-market matching is done by first selecting all companies that have an equity market value within 30% of that of the issuer. The company with the closest book-to-market value is chosen as the matching firm. The size rankings are from the month-end prior to the start date of return cumulation. The figures reported below are for $BH \times 100$, where

$$BH = \frac{1}{N}\sum_{i=1}^{N} BH_i = \frac{1}{N}\sum_{i=1}^{N}\left[\prod_{t=\tau_i}^{T_i}(1+R_{it}) - \prod_{t=\tau_i}^{T_i}(1+R_{it}^{m})\right]$$

Numbers in parentheses are number of observations and p-values derived from t-statistics. The p-values are for a two-sided test of the hypothesis that BH equals zero.

Period	Industrials		Utilities
[a ± months][a]	Straight	Convertible	Straight
(a) 12-months performance intervals during five years prior to issue announcement			
$[a-60, a-49]$	−2.48 (328; .191)	−0.89 (45; .857)	0.03 (194; .977)
$[a-48, a-37]$	−6.55 (335; .000)	10.37 (50; .023)	−2.74 (202; .008)
$[a-36, a-25]$	−1.95 (339; .273)	9.08 (55; .185)	−0.72 (209; .477)
$[a-24, a-13]$	−0.14 (346; .944)	5.92 (57; .230)	−2.47 (211; .015)
$[a-12, a-1]$	0.22 (347; .910)	7.25 (57; .214)	−1.71 (211; .098)
(b) 12-months performance intervals during five years after issue announcement			
$[a, a+11]$	−2.78 (347; .114)	−6.12 (57; .137)	−0.87 (211; .391)
$[a+12, a+23]$	−2.98 (244; .180)	−8.85 (47; .105)	1.01 (102; .621)
$[a+24, a+35]$	3.54 (207; .145)	−9.40 (40; .154)	−7.08 (68; .006)
$[a+36, a+47]$	1.86 (183; .442)	0.97 (32; .894)	1.12 (51; .732)
$[a+48, a+59]$	−1.00 (163; .728)	2.08 (29; .830)	3.29 (45; .376)
(c) One-year to five-year performance intervals			
$[a, a+11]$	−2.78 (347; .114)	−6.12 (57; .137)	−0.87 (211; .391)
$[a, a+23]$	−4.72 (347; .063)	−15.92 (57; .022)	−0.47 (211; .702)
$[a, a+35]$	−3.48 (347; .243)	−17.01 (57; .067)	−3.27 (211; .075)
$[a, a+47]$	−1.42 (347; .698)	−19.26 (57; .059)	−4.66 (211; .053)
$[a, a+59]$	−1.70 (347; .706)	−16.57 (57; .119)	−4.67 (211; .097)

[a] a is the announcement month of an issue.

Panel (a) also documents that utilities making SEOs using a firm commitment underwriting performed negatively relative to the control firms in each of the five years prior to the issue announcement. While statistically significant, the magnitude of the underperformance is relatively small, ranging from −7.96% to −1.79%. Moreover, the underperformance continues in the three to four years after the offering date (see panel (b)), resulting in the overall underperformance of 17.5% (panel (c), last row), consistent with the earlier tables.

Table 4.7 classifies the long-run performance following bond offerings by whether the debt is straight or convertible. As was shown earlier in Table 4.3, panel (a) shows that industrial issuers of convertible debt are associated with a stock price run-up prior to the announcement, while there is no prior run-up for straight debt issuers. More importantly, shifting focus to industrial issuers, panels (b) and (c) fail to indicate any abnormal performance for straight debt offerings on any of the one-to-five-year horizons. However, convertible debt issuers exhibit more negative performance, showing statistically significant (at the 5% level) two-year stock return underperformance. This result differs from that of Spiess and Affleck-Graves (1996). Using a later sample period, they report that firms making debt offerings where straight and convertibles are combined underperform relative to a set of size- and book-to-market matched control firms. Their results appear to be driven by the convertible issues, which are more prevalent in recent years.

In order to determine whether the lack of underperformance in our sample of issuing firms is caused by a confounding effect between size or book-to-market groups, Table 4.8 reports abnormal performance for size and book-to-market quartiles. The table provides no indication that our overall sample results "hide" significant abnormal performance in any of the size or book-to-market quartiles. For industrial issuers, none of the quartiles show significant performance at the 5% level or better. Moreover, utility issuers underperform the matching firms in all the quartiles.

4.5 Performance estimates using factor models

In this section we present results from tests of abnormal long-run performance using the two alternative multifactor asset pricing models explained in section 2.

4.5.1 Construction of portfolio returns

The dependent variables in the factor model regressions are the monthly returns to equal- and value-weighted portfolios of security issuers, respectively. To illustrate, the value-weighted industrial SEO portfolio is constructed as follows: Invest one dollar in the first

Table 4.8
Mean five-year buy-and-hold returns (%) in size and book-to-market quartiles for issuers of seasoned equity offerings and their matching firms on the NYSE and Amex, 1963–1983

The matching firms are selected to have similar size and book-to-market ratios. The size and book-to-market rankings are from the month prior to the start of the return cumulation. Monthly book-to-market rankings in year t are created by dividing the end-of-year book-value from year $t-1$ with monthly market capitalizations for year t. The p-values are computed using the t-statistic for the stock return difference between issuer and matching firm. The buy-and-hold returns are computed using:

$$\frac{1}{N}\sum_{i=1}^{N}\left[\prod_{t=\tau_i}^{T_i}(1+R_{it})-1\right]\times 100$$

The p-values in the column marked $p(t)$ are p-values of the t-statistic using a two-sided test of no difference in average five-year buy-and-hold return for issuer and matching firms.

	Industrial issuers					Utility issuers				
Quartile[a]	N	Issuer	Match	Issuer–Match	$p(t)$	N	Issuer	Match	Issuer–Match	$p(t)$
(a) Size-quartiles										
Q1	28	-8.9	15.7	-24.6	.358	66	55.9	66.8	-10.9	.235
Q2	105	46.0	39.7	6.3	.695	155	42.9	62.4	-19.5	.000
Q3	198	49.2	52.3	-3.1	.776	232	32.1	50.4	-18.3	.000
Q4	279	52.0	68.7	-16.7	.073	271	21.9	36.2	-14.3	.000
(b) Book-to-market-quartiles										
Q1	305	42.5	42.3	0.2	.979	155	29.6	44.3	-14.6	.007
Q2	175	48.0	71.2	-23.2	.072	199	28.7	51.0	-22.2	.000
Q3	99	54.9	67.1	-12.2	.371	195	39.5	55.3	-15.8	.000
Q4	31	65.1	69.3	-4.2	.857	175	32.6	44.6	-11.9	.002

[a]Q1 is the quartile with smallest firms and lowest book-to-market ratios, Q4 is the quartile with largest firms and highest book-to-market ratios.

industrial firm's stock that announces a seasoned equity offering. At the beginning of the first month after the second equity announcement, the portfolio is rebalanced to include the new company using current value-weights. This process is continued as additional firms issue securities, until the first firm reaches its five-year anniversary or an issuing firm's stock in the portfolio is delisted, at which point it is removed and the portfolio is again rebalanced using value-weights.

There is a total of eight such "issuer portfolios" when classified by security type (equity/debt issue) and issuer type (industrial/utility). Similarly, we construct eight "match-portfolios" using the corresponding set of size- and book-to-market matched firms identified in the preceding analysis, and eight "zero-investment portfolios" where the issuer-portfolio is sold short to finance a long position in the match-portfolio.

Table 4.9, Figure 4.1 and Figure 4.2 show the raw returns to portfolios of industrial firms making SEOs (Table 4.9 and Figure 4.1) and bond offerings (Figure 4.2). Referring to the figures, notice that it makes a difference whether the portfolios are value-weighted or equally-weighted. With equal-weights the raw return on issuer and match portfolios move together closely, while with value-weights there are more pronounced differences between the pairs of portfolios. Moreover, returns to the equally-weighted issuer and match portfolios substantially underperform relative to the returns on the equally weighted market index, while the value-weighted issuer and match portfolios outperform the value-weighted market index (except for early in the observation period). As shown in Table 4.9, over the total sample period, the average equal-weighted five-year stock return for issuers is 5.1 percent less than the average five-year stock return for the control sample of matches. This relationship changes with value-weights, where issuers outperform matching firms by 4.5 percent. It is noteworthy that the underperformance is more severe in the earlier 1963–74 subperiod for both equal weighted and value weighted portfolios.[28]

4.5.2 Risk factors and information variables

The Fama and French (1993) model contains three risk factors. The stock market risk factor is the return difference between the value-weighted CRSP-index of all NYSE, Amex and NASDAQ firms and the return on a 1-month Treasury bill, while the other two (SMB and HML) are the returns on zero-investment portfolios. Motivated by the cross-sectional relationship between returns and size (market capitalization) and book-to-market ratio, the SMB and HML portfolios are constructed to be a "size" factor and a "book-to-market"

[28]There is actually overperformance for the issuer value weighted portfolio in the 1975–1983 period.

Table 4.9
Mean buy-and-hold returns (%) over five-year intervals for industrial issuers of seasoned equity classified by portfolio weighting (equal- and value-weights)

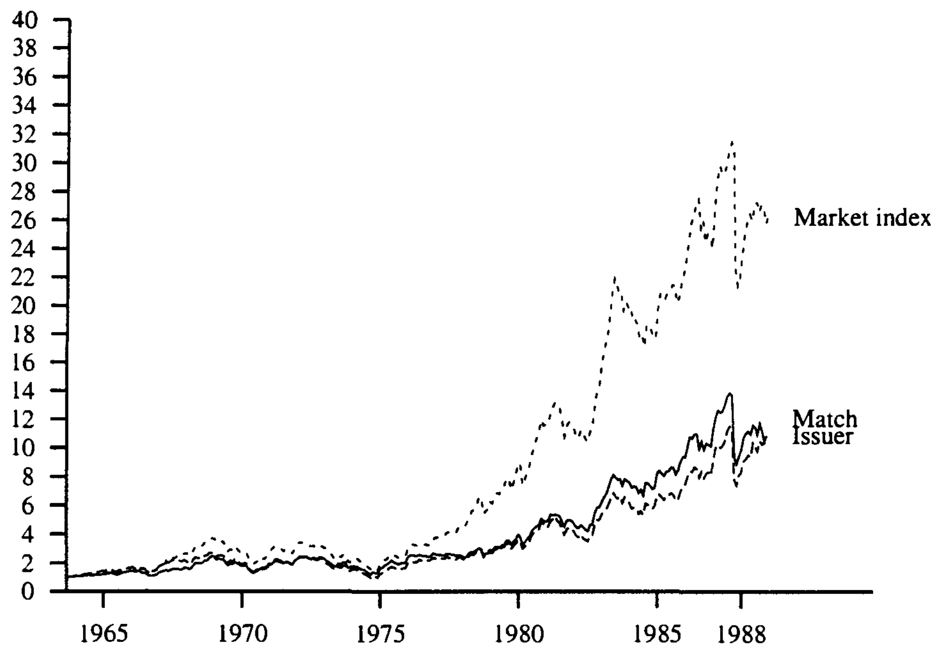
The matching firms are chosen using size- and book-to-market matching and without the five-year pre-event return requirement. The portfolio return in month t is $\omega_t' R_t$, where ω_t is the vector of weights and R_t is the vector of monthly raw returns. The buy-and-hold return in a five-year subperiod starting in month τ is computed using:

$$\left[\prod_{t=\tau}^{\tau+59} (1 + \omega_t' R_t) - 1\right] \times 100$$

		Equal-weighted portfolios		Value-weighted portfolios	
Year	N	Issuers	Matches	Issuers	Matches
1963 – 67	10	119.0	65.3	55.3	20.2
1964 – 68	15	152.9	138.7	61.6	38.4
1965 – 69	24	42.3	60.9	16.6	33.1
1966 – 70	34	8.0	19.9	11.7	31.0
1967 – 71	48	38.9	84.7	38.8	117.0
1968 – 72	65	3.7	42.4	44.8	116.1
1969 – 73	79	-46.7	-32.2	17.9	41.4
1970 – 74	88	-51.6	-31.8	2.6	11.0
1971 – 75	93	-7.8	19.3	61.5	64.0
1972 – 76	92	10.5	26.1	79.7	57.9
1973 – 77	84	5.2	8.0	43.6	18.9
1974 – 78	76	81.5	60.8	75.7	65.2
1975 – 79	73	245.8	188.6	175.0	143.3
1976 – 80	75	189.0	147.8	120.1	90.1
1977 – 81	85	84.2	86.2	67.0	54.6
1978 – 82	100	109.4	130.9	108.1	118.6
1979 – 83	137	149.2	188.8	143.5	159.2
1980 – 84	179	78.1	97.2	130.2	143.0
1981 – 85	209	54.8	90.0	141.8	166.7
1982 – 86	223	86.2	104.0	194.8	178.6
1983 – 87	226	61.7	59.5	184.0	117.8
1984 – 88	188	68.4	40.6	203.2	91.9
1963 – 74		29.7	38.5	42.5	51.2
1975 – 88		112.7	113.4	146.8	126.4
1963 – 88		67.4	72.5	89.9	85.4

Figure 4.1
Value of $1 invested in CRSP indices, industrial issuer-portfolios, and matching-firm portfolios for the sample of seasoned equity offerings

A. Equally-weighted portfolios



B. Value-weighted portfolios

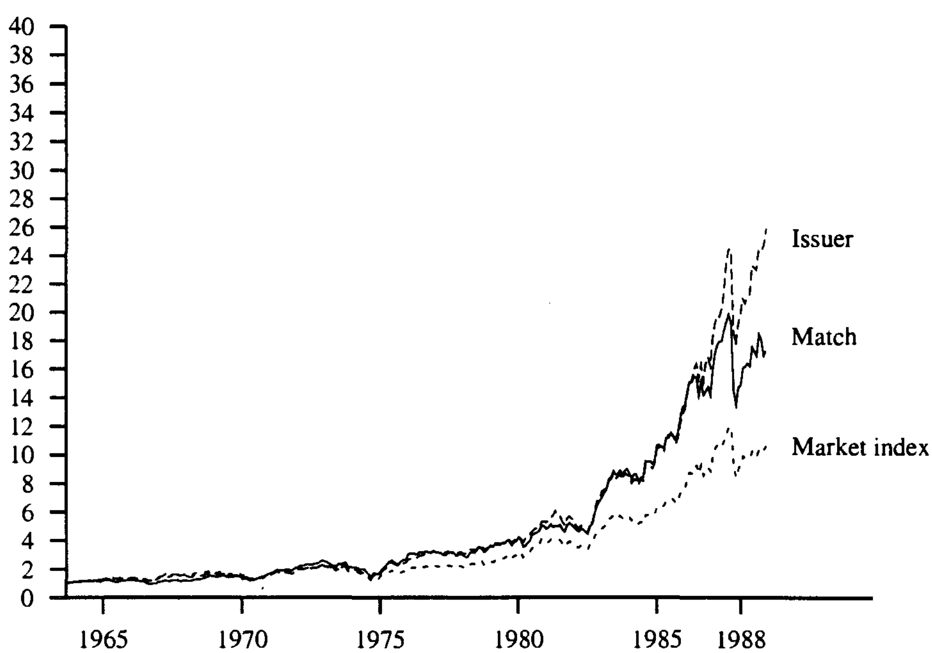
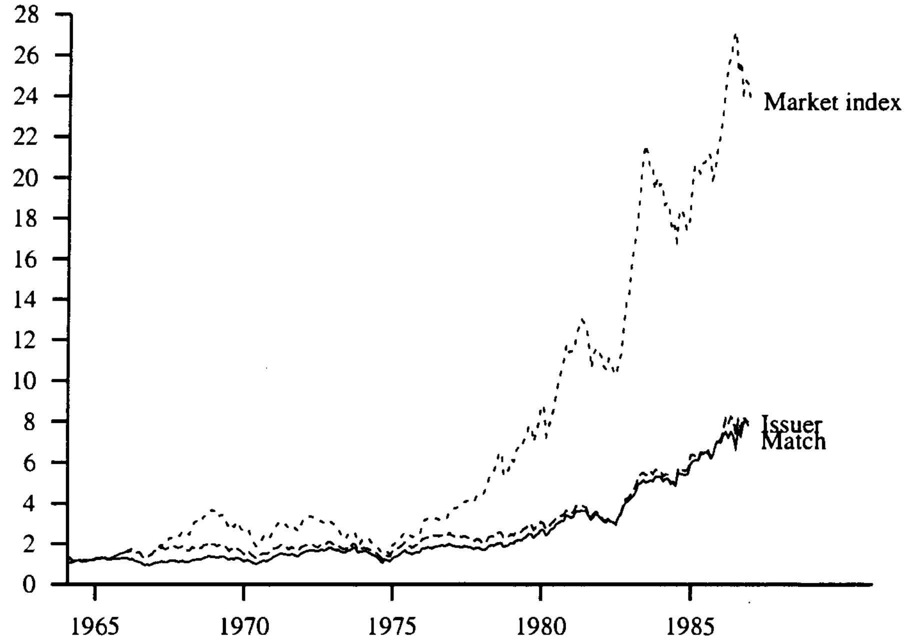
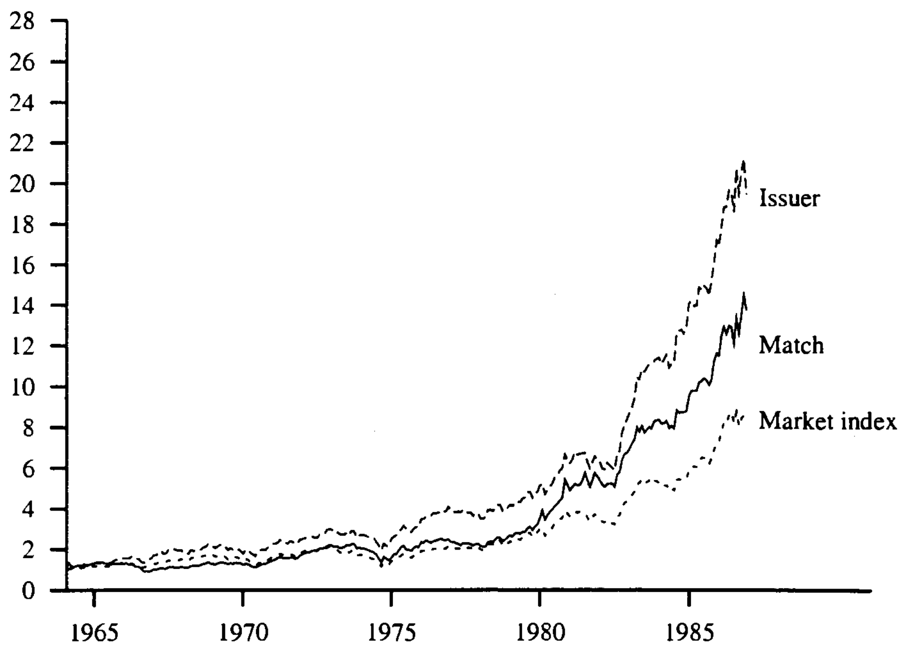


Figure 4.2
Value of $1 invested in CRSP indices, industrial issuer-portfolios, and matching-firm portfolios for the sample of debt offerings

A. Equally-weighted portfolios



B. Value-weighted portfolios



factor respectively. Using size sorted portfolios with the same average book-to-market ratio, the SMB portfolio is long in a portfolio of small stocks and short in a portfolio of big stocks. Similarly, the HML portfolio is constructed from portfolio sorted on book-to-market ratio but with the same average size, and is long in a portfolio of stocks with high book-to-market ratio and short in a portfolio of stocks with low book-to-market ratio. Thus, the SMB factor is constructed to be orthogonal to HML and vice versa. Panel (d) of Table 4.10 shows that the procedure succede in making the SMB and HML factors orthogonal in that the correlation between the factors is only -0.08.

The five risk factors used in the conditional expectation models are described in panel (b) of Table 4.10.[29] The stock market risk factor is the same as in the Fama-French model. The bond market risk factor is the return spread on lo·v-grade and high-grade corporate bonds. We use data from Ibbotson Associates, and high grade bonds are bonds rated AAA by Moody's and low grade bonds are bonds with grading below BAA or that are ungraded. As a term structure factor we follow Ferson and Korajczyk (1995) and use the difference in return on Treasury bonds with 20 years to maturity and 30-day Treasury bills. We also include the real per capita growth rate in personal consumption expenditures for nondurable goods. This proxy should be inversely related to the marginal utility of wealth.[30] Both Ferson and Harvey (1991) and Ferson and Korajczyk (1995) include unanticipated inflation as a risk factor. Let R_{ft} be the yield on a 1-month Treasury bill. Let i_t be inflation during period t. Model inflation as $i_t = i_t^e + i_t^u$ where i_t^e is expected inflation and i_t^u is unexpected inflation. Let realized real return on the 1-month T-bill be $r_{ft} = R_{ft} - i_t$. Consider a regression of real return on a constant and 12 of it's past lagged values:

$$r(\tau) = X(\tau)b(\tau) + u(\tau),$$

where the $(\tau \times 13)$ matrix $X(\tau)$ contains a constant and 12 lagged values of r_{ft} for the period starting at $t = 1$ and ending at $t = \tau$, and $b(\tau)$ is the coefficient vector. Let X_t be a row in $X(\tau)$, and define

$$\hat{R}_t = X_t \hat{b}(t-1),$$

where $\hat{b}(t-1)$ is the estimate of $b(\tau)$ for $\tau = t - 1$. We follow Ferson and Korajczyk (1995) and use $\hat{R}_t$ as an estimate the expected real return. Since unanticipated inflation is expected to be zero by definition, expected inflation for period t may be proxied by $\hat{i}_t^e = r_t - \hat{R}_t$.

[29]See Ferson and Harvey (1991) for a thorough motivation for the inclusion of each of these risk variables.

[30]The time series of the real per capita growth rate in personal consumption expenditures for nondurable goods are from the Federal Reserve Bank of St. Louis' FRED database.

Table 4.10
Summary statistics on risk factors and information variables used in the Fama-French and the conditional factor models of expected returns

SMB is the difference in stock returns between a portfolio of small firms and a portfolio of large firms, where size is measured using market capitalization, HML is the difference in stock returns between portfolios of firms with high and low book-to-market ratio (See Fama and French (1993) for the details on how to compute SMB and HML.) The market portfolio excess return (MR) is computed as the difference between the CRSP value weighted index of NYSE, Amex, and NASDAQ firms and the return on a 1-month Treasury bill. LG−HG is the return difference between high-grade corporate bonds and low-grade corporate bonds from the Ibbotson Associate Corporate Bond Module. ΔRPC is real per capita growth rate of personal consumption of nondurable goods. UI is unanticipated inflation measured as the difference between realized inflation and expected inflation. The expected inflation is modeled by running a regression of real returns (returns on 30-day Treasury bills less inflation) on a constant and 12 of it's lagged values. Tspread is the return difference between Treasury bonds with 20 years to maturity and 30-day Treasury bills. ΔRPC and UI are converted into excess returns using Breeden, Gibbons and Litzenberger (1989) factor-mimicking portfolios. The numbers in panel (a) and (b) are in percentages.

	Mean	Min	Max	Std.
(a) Fama-French model risk factors				
MR	0.38	-22.85	15.99	4.57
SMB	0.27	-10.09	11.05	2.93
HML	0.43	-9.75	8.92	2.56
(b) Risk factors in the conditional factor model				
MR	0.38	-22.85	15.99	4.57
LG−HG	-0.02	-12.24	9.67	2.45
ΔRPC	0.86	-32.26	42.83	8.54
UI	1.24	-47.91	59.12	14.79
Tspread	0.02	-9.37	13.95	3.02
(c) Information variables Z_{t-1}				
Nominal yield on 1-month Treasury bill	0.52	0.20	1.35	0.23
Dividend yield	3.88	2.60	6.03	0.86
3-month/1-month Treasury bill spread	0.04	-0.05	0.26	0.04
BAA/AAA corporate bond yield-spread	0.08	-0.08	0.28	0.06

(d) Correlation matrix for Fama-French model factors

MR	SMB	HML	
MR	1.00		
SMB	0.30	1.00	
HML	-0.40	-0.08	1.00

(e) Correlation matrix for risks in the conditional factor model

	MR	LG−HG	ΔRPC	UI	Tspread
MR	1.00				
LG−HG	0.20	1.00			
ΔRPC	0.71	0.18	1.00		
UI	0.51	0.16	0.52	1.00	
Tspread	0.29	-0.11	0.11	0.16	1.00

Unexpected inflation is measured as the difference between realized inflation and expected inflation.

Since the real per capita growth rate of personal consumption of nondurable goods and unanticipated inflation are not measured as portfolio returns, they are converted into excess returns using Breeden, Gibbons and Litzenberger (1989) factor-mimicking portfolios. Using unanticipated inflation as an example, this involves a regression of the time-series of unanticipated inflation on equally-weighted decile portfolios and the set Z_{t-1} of predetermined information variables. The regression coefficients on the decile portfolios are used to form portfolio weights by dividing each portfolio's coefficient by the sum of all the coefficients. The factor mimicking portfolio is the time-series of return constructed from the decile portfolios using the weights constructed from the regression coefficients. Simple correlations of the risk factors are reported in panel (e) of Table 4.10.

The predetermined information variables that we use have all been shown in the literature to have predictive power for future stock and bond returns.[31] Panel (b) of Table 4.10 summarizes the information variables. The nominal one-month and three-month Treasury bill returns are from the CRSP-Fama files. The monthly dividend yield is computed as the aggregate cash dividends paid in the current month and the previous eleven months divided by the value of the index (without dividends reinvested) in the month before the start month of dividend summation. As a proxy for the slope of the near-maturity term structure, we follow Ferson and Harvey (1991) and use the difference in monthly returns on a 3-month Treasury bill and a 1-month Treasury bill. To capture corporate default risk, we use the Ibbotson Associates' yield spread between corporate bonds rated BAA and corporate bonds rated AAA. The short (one-month T-bill) rate is included as a current information variable since it is highly correlated with expected inflation, which in turn is negatively correlated with future expected stock returns (Fama and Schwert (1977)). There is also evidence that T-bill rates are related to the conditional second moments of stock returns (e.g., Campbell (1987)). Furthermore, both the dividend yield and the near-maturity term structure slope are shown in prior studies to have predictive power for future stock returns and are consequently included as information variables (e.g., Fama and French (1988), Ferson and Harvey (1991)). Finally, the default-related yield spread variable captures changes in risk premiums and in market expectations of losses on risky bonds, where both effects are driven by changes in economic conditions (e.g., business cycles). This variable is also documented to be positively related to future stock returns (e.g., Keim and Stambaugh (1986), Fama

[31] See Ferson and Harvey (1991), Evans (1994), Ferson and Korajczyk (1995), Ferson and Schadt (1996), and Chen and Knez (1996) among others.

and French (1989)).

4.5.3 Risk adjustment using the Fama-French three-factor model

Under the null hypothesis of no abnormal performance for the constructed portfolios, the constant term in equation (4.1) should be indistinguishable from zero. Tests of this prediction are found in Table 4.11. The estimated constant term is found in the second column of this table marked $\hat{\alpha}^{\text{ff}}$. Looking at industrial firms in panel (a), the alphas are significantly negative for equally-weighted matching firms but not for issuers. This suggests that only the matching firms underperform relative to the three-factor model when aggregated into portfolios with equal weights. When using value-weights, however, the alpha estimates indicate that both issuers and matching firms signifiantly *over*perform by approximately .3 percent per month.

Statistically significant values of $\hat{\alpha}^{\text{ff}}$ for the matching firm portfolios potentially reflect model misspecification or market mispricing. Model misspecification can be caused by omitted factors and/or violation of the assumption of constant expected returns. Suppose that the model omits factor F_t^* and that the issuing firms and their matches have similar exposures to this factor. In this case, market efficiency implies that a self-financing investment strategy of shorting issuers and going long in the matching firms generates zero values of $\hat{\alpha}^{\text{ff}}$. As shown in Table 4.11 for industrial issuers, the alpha values for the difference portfolio (zero investment strategy) are insignificantly different from zero at conventional levels.

Turning to utility issuers, there is no evidence in panel (b) of Table 4.11 of significant alphas for either issuers or matching firms. However, the difference portfolios (long in matching firms and short in issuers) generate significant *positive* alphas. These results are largely consistent with the evidence of underperformance found earlier for utility issuers using the matching-firm procedure (Table 4.4). Thus, the Fama-French model is not able to explain the negative return difference between utility SEOs and utility matching firms.

Turning to bond issuers, panel (a) of Table 4.12 presents evidence of statistically significant alphas for value-weighted portfolios of industrial issuers and matching firms, similar to the SEO issuer evidence. Again, for industrial issuers, the alphas for the difference portfolios (long in matching firms and short in issuers) are not significant. Furthermore, the pattern of abnormal returns for utility bond issuers is similar to the pattern observed earlier for utility SEOs in Table 4.11. That is, utility issuers seems to underperform relative to their benchmark firms. In panel (b) of Table 4.12 this show up as a negative and significant (at the 6% level) alpha for the equally-weighted issuer portfolio and positive alphas for the

Table 4.11
Jensen's alpha for SEO issuer and match portfolios using the Fama-French three-factor model as an expected return benchmark

The risk factors are the difference in return between small and large firms (SMB), where size is measured using market capitalization, the return difference between high and low book-to-market firms (HML), and the excess return on a value weighted index (MR). See Fama and French (1993) for the details on how to compute SMB and HML. The excess return is computed as the difference between the CRSP value weighted index for NYSE, Amex, and NASDAQ firms and the return on a 1-month Treasury bill. The Fama and French (1993) three-factor model is:

$$r_{pt} = \alpha_p^{ff} + \beta_{p1}\text{MR}_t + \beta_{p2}\text{SMB}_t + \beta_{p3}\text{HML}_t + e_{pt}^{ff}$$

where r_{pt} is portfolio return. The coefficients are estimated using OLS. Standard errors are computed using the heteroskedasticity consistent estimator of White (1980). The numbers in parentheses are p-values.

Portfolio[a]	$\hat{\alpha}^{ff}$	MR	SMB	HML	R^2	$F(Z_{t-1})$
(a) Industrials						
Issuing firms, EW	−0.14 (.174)	1.15 (.000)	0.61 (.000)	−0.04 (.466)	0.927	0.031
Matching firms, EW	−0.18 (.058)	1.08 (.000)	0.63 (.000)	0.02 (.523)	0.930	0.314
Issuing firms, VW	0.37 (.002)	1.02 (.000)	−0.09 (.022)	−0.06 (.246)	0.853	0.004
Matching firms, VW	0.24 (.035)	1.04 (.000)	−0.01 (.819)	−0.11 (.023)	0.866	0.014
Match-issuer, EW	−0.03 (.767)	−0.07 (.004)	0.02 (.645)	0.06 (.187)	0.028	0.269
Match-issuer, VW	−0.13 (.387)	0.03 (.514)	0.08 (.144)	−0.05 (.446)	0.009	0.009
(b) Utilities						
Issuing firms, EW	−0.20 (.215)	0.74 (.000)	−0.15 (.032)	0.59 (.000)	0.536	0.000
Matching firms, EW	0.08 (.570)	0.81 (.000)	−0.12 (.032)	0.42 (.000)	0.629	0.002
Issuing firms, VW	−0.13 (.465)	0.77 (.000)	−0.24 (.000)	0.55 (.000)	0.511	0.002
Matching firms, VW	0.29 (.115)	0.86 (.000)	−0.27 (.000)	0.31 (.000)	0.570	0.030
Match-issuer, EW	0.28 (.046)	0.07 (.096)	0.03 (.590)	−0.17 (.038)	0.071	0.000
Match-issuer, VW	0.42 (.024)	0.09 (.092)	−0.03 (.673)	−0.24 (.030)	0.073	0.000

[a]Portfolios are either equal-weighted ('EW') or value-weighted ('VW'). The 'Match-issuer' is a zero-investment portfolio with a long position in the matching firms and a short position in the issuing firms.

Table 4.12
Jensen's alpha for bond issuer and match portfolios using the Fama-French three-factor model as an expected return benchmark

The risk factors are the difference in return between small and big firms (SMB), where size is measured using market capitalization, the return difference between high and low book-to-market firms (HML), and the excess return on a value weighted index (MR). See Fama and French (1993) for the details on how to compute SMB and HML. The excess return is computed as the difference between the CRSP value weighted index for NYSE, Amex, and NASDAQ firms and the return on a 1-month Treasury bill. The Fama and French (1993) three-factor model is:

$$r_{pt} = \alpha_p^{ff} + \beta_{p1}\text{MR}_t + \beta_{p2}\text{SMB}_t + \beta_{p3}\text{HML}_t + e_{pt}^{ff}$$

where r_{pt} is portfolio return. The coefficients are estimated using OLS. Standard errors are computed using the heteroskedasticity consistent estimator of White (1980). The numbers in parentheses are p-values.

Portfolio[a]	$\hat{\alpha}^{ff}$	MR	SMB	HML	R^2	$F(Z_{t-1})$
(a) Industrials						
Issuing firms, EW	−0.15 (.264)	1.15 (.000)	0.03 (.447)	0.22 (.000)	0.836	0.730
Matching firms, EW	−0.09 (.368)	1.08 (.000)	−0.06 (.136)	0.15 (.000)	0.890	0.587
Issuing firms, VW	0.30 (.012)	1.06 (.000)	−0.20 (.000)	0.17 (.000)	0.824	0.477
Matching firms, VW	0.43 (.003)	1.03 (.000)	−0.34 (.000)	−0.16 (.008)	0.803	0.703
Match-issuer, EW	0.06 (.673)	−0.07 (.036)	−0.09 (.052)	−0.07 (.254)	0.028	0.966
Match-issuer, VW	0.13 (.483)	−0.03 (.478)	−0.14 (.021)	−0.32 (.000)	0.086	0.887
(b) Utilities						
Issuing firms, EW	−0.36 (.060)	0.84 (.000)	−0.13 (.072)	0.55 (.000)	0.547	0.005
Matching firms, EW	−0.04 (.852)	0.84 (.000)	−0.33 (.000)	0.38 (.000)	0.513	0.170
Issuing firms, VW	−0.19 (.304)	0.83 (.000)	−0.20 (.004)	0.56 (.000)	0.540	0.006
Matching firms, VW	0.08 (.728)	0.85 (.000)	−0.40 (.000)	0.33 (.000)	0.458	0.165
Match-issuer, EW	0.32 (.034)	0.00 (.977)	−0.20 (.003)	−0.16 (.054)	0.074	0.000
Match-issuer, VW	0.28 (.094)	0.01 (.809)	−0.21 (.005)	−0.22 (.013)	0.083	0.001

[a]Portfolios are either equal-weighted ('EW') or value-weighted ('VW'). The 'Match-issuer' is a zero-investment portfolio with a long position in the matching firms and a short position in the issuing firms.

both the equal- and value weighted difference portfolios.

In sum, the evidence in this section indicates some abnormal performance in the utility samples. However, the abnormal performance may very well be caused by a misspecification of the factor model used to control for risk. This suspicion is supported by the F-tests in the last column of table 4.11 and 4.12. These columns report the p-values of a null hypothesis that the betas in the Fama-French model is not time-varying. The test is constructed by augmenting the Fama-French model with the following model for time-variation in factor betas:

$$\beta_{pj} = b_{p0j} + b'_{p1j} Z_{t-1}, \quad j = 1, 2, 3.$$

The F-test is for the null hypothesis that all the elements in the L-vector b_{p1j} is zero. For the utility portfolios, the null is rejected for ten out of twelve portfolios. This support that the Fama-French three-factor model with constant betas is misspecified. Further evidence of misspecifiaction appear in separate regressions of size-decile portfolio returns against the Fama-French factors, which produces significant constant terms for some portfolios (not reported in tables).

We now turn to a different specification of the factor model, namely one which allows expected returns to be time varying based on various macroeconomic risks identified in the empirical literature on arbitrage pricing.

4.5.4 Risk adjustment using a conditional four-factor model

The null hypothesis underlying the analysis in this section is that α_p in equation (4.2) is zero *conditional* on the current information variables in Z_{t-1}.

In order to check the robustness of the model, we first estimate alphas using size decile portfolios. Under the null hypothesis that the model captures all priced risk factors, the alphas for each of the decile portfolios should be indistinguishable from zero. Table 4.13 contains the coefficients from these estimations. Only one of the portfolios (Decile 8) has an alpha that is different from zero at the 10% level, and none of the decile portfolios have alphas that are different from zero at the 5% level of statistical significance. Column 3 through 7 report the estimated betas at the mean level of the information variables, that is

$$\hat{\beta}_{pt-1} = \hat{b}_{i0} + \hat{b}_{i1} \overline{Z}_{t-1},$$

where $\overline{Z}_{t-1}$ is a L-vector containing the time-series mean of the information variables. The market index (MR), the real per capita growth rate of personal consumption of nondurables

Table 4.13
Jensen's alpha for portfolios of equally-weighted size deciles using a multi-factor asset pricing model as an expected return benchmark

The market portfolio excess return (MR) is computed as the difference between the CRSP value weighted index of NYSE, Amex, and NASDAQ firms and the return on a 1-month Treasury bill. LG−HG is the return difference between high-grade corporate bonds and low-grade corporate bonds from the Ibbotson Associate Corporate Bond Module. ΔRPC is real per capita growth rate of personal consumption of nondurable goods. UI is unanticipated inflation measured as the difference between realized inflation and expected inflation. The expected inflation is modeled by running a regression of real returns (returns on 30-day Treasury bills less inflation) on a constant and 12 of it's lagged values. Tspread is the return difference between Treasury bonds with 20 years to maturity and 30-day Treasury bills. ΔRPC and UI are converted into excess returns using Breeden, Gibbons and Litzenberger (1989) factor-mimicking portfolios. The information variables in Z_{t-1} are listed in table 4.10. The model

$$r_{it} = \alpha_p + b_{i0}' r_{mt} + b_{i1}'(Z_{t-1} \otimes r_{mt}) + e_{it}$$

is estimated using OLS. Standard errors are computed using the heteroskedasticity consistent estimator of White (1980). The numbers in parentheses are p-values.

		Betas at mean Z_{t-1}						
Portfolio[a]	$\hat{\alpha}_p$	MR	LH−HG	ΔRPC	UI	Tspread	R^2	$F(Z_{t-1})$
Decile 1	0.10 (0.643)	0.48 (0.000)	0.17 (0.134)	0.44 (0.000)	0.11 (0.000)	0.04 (0.631)	0.000	0.787
Decile 2	0.06 (0.632)	0.58 (0.000)	0.08 (0.578)	0.33 (0.000)	0.11 (0.000)	0.01 (0.772)	0.003	0.909
Decile 3	−0.00 (0.979)	0.55 (0.000)	0.07 (0.071)	0.36 (0.000)	0.10 (0.000)	0.00 (0.920)	0.001	0.966
Decile 4	0.07 (0.329)	0.66 (0.000)	0.03 (0.432)	0.30 (0.000)	0.07 (0.000)	0.05 (0.623)	0.339	0.963
Decile 5	−0.01 (0.893)	0.75 (0.000)	0.07 (0.419)	0.17 (0.000)	0.11 (0.000)	−0.02 (0.732)	0.676	0.952
Decile 6	0.06 (0.255)	0.74 (0.000)	0.04 (0.313)	0.18 (0.000)	0.08 (0.000)	0.03 (0.628)	0.032	0.966
Decile 7	0.05 (0.277)	0.86 (0.000)	−0.02 (0.802)	0.10 (0.000)	0.08 (0.000)	−0.00 (0.802)	0.006	0.974
Decile 8	0.10 (0.075)	0.91 (0.000)	0.04 (0.527)	0.05 (0.000)	0.06 (0.000)	0.05 (0.023)	0.008	0.963
Decile 9	0.05 (0.246)	0.97 (0.000)	−0.02 (0.282)	−0.03 (0.000)	0.06 (0.000)	0.08 (0.000)	0.000	0.981
Decile 10	0.01 (0.706)	1.09 (0.000)	−0.02 (0.075)	−0.06 (0.000)	−0.03 (0.000)	0.04 (0.006)	0.000	0.982

[a]Decile 1 is the portfolio with the smallest equity capitalization stocks while Decile 10 contain the largest equity capitalization stocks.

(ΔRPC), and unanticipated inflation (UI) are all associated with statistically significant betas for all the portfolios. The measure of corporate default risk (LG−HG) and the term-structure factor (Tspread) also appears as a relevant risk factor for some of the portfolios. However, note that the reported p-values are for the joint hypothesis of relevance of the factor (the "average" beta $\hat{b}_{i0}$ is different from zero) and the time-varyation in the factor loadings (the effect of $\hat{b}_{i1}' Z_{t-1}$ is different from zero). The last column in Table 4.10 reports p-values for the null hypothesis that betas are constant. For all but two portfolios, this hypothesis is rejected.

Tables 4.14 and 4.15 present the conditional factor model results for the issuer-, matching firm-, and zero-investment portfolios for the samples of SEOs and bond offerings. The tables show that the alpha estimates tend to be smaller and less significant than the alpha estimates in Fama-French three-factor model reported above.

Panel (a) of Table 4.14 presents the results for industrial SEOs. Except for the value-weighted issuer portfolio, none of the alphas are distinguishable from zero. The estimate $\hat{\alpha}_p$ for the value-weighted issuer portfolios are positive and significant. However, as with the Fama-French regressions (Table 4.11), the alpha estimates for the value-weighted zero-investment portfolio are again statistically indistinguishable from zero.

Panel (b) of Table 4.14 contains the corresponding results for SEOs by utilities. The table shows insignificant alphas for both equal- and value-weighted issuer portfolios. However, the utility matching firm portfolios have statistically significant alphas, implying *over-performance* relative to the benchmark. Also the difference (match−issuer) portfolios for utilities produce positive and significant alphas. However, what earlier appeared as underperformance of issuers relative to the matching firms now shows up in Table 4.14 as overperformance of the matching firm portfolios relative to the benchmark.

Table 4.15 reports the corresponding alpha values for portfolios of debt issuers. The table shows significant and positive alphas for the value-weighted portfolios of industrial issuers and matched firms. However, none of the alphas for the difference portfolios (match−issuer) are statistically distinguishable from zero. For utility issuers there are no clear pattern, only one alpha is significantly different from zero at the 5% level. Given the number of alphas that are estimated, this could easily occur by chance.

Overall, the evidence presented in this section is consistent with the hypothesis that market reactions to security issue announcements are unbiased. That is, using a time-varying expected return framework, we fail to document significant abnormal returns from a self financing investment strategy involving shorting issuing firms and going long in matching

Table 4.14
Jensen's alpha for SEO issuer and match portfolios using a conditional multi-factor asset pricing model as an expected return benchmark

The market portfolio excess return (MR) is computed as the difference between the CRSP value weighted index of NYSE, Amex, and NASDAQ firms and the return on a 1-month Treasury bill. LG−HG is the return difference between high-grade corporate bonds and low-grade corporate bonds from the Ibbotson Associate Corporate Bond Module. ΔRPC is real per capita growth rate of personal consumption of nondurable goods. UI is unanticipated inflation measured as the difference between realized inflation and expected inflation. The expected inflation is modeled by running a regression of real returns (returns on 30-day Treasury bills less inflation) on a constant and 12 of it's lagged values. Tspread is the return difference between Treasury bonds with 20 years to maturity and 30-day Treasury bills. ΔRPC and UI are converted into excess returns using Breeden, Gibbons and Litzenberger (1989) factor-mimicking portfolios. The information variables in Z_{t-1} are listed in table 4.10. The model

$$r_{it} = \alpha_p + b_{i0}' r_{mt} + b_{i1}'(Z_{t-1} \otimes r_{mt}) + e_{it}$$

is estimated using OLS. Standard errors are computed using the heteroskedasticity consistent estimator of White (1980). The numbers in parentheses are p-values.

		Betas at mean Z_{t-1}						
Portfolio[a]	$\hat{\alpha}_p$	MR	LH−HG	ΔRPC	UI	Tspread	R^2	$F(Z_{t-1})$
(a) Industrials								
Issuing firms, EW	−0.06	0.99	0.05	0.14	0.08	−0.10	0.924	0.406
	(0.549)	(0.000)	(0.205)	(0.000)	(0.000)	(0.066)		
Matching firms, EW	−0.11	0.93	0.01	0.16	0.06	−0.09	0.908	0.996
	(0.318)	(0.000)	(0.808)	(0.000)	(0.000)	(0.169)		
Issuing firms, VW	0.30	1.06	−0.02	−0.05	0.02	−0.06	0.868	0.000
	(0.003)	(0.000)	(0.228)	(0.068)	(0.011)	(0.009)		
Matching firms, VW	0.08	1.11	0.00	−0.01	−0.01	0.03	0.872	0.007
	(0.448)	(0.000)	(0.385)	(0.087)	(0.005)	(0.907)		
Match-issuer, EW	−0.05	−0.06	−0.04	0.02	−0.02	0.01	0.042	0.425
	(0.693)	(0.218)	(0.136)	(0.273)	(0.095)	(0.198)		
Match-issuer, VW	−0.21	0.05	0.03	0.04	−0.02	0.09	0.079	0.006
	(0.152)	(0.041)	(0.355)	(0.063)	(0.270)	(0.022)		
(b) Utilities								
Issuing firms, EW	0.02	0.41	0.06	−0.02	0.05	0.45	0.597	0.000
	(0.920)	(0.000)	(0.503)	(0.164)	(0.000)	(0.000)		
Matching firms, EW	0.29	0.55	0.10	−0.02	0.05	0.28	0.631	0.065
	(0.055)	(0.000)	(0.047)	(0.112)	(0.000)	(0.000)		
Issuing firms, VW	0.08	0.47	0.03	−0.04	0.04	0.47	0.566	0.000
	(0.632)	(0.000)	(0.499)	(0.124)	(0.000)	(0.000)		
Matching firms, VW	0.49	0.69	0.13	−0.08	0.03	0.27	0.594	0.001
	(0.006)	(0.000)	(0.004)	(0.005)	(0.004)	(0.001)		
Match-issuer, EW	0.28	0.15	0.04	−0.01	0.00	−0.17	0.330	0.000
	(0.013)	(0.000)	(0.185)	(0.744)	(0.003)	(0.000)		
Match-issuer, VW	0.41	0.22	0.11	−0.03	−0.01	−0.20	0.332	0.000
	(0.005)	(0.000)	(0.069)	(0.275)	(0.002)	(0.000)		

[a]Portfolios are either equal-weighted ('EW') or value-weighted ('VW'). The 'Match-issuer' is a zero-investment portfolio with a long position in the matching firms and a short position in the issuing firms.

Table 4.15
Jensen's alpha for bond issuer and match portfolios using a conditional multi-factor asset pricing model as an expected return benchmark

The market portfolio excess return (MR) is computed as the difference between the CRSP value weighted index of NYSE, Amex, and NASDAQ firms and the return on a 1-month Treasury bill. LG−HG is the return difference between high-grade corporate bonds and low-grade corporate bonds from the Ibbotson Associate Corporate Bond Module. ΔRPC is real per capita growth rate of personal consumption of nondurable goods. UI is unanticipated inflation measured as the difference between realized inflation and expected inflation. The expected inflation is modeled by running a regression of real returns (returns on 30-day Treasury bills less inflation) on a constant and 12 of it's lagged values. Tspread is the return difference between Treasury bonds with 20 years to maturity and 30-day Treasury bills. ΔRPC and UI are converted into excess returns using Breeden, Gibbons and Litzenberger (1989) factor-mimicking portfolios. The information variables in Z_{t-1} are listed in table 4.10. The model

$$r_{it} = \alpha_p + b_{i0}' r_{mt} + b_{i1}'(Z_{t-1} \otimes r_{mt}) + e_{it}$$

is estimated using OLS. Standard errors are computed using the heteroskedasticity consistent estimator of White (1980). The numbers in parentheses are p-values.

		Betas at mean Z_{t-1}						
Portfolio[a]	$\hat{\alpha}_p$	MR	LH−HG	ΔRPC	UI	Tspread	R^2	$F(Z_{t-1})$
(a) Industrials								
Issuing firms, EW	−0.06 (0.678)	1.05 (0.000)	0.02 (0.797)	0.02 (0.910)	0.02 (0.381)	0.01 (0.897)	0.824	0.880
Matching firms, EW	−0.04 (0.694)	1.04 (0.000)	0.01 (0.540)	−0.00 (0.726)	0.00 (0.583)	0.00 (0.431)	0.883	0.453
Issuing firms, VW	0.30 (0.034)	1.05 (0.000)	−0.06 (0.912)	−0.04 (0.710)	−0.01 (0.606)	−0.01 (0.975)	0.803	0.904
Matching firms, VW	0.35 (0.015)	1.19 (0.000)	0.02 (0.102)	−0.08 (0.023)	−0.07 (0.000)	−0.02 (0.619)	0.808	0.647
Match-issuer, EW	0.02 (0.901)	−0.01 (0.897)	−0.01 (0.539)	−0.03 (0.672)	−0.02 (0.670)	−0.01 (0.493)	-0.001	0.940
Match-issuer, VW	0.05 (0.792)	0.15 (0.192)	0.07 (0.102)	−0.04 (0.668)	−0.05 (0.006)	−0.01 (0.860)	0.060	0.649
(b) Utilities								
Issuing firms, EW	−0.20 (0.291)	0.42 (0.000)	−0.02 (0.236)	−0.01 (0.227)	0.07 (0.000)	0.42 (0.000)	0.608	0.000
Matching firms, EW	0.12 (0.545)	0.59 (0.000)	0.02 (0.667)	−0.09 (0.012)	0.04 (0.068)	0.46 (0.000)	0.533	0.256
Issuing firms, VW	−0.02 (0.937)	0.44 (0.000)	−0.04 (0.360)	−0.03 (0.241)	0.06 (0.000)	0.45 (0.000)	0.591	0.000
Matching firms, VW	0.22 (0.333)	0.64 (0.000)	−0.00 (0.532)	−0.11 (0.004)	0.02 (0.137)	0.47 (0.000)	0.482	0.131
Match-issuer, EW	0.33 (0.032)	0.17 (0.001)	0.04 (0.222)	−0.08 (0.001)	−0.03 (0.000)	0.04 (0.319)	0.201	0.000
Match-issuer, VW	0.23 (0.158)	0.19 (0.001)	0.04 (0.157)	−0.08 (0.001)	−0.03 (0.000)	0.02 (0.503)	0.184	0.000

[a] Portfolios are either equal-weighted ('EW') or value-weighted ('VW'). The 'Match-issuer' is a zero-investment portfolio with a long position in the matching firms and a short position in the issuing firms.

firms.[32]

4.6 Conclusions

Capital market participants react to security issue announcements by revaluing the issuer's stock price. This revaluation depends in part on the market's perception of the issuing firm's objectives and in part on the nature of the information asymmetry between investors and the firm concerning the true value of its securities. As surveyed Eckbo and Masulis (1995), substantial empirical research has established that the market reaction to seasoned security offerings is swift and consistent with the hypothesis that investors are concerned with adverse selection.

However, recent studies of the long-run performance following security offers raise an important question of whether this market reaction is unbiased. In particular, Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995) find that firms making seasoned equity offers have stock returns which systematically underperform a control group of non-issuing firms over a long horizon; thus generating what has been termed the "new issues puzzle".

The results of this study raise doubts about the econometric foundation of the "new issues puzzle". First, we draw on recent asset pricing econometrics to produce time-varying expected return benchmarks for estimating long-run abnormal stock price performance. This approach also alleviates a concern that risk factors constructed using broad-based stock portfolios, such as those in the Fama-French model, are "contaminated" as they themselves contain a large proportion of issuing firms. More specifically, Loughran and Ritter (1999) argue that portfolios of small growth stocks (i.e., small size and low book-to-market) contain a disproportionately large number of issuers and therefore should not be used as a benchmark for abnormal return estimation. Since our factors are based on macroeconomic variables, this criticism does not apply to our expected stock returns or abnormal returns estimates.

Second, we examine a range of issue and issuer characteristics not presented in earlier research on security offerings, including convertible and straight debt, in addition to examining equity offering flotation methods, and public utility issuers as well as industrial issuers. Since characterizations of the nature of the market underreactions to security of-

[32]This conclusion receives indirect support from a recent study by Brous, Datar and Kini (1998) who examine corporate earnings following SEOs. They report no evidence of significant abnormal earnings in the five-year period following SEOs.

ferings ought to be consistent across these various issue characteristics, this study furthers our understanding of potential underperformance effects and thus reinforces our overall conclusions.

Before making methodological improvements to the conventional procedures, we establish that our industrial sample of seasoned equity offers produces significant underperformance similar to Loughran and Ritter (1995) when using their estimation techniques. We also document that utility issuers of seasoned equity offers exhibit significant underperformance, albeit of a lesser magnitude. We then show that, for industrial issuers, this long-run underperformance is eliminated when employing either matching firm techniques sorted on size and book-to-market, or time-varying multifactor expected return benchmarks. This conclusion is also robust to the portfolio weighting procedure and the equity offering flotation method chosen. Moreover, we find statistically insignificant long-run abnormal performance following debt issues. Overall, the evidence in this paper is consistent with the hypothesis that the typical market reaction to seasoned security offerings announcements is unbiased.

Chapter 5

Equity Flotation Method Choice

Eckbo and Masulis (1992) develop a model framework for firms equity flotation method choice under adverse selection which captures a number of empirical regularities. This chapter formalizes and extends the Eckbo-Masulis model to include private placements as an alternative to rights and underwritten offerings, and to allow the possibility of issue rejection by either the private placement investor or the underwriter. This extension produces a set of empirical implications that are tested on the Oslo Stock Exchange where all methods used to issue seasoned equity is in fact covered by the model. The model implications are supported by the empirical results. The chapter also examines competing theories for the announcement period returns. For example, there is no evidence that the positive market reaction to private placements is due to increased monitoring by large shareholders.

5.1 Introduction

A company that wishes to raise new equity must decide on the flotation method. For public equity offerings, flotation methods commonly used internationally include firm commitment underwritten offering, rights offering with standby underwriting, and uninsured rights offering.[1] If companies rank flotation methods using direct costs only (underwriter fees, legal fees, mailing costs etc.), we should see strong preferences for the relatively inexpensive rights

[1]In rights offers, current shareholders are given tradeable short-term warrants (rights) that prevent wealth loss due to dilution either by allowing shareholders to maintain their ownership level by purchase new shares at a discount, or obtain cash compensation by selling the warrants. Standby rights offers differ from uninsured rights offers in that an underwriter (normally an investment bank) insure the issuer against offer failure by committing to by a certain fraction of the issue. In a firm commitment underwritten offer, an investment bank acquires the whole issue.

offer method.[2] However, international evidence suggest that issuers prefer to raise equity using the services of an underwriter, either in the form of standby underwriting or using firm commitment contracts—despite the fact that these methods are substantially more costly in terms of direct costs.[3] This puzzle is commonly referred to as the "rights offer paradox".

The resolution of the rights offer paradox involves identifying indirect costs of rights, or managerial benefits associated with underwriting, that may explain the observed trend away from rights issues. Eckbo and Masulis (1992) develop theoretically and empirically the hypothesis that observed flotation method choices in the U.S. are driven by adverse selection in issue markets. Their model extends the basic Myers and Majluf (1984) adverse selection argument (which focuses on direct public sales only) to a situation where the firm may sell some or all of the issue to current shareholders and/or to an underwriter. Shareholder takeup and sale to the public via an underwriter are modeled as substitute mechanisms for reducing wealth transfers between current shareholders and outside investors. Assuming exogenous shareholder demand for the issue, and imperfect quality certification by the underwriter, Eckbo and Masulis (1992) characterize an equilibrium where only firms with a high expected shareholder chose uninsured rights. Moreover, in this equilibrium, the average issuer value in uninsured rights offers exceeds that in underwritten offers, and the market reaction to the issue announcement is predicted to be inversely related to current shareholder takeup.

These predictions are found to be consistent with the evidence in several empirical studies. Using samples of U.S. and Norwegian seasoned equity offerings, respectively, Eckbo and Masulis (1992) and Bøhren, Eckbo and Michalsen (1997) find that the average announcement period abnormal return for underwritten offerings is less favorable than for uninsured rights offerings. The later study also documents that the probability that the issuer underwrites a rights offer is inversely related to expected shareholder takeup, as predicted. Moreover, Singh (1997) present additional evidence that adverse selection is greater among standby rights offerings with low current shareholder takeup than for offerings with high takeup.

[2] Eckbo and Masulis (1992) find that the average cost in percent of total offering proceeds is 6% for firm commitment offerings, 4% for standby rights, and 1.5% for uninsured rights offerings by U.S. industrial firms. The costs for public utilities are somewhat lower. See also Eckbo and Masulis (1995) for a survey.

[3] Firms listed on the New York Stock Exchange (NYSE) or the American Stock Exchange (AMEX) have over the post 1981 period almost exclusively have seasoned equity using firm commitment underwritten offerings (See for example Eckbo and Masulis (1992)). In several other countries there is a trend towards the exclusive use of underwritten offerings. See Bøhren, Eckbo and Michalsen (1997) for Norway and Kato and Schallheim (1995) for Japan.

This chapter contributes to our understanding of the equity issue process in at least three ways. First, the theoretical developments in Eckbo and Masulis (1992) are formalized and extended to include private placements as an alternative to rights and underwritten offerings. In this model, both the private investor and the underwriter have the ability to partially reveal the true quality of the issuer, however, private placements are less expensive in terms of direct issue costs. Negotiations with the private investor may fail, in which case the firm has the option to withdraw the issue or to select uninsured rights or standby underwriting. In equilibrium, the average firm quality is highest in the pool of private placements and lowest in the pool of standby rights, with uninsured rights in between. While not modeled explicitly, the choice of private placements may also be driven by corporate control considerations. Contrary to rights offers, private placements create a large block holder. This blockholder may increase monitoring of management or form an alliance with management to extract corporate control benefits not available to other shareholders. If the latter effect dominates, the market reaction to private placements is predicted to be negative.

Second, the empirical analysis in Bøhren, Eckbo and Michalsen (1997) is extended to include a comparison of rights and standby offerings with the population of private placements on the Oslo Stock Exchange (OSE). Since, on the OSE, all seasoned equity offerings are issued using either rights, standby, or private placements, this extension allows more powerful tests of alternative theories for the flotation method choice. The evidence in this chapter also complements earlier studies of private placements by Wruck (1989) and Hertzel and Smith (1993) on U.S. data, and it is the first to examine private placements decision in the context of the remaining issue methods available to firms.

The rest of the chapter is organized as follows. Section 5.2 develops a theoretical model that seeks to explain what determines the choice of equity flotation method when the available methods are private placement, standby rights offerings, and uninsured rights offerings. Section 5.3 presents some descriptive characteristics of the sample of equity offerings. Section 5.4 looks at the announcement period abnormal returns for private and public equity offerings, and discusses the implications for the model developed in section 5.2. Section 5.5 concludes.

5.2 Model

To focus on the choice among the equity offering methods described in the previous section, it is assumed that firms face a short-lived investment project that must be financed using private placements, uninsured rights or standby rights. As indicated in the introduction, these flotation methods comprise the entire choice set available to OSE-listed firms. In a private placement a private investor agrees to fund the entire project in return for a fraction of the post issue company. In a standby rights offering the underwriter provides the issuer with a quality certification by committing to buy shares not taken up by the market. In an uninsured rights offering the issuer sells shares at the price determined by perfect competition among market investors. The current shareholders that subscribe to the issue in rights offerings are regarded as market investors. The game between the issuer's management and investors is modeled as depicted in figure 1.

Before the game starts, nature draws the issuer type from the discrete set Θ. The probability of drawing type θ is denoted $p(\theta)$. Thus, when investors are approached by a firm seeking financing, they do not know the full information value of the firm. In the first subgame, the issuer chooses the flotation method. If the issuer chooses a private placement or a standby rights offering, the private investor or the underwriter inspects the issuer. The inspection reduces but does not completely eliminate uncertainty about the issuers full information value. Let $q(\theta' \mid \theta)$ denote the prior probability of concluding an inspection with "this is type θ'" when the actual type is θ $(\theta, \theta' \in \Theta)$. The prior distributions $p(\cdot)$ and $q(\cdot \mid \cdot)$ are common knowledge. Following inspection, the issuer and the private investor/underwriter bargain over the offerprice (illustrated by the ovals in figure 1).[4] The bargaining proceeds as follows. The issuer suggests a price and the investor accepts or rejects. The rejection decision is private information. If the investor rejects, the issuer decides not to issue (forgoing the project) or tries another flotation method.

The second subgame (see figure 1) starts after the offerprice bargaining in private placements or standby rights fails. This subgame is identical to the first subgame, except that the remaining options are fewer. The third and last subgame is only reached if the issuer is rejected twice, once by the private investor and once by the underwriter. Consequently, there is only two options in this subgame, either issue using uninsured rights or not issue at all.

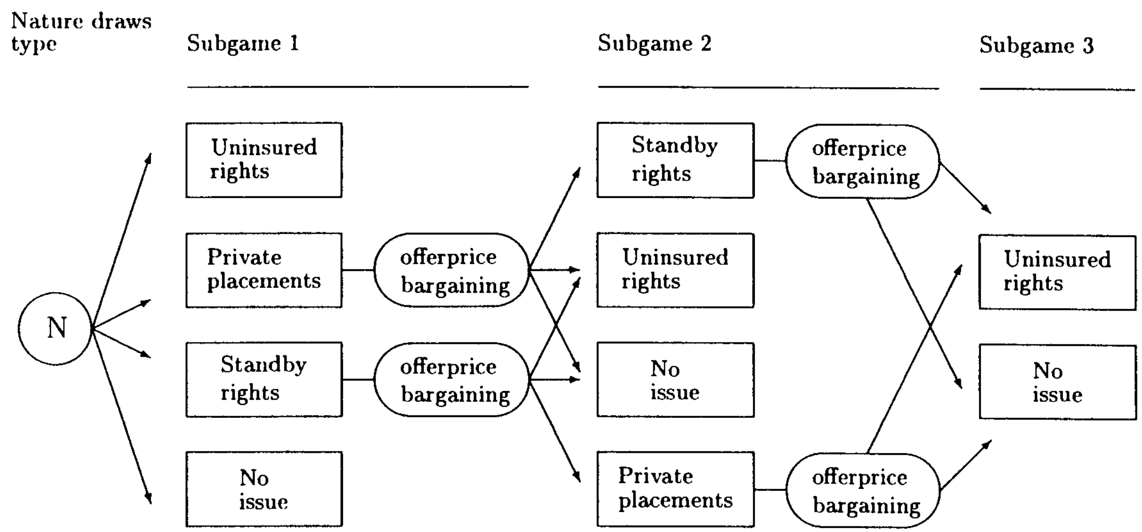
The sequential equilibrium of Kreps and Wilson (1982) is used to characterize pure strategy equilibria in this game. The optimal strategy in a sequential equilibrium must give higher expected payoff than any other strategy given the beliefs of the players. In addition, the beliefs of the players are formed using Bayes' rule and the equilibrium strategies.[5] The

[4] Giammarino and Lewis (1989) develop a theory of negotiated equity financing. The bargaining game in their model is similar to the one used here. However, their model focuses on firm commitment underwritten offerings and does not model other flotation methods.

[5] Moreover, a sequential equilibrium specifies how to form beliefs when an "off-the-equilibrium-path"

Figure 5.1: The game between equity issuers and investors

The game proceeds to a new subgame if the bargaining between issuer and investor fails. If the bargaining succeeds or either the "not issue" or the "uninsured rights" offering is chosen, the game ends.



next three paragraphs define strategies, beliefs, and payoffs in the game.

Before nature draws the issuer type, a pure strategy for this game specifies what the issuer should do throughout the game conditional on all previous actions taken by either of the players. In other words, a pure strategy maps possible types and action histories into a particular action at the next step in the game. In the bargaining game that starts when an issuer have decided on a private placement or a standby rights offering, a pure strategy for the issuer maps inspection results into an offerprice while a pure strategy for investors maps an offerprice into a decision to "accept" or "reject" the offer. Let a denote a pure strategy for the issuer.[6]

Given strategies and prior distribution of types, investors compute the probability of facing a particular type when asked for project funding. Denote the probability of facing type θ given equilibrium strategy a by $\mu^m(\theta \mid a)$, where $m \in \{\text{pp}, \text{sr}, \text{ur}\}$ indicates that the beliefs are formed by the private investor if a private placement is chosen ($m = \text{pp}$), or by underwriter if a standby offering us chosen ($m = \text{sr}$), or by market investors if an uninsured rights offering is chosen ($m = \text{ur}$). For market investors $\mu(\cdot \mid \cdot)$ represent the equilibrium beliefs about the issuer type. Private investors and underwriters use inspection results to update $\mu(\cdot \mid \cdot)$ to form equilibrium (posterior) beliefs. Let $\psi^m(\theta \mid \theta', a)$, $m \in \{\text{pp}, \text{sr}\}$, be the probability the investor put on type θ in equilibrium given that inspection concluded type θ' and issuers follow strategy a.

Managers decision whether to issue or not, depends on the value of the fraction old shareholders have in the post issue company. As in Myers and Majluf (1984), managers only issue equity if old shareholders claim in the post issue company is worth more than the value of the company when equity is not issued and the project is dropped. Suppose the assets in place are worth a_θ for a type θ issuer, the net present value of the project is b_θ, the required level of investments is I, the direct issue cost is d (underwriter fees, legal fees, mailing costs etc.), the post issue number of shares is S, the ratio of the new number of shares to S is δ, and the fraction of the issue taken up by current shareholders is k.[7] Thus, the fraction of the post issue company held by new shareholders is $\Delta \equiv \delta(1-k)$, and managers issue new equity and invest in the project only if $(1-\Delta)(a_\theta + b_\theta + I - d) \geq a_\theta + kI$. When

action is observed. Strategies and beliefs in a a sequential equilibrium are *consistent* if beliefs after an "off-the-equilibrium-path" action is formed assuming arbitrarily small "trembles" by all types that could possibly reach the stage preceding the "off-the-equilibrium-path" action. The possibility of "trembles" implies that all actions have positive probability, and therefore, Bayes' rule can be used to form beliefs.

[6]As an example, a pure strategy for the issuer could be: If type θ is draw, seek funding from a private investor. If inspection says type θ', propose price $P_{\theta'}$. If private investor rejects, try an uninsured rights offer. If inspection says type θ'', propose price $P_{\theta''}$. If underwriter rejects, do not issue.

[7]Myers and Majluf (1984) assume $k = 0$, while Eckbo and Masulis (1992) allow $k \leq 1$.

this condition is satisfied the value of the claim old shareholders have in the post issue company exceeds the value of the company without the project plus the funds provided by old shareholders. Given the post-issue number of shares, let $P_\theta(a)$ be the expected offerprice for a type θ issuer that follows strategy a.[8] Rearranging the issue-condition gives:

$$b_\theta \geq d + \Delta\left[(a_\theta + b_\theta + I - d) - P_\theta(a)S\right]. \tag{5.1}$$

The rights hand side of equation (5.1) is the issue costs. The first term is the direct issue costs, while the second term is the wealth transfer between old and new shareholders. If the equilibrium price times the post issue number of shares is less than the full information value of the post issue company, old shareholders experience a wealth transfer cost. The issuer decides to issue as long as the net present value of the project exceeds the issue costs.

The direct costs for standby rights offerings exceed the direct costs for uninsured rights offerings because the former include standby fees. The direct costs for private placements do not include mailing costs and costs of administrating the rights offer subscriptions. Thus, the direct costs for a private placement are less than the direct costs for uninsured rights.[9] Denoting direct costs by $d(m)$, we have $d(\mathrm{pp}) < d(\mathrm{ur}) < d(\mathrm{sr})$.

As in Eckbo and Masulis (1992) the significance of the wealth transfer between old and new shareholders depends on the number of old shareholders that subscribe to the issue. If old shareholders subscribe to the whole issue, the wealth transfer is zero. On the other hand, if old shareholders do not participate in the issue (as in a private placement), the wealth transfer is potentially large.

5.2.1 A sequential equilibrium

This section characterizes and discusses a sequential equilibrium that may occur in the game between equity issuers and investors. Suppose the full information value of an issuer can take on one of three values: v_1, v_2, and v_3, with $v_1 < v_2 < v_3$. Assume that there are two different takeup levels $\underline{k}$ and $\overline{k}$, with $\underline{k} < \overline{k}$. The set of possible types is then $T \equiv \{\underline{v}_1, \underline{v}_2, \underline{v}_3, \overline{v}_1, \overline{v}_2, \overline{v}_3\}$, where $\underline{v}_j$ and $\overline{v}_j$ denote firms with full information value v_j and takeup levels $\underline{k}$ and $\overline{k}$ respectively ($j = 1, 2, 3$). The inspection is assumed to be informative

[8] The offerprice is determined under the equilibrium beliefs of investors. For rights offerings, the offerprice is the sum of the subscription price set by the issuer and the total market value of the number of rights necessary to acquire one new share. In private placements, the offerprice depends on the outcome of the bargaining game between issuer and investor.

[9] Bøhren, Eckbo and Michalsen (1997) report that the direct cost of uninsured rights offerings are about 70% of the direct cost for standby rights offerings. Moreover, they report that rights subscriptions are handled by a bank that normally charges a fixed fee plus about 1.5% of the value of shareholder subscriptions.

and symmetric in the sense of definition 1 and 2.

Definition 1 *An inspection is said to be informative if* $q(v_i \mid v_i) > q(v_i \mid v_j) > q(v_i \mid v_k)$ *for* i *adjacent to* j*, and* i *not adjacent to* k.

Definition 2 *An inspection is said to be symmetric if* $q(v_i \mid v_j) = q(v_j \mid v_i)\ \forall i, j$.

Definition 1 says that the probability of concluding any particular type v_i decreases with the difference in value between v_i and the type being inspected. Let the pure strategy $a_\theta = \{\mathrm{pp}_\theta, \mathrm{sr}_\theta, \mathrm{ur}_\theta\}$ have the interpretation that a type θ issuer first approaches a private investor and then follows the optimal strategy in this subgame. If the private investor rejects the issuer, use a standby underwritten offering and follow the optimal strategy in this subgame. If the underwriter rejects the offer, use an uninsured rights offering.

Taking the number of new shares as given, the bargaining subgame that starts after an inspection can only succeed if the breach-price of the issuer (his lowest acceptable price) is below the breach-price of the the private investor or the underwriter (their highest acceptable prices). Denote the breach-prices of a type v_i issuer and the investor given inspection result v_j as $B_i^m(v_j)$ and $B^m(v_j)$, respectively (m = pp, sr and $i, j = 1, 2, 3$). The breach-price of the issuer is determined by the expected price that he will get in the continuation game that starts after a bargaining fails. The breach-price of the investor is determined by the expected value of the issuer under the post-inspection equilibrium beliefs $\psi^m(\cdot)$. Consider the following proposition.

Proposition 3 *If the inspection is informative and symmetric, the following strategies are part of a sequential equilibrium* ($i = 1, 2, 3$ *and* $m = \mathrm{pp}, \mathrm{sr}$)*:*

(a) $\{\mathrm{pp}_\theta, \mathrm{sr}_\theta, \mathrm{ur}_\theta\}$ *for* $\theta \in \{\underline{v}_1, \underline{v}_2\}$, $\{\mathrm{pp}_\theta, \mathrm{sr}_\theta\}$ *for* $\theta = \underline{v}_3$*, and* $\{\mathrm{ur}_\theta\}$ *for* $\theta \in \{\overline{v}_1, \overline{v}_2 \overline{v}_3\}$

(b) If $B_i^{\mathrm{pp}}(\theta) <= B^{\mathrm{sr}}(\theta)$ *then propose offerprice* $B^{\mathrm{sr}}(\theta)$*, otherwise propose offerprice* $B_i^{\mathrm{pp}}(\theta)$.

(c) For the investors, "accept" all offerprices greater than or equal to $B^m(\theta)$*, "reject" all offerprices lower than* $B^m(\theta)$

Proof By example ■

Table 5.1 presents an example of an equilibrium that satisfies proposition 3. Any issuer with low current shareholder takeup first tries to float the new equity using a private placement. If the inspection by the private investor concludes type $\underline{v}_3$, all types propose an offerprice of 185 to the private investor, which makes him indifferent between accepting and rejecting the offer. If the inspection by the private investor concludes type $\underline{v}_2$, the lowest offerprice acceptable to the private investor is 136, while the highest offerprice acceptable to the high-value issuer is 150. Thus, there exists no offerprice that is optimally proposed by the issuer and that is accepted by the private investor. The reason is that using a standby rights offer has expected issue cost for the high-value type of 6.98, which is less than the 9.86 issue costs for private placement following the type $\underline{v}_2$ inspection conclusion. In the standby rights subgame, if the underwriter concludes the inspection with type $\underline{v}_3$, all issuer types propose offerprice 160, which is accepted by the underwriter. This gives the high-value type issue costs of 6.23. If the inspection conclusion is type $\underline{v}_2$, the high-value type would experience issue costs of 12.98 when proposing the lowest acceptable offerprice to the underwriter. This exceeds the net present value of the projects. Since the issue costs for the high-value type in an uninsured rights offering will be 19.96, it is optimal for the high-value type not to issue. Even if $\underline{v}_3$ foregoes the project, types $\underline{v}_1$ and $\underline{v}_2$ optimally propose an offerprice that the underwriter accepts. This results in costs of 1.41 for type $\underline{v}_2$ in the standby rights offering. If the inspection conclusion is $\underline{v}_1$, type $\underline{v}_2$ is best off proposing an offerprice that the underwriter rejects, and then using an uninsured rights offering to save on the direct issue costs. This strategy yields expected costs in the uninsured rights offer of 5.68, while the standby rights offer costs would have been 5.83.

Given equilibrium beliefs formed using the prior distribution of types, the equilibrium strategies and inspection results, the proposed equilibrium is sustained if no type can obtain lower issue costs by deviating. The intuition for the existence of this equilibrium is explained in two steps. First, it is argued why low-k issuers follow the equilibrium strategies $\{pp_\theta, sr_\theta, ur_\theta\}$ and $\{pp_\theta, sr_\theta\}$. Second, the equilibrium strategy of high-k issuers is explained.

Any optimal strategy that includes inspection, must have pp_θ as the first action. This result rests on the assumption that private placement investors and underwriters have identical inspection ability, and that the issuer-investor bargaining situations are set up so that the issuer gets all the bargaining power. These two assumptions imply that the expected wealth transfers are the same for these flotation methods. Thus, a type that optimally chooses to be inspected, selects a private placement first since this has the lowest direct costs. Given that inspection is informative, high-value firms with low-k also prefer a private placement to an uninsured rights offer. There are two reasons for this. First, informative inspection increases the probability put on the high-value type relative to the prior distribution. Second, in the proposed equilibrium, low-k issuers that use an uninsured rights

Table 5.1
Example of equilibrium in proposition 3

The inspection performed by the private investor and the underwriter is informative and symmetric. The direct costs are highest for standby rights and lowest for private placements. The probability private investors and underwriters put on facing each type is consistent with Bayes' rule using the prior distribution, the inspection ability, and the equilibrium strategy. The equilibrium beliefs are not reported for this example.

(a) Exogenous parameters

Types	Prior	Inspection	Direct costs	Investment	Project NPV
$v_1 = 100$	$p(\underline{v}_1) = p(\bar{v}_1) = 0.17$	$q(v_1 \mid v_1) = 0.80$	$d_{pp} = 0.15$	30	10
$v_2 = 130$	$p(\underline{v}_2) = p(\bar{v}_2) = 0.16$	$q(v_2 \mid v_2) = 0.70$	$d_{ur} = 0.20$		
$v_3 = 200$	$p(\underline{v}_3) = p(\bar{v}_3) = 0.17$	$q(v_3 \mid v_3) = 0.80$	$d_{sr} = 1.10$		
$\bar{k} = 0.95$		$q(v_1 \mid v_2) = 0.05$			
$\underline{k} = 0.30$					

(b) Equilibrium private placement subgame

			Inspection conclusions					
			Type $\underline{v}_1$		Type $\underline{v}_2$		Type $\underline{v}_3$	
Type	Breach prices issuer	Expected cost if rejected	Breach prices investor	Issue cost	Breach prices investor	Issue cost	Breach prices investor	Issue cost
$\underline{v}_1$	106	−1.23	109	−1.75	136	−5.55	185	−9.60
$\underline{v}_2$	122	1.39	109	4.01	136	−0.92	185	−6.20
$\underline{v}_3$	150	6.98	109	17.46	136	9.86	185	1.74

(c) Equilibrium standby rights subgame

			Inspection conclusions					
			Type $\underline{v}_1$		Type $\underline{v}_2$		Type $\underline{v}_3$	
Type	Breach prices issuer	Expected costs if rejected	Breach prices investor	Issue cost	Breach prices investor	Issue cost	Breach prices investor	Issue cost
$\underline{v}_1$	107	−0.44	105	−0.15	127	−3.55	160	−6.91
$\underline{v}_2$	106	5.68	105	5.83	127	1.41	160	−2.97
$\underline{v}_3$	140	10.00	105	19.82	127	12.98	160	6.23

(d) Equilibrium uninsured rights subgame

Low takeup types ($\underline{k}$)				High takeup types ($\bar{k}$)				
					Breach costs			
Type	Breach costs	Breach prices investor	Issue cost	Type	Forgo project	Standby rights	Breach prices investor	Issue cost
$\underline{v}_1$	10	103	−0.44	$\bar{v}_1$	10	1.29	144	−0.26
$\underline{v}_2$	10	103	5.68	$\bar{v}_2$	10	1.33	144	0.06
$\underline{v}_3$	10	103	19.96	$\bar{v}_3$	10	1.32	144	0.79

offer have been rejected both by a private investor and an underwriter. The probability that this happens is much larger for a low-value type than for a high-value type. Thus, the equilibrium beliefs of market investor when they observe an uninsured rights offer by a low-k type put large weight on the low-value type.

If the issuer is rejected by the private investor, the optimal strategy specifies that all low-k firms contact an underwriter in order to discuss terms for a standby rights offering. Conditional on the inspection result, $\underline{v}_3$ foregoes the project if the inspection concludes with a wrong type. This is optimal since underwriter in this situation puts very little probability on type v_3 when forming beliefs. Thus the wealth transfer cost for old shareholders outweighs the net present value of the project. For the remaining types $\underline{v}_1$ and $\underline{v}_2$ the optimal strategy specifies that they propose an offerprice that will be accepted by the underwriter if the inspection concludes v_2 and an offerprice that will be rejected if inspection concludes v_1. This is optimal as long as the expected costs of a standby rights offering are less than the expected costs of forgoing the project or using an uninsured rights offering for type $\underline{v}_2$. This is more likely to be the case the higher the precision of the underwriter inspection. More precise inspection has the direct effect of reducing the wealth transfer costs and the indirect effect of reducing the attractiveness of an uninsured rights offer. The indirect effect is caused by an increase in the equilibrium probability market investors put on the low-value firm when they observe an uninsured rights offer with low current shareholder takeup. If the inspection concludes type $\underline{v}_1$ such that the issuer also is rejected by the underwriter, issuer types $\underline{v}_1$ and $\underline{v}_2$ choose an uninsured rights offering as long as the net present value of the project is higher than the issue costs.

Since preemptive rights are not waived for firms with a high takeup level, the high-k types $\{\overline{v}_1, \overline{v}_2, \overline{v}_3\}$ cannot deviate from the equilibrium using private placements. The high-value type $\overline{v}_3$ does not deviate using a standby offering since the low direct costs of uninsured rights and the relatively low wealth transfer caused by a high takeup outweigh the advantage of reducing uncertainty about the type through underwriter inspection. The types $\overline{v}_1$ and $\overline{v}_2$ do not deviate because inspection is informative, implying that the weight underwriters put on the high-value type when they are asked to stand by the issue of types $\overline{v}_1$ and $\overline{v}_2$ is less than the prior probability of this type.

Finally, the equilibrium specifies that private investors and underwriters optimally accept all offers larger or equal to their breach-prices. This is optimal because investors are indifferent between accepting or rejecting the offer at the breach price, but make a profit if the offerprice exceeds the breach-price. It follows that it cannot be optimal for the issuer to offer a price higher than the investor's breach price. When the inspection in the private placement subgame concludes that the issuer is the high-value type $\underline{v}_3$, the breach-price of the investor exceeds the breach-price of all issuer types, causing $B^m(v_3)$ to be the optimal offer for the issuer. On the other hand, if the private investor concludes that the issuer is the medium-value type $\underline{v}_2$, the breach-price of the investor is below the breach-price of the high-value issuer. Thus, the set of offerprices that gives both players non-negative payoff is empty. The optimal action for the issuer is thus to propose any price in $[0, B^m_{\text{iss}}]$, which causes the investor to reject. A similar situation arises in the standby rights offering subgame. If the inspection conclusion is type $\underline{v}_2$, any issuer type optimally proposes the offerprice $B^m(v_3)$. When the inspection conclusion is type $\underline{v}_1$, the breach-price of the medium-value investor $\underline{v}_2$ exceeds the breach price of the underwriter and the bargaining fails.

5.2.2 Empirical implications

Proposition 3 yields testable predictions concerning the relative magnitude and sign of abnormal returns around equity issue announcements. Since inspections are informative, the probability that the private investor will accept a request for funds by a high-value issuer is larger than the prior probability of a high-value type. Thus, when a private placement is announced, the market infers that the probability of a high-value type is larger than that given by the prior probability, causing a positive market reaction around the announcement.

When a standby rights offering is announced, market investors infer that the inspection result of the private investor was either $\underline{v}_1$ or $\underline{v}_2$. Again, since inspections are informative, this is a relatively unlikely outcome if the type of the issuer is $\underline{v}_3$. Thus, the probability that the type is $\underline{v}_3$ in a standby offering is less than the prior probability $p(\underline{v}_3)$. This results in a negative market reaction when a standby offering is announced.

The market reaction at the announcement of an uninsured rights offering depends on the expected current shareholder takeup level k. If an uninsured rights offering is announced and the market expects the issue to get a high current shareholder takeup, the market reaction should be positive. The good news come from the fact that ex ante, the pool of potential issuers $\{\underline{v}_1, \underline{v}_2, \underline{v}_3, \overline{v}_1, \overline{v}_2, \overline{v}_3\}$ is priced with positive probability on the event that the high-value type with low current shareholder takeup $\underline{v}_3$ forgos the project in equilibrium. Thus, given that the market expect k to be high, the set of potential issuers is reduced to $\{\overline{v}_1, \overline{v}_2, \overline{v}_3\}$. However, since no inspection is performed, the market reaction should be positive but closer to zero than the market reaction to private placements. On the other hand, if the expected takeup level is low, the issuer has been rejected by both a private investor and an underwriter—which is a much more likely event if the issuer is the low-value type. Thus, the revised beliefs of the market put a relatively high probability on the low-value type, causing a negative market reaction when the issue is announced. Moreover, since the revised beliefs are formed conditional on a rejection by the underwriter, the probability of a low-value type is larger for uninsured than for standby rights offering. This implies that the market reaction to an uninsured rights offering by a firm with low expected takeup should be more negative than the market reaction to a standby rights offering.

Proposition 3 also predicts that the average current shareholder takeup should be higher for uninsured rights offerings than for standby rights offerings. This is consistent with the findings of Eckbo and Masulis (1992) and Bøhren, Eckbo and Michalsen (1997).

5.3 Sample characteristics

The empirical analysis examines a total of 381 rights offers and private placements on the OSE over the period 1980–1996. The rights offer sample consists of 200 cases compiled by Bøhren, Eckbo and Michalsen (1997) from the years 1980–1993,[10] as well as an added 28 cases collected here. These 28 cases and the private placements are identified by reading annual reports published by the OSE. The announcement dates are found in either the OSE Daily Bulletin or in newspapers covering business in Norway.[11]

Table 5.2 and table 5.3 show some descriptive statistics for the sample of Norwegian equity issues and issuers. The SEO sample contains the population of offerings during the period 1980–1996. The sample of private placements includes all offerings made to outside investors. Offerings made under employee stock ownership plans are excluded. This gives a total of 153 private equity offerings over the years 1984–1996. If the private placement is made as part of a merger or an acquisition, other means of floating equity might not be a viable option. For example, if the financing package in an acquisition involves bidder shares, a rights offer cannot be used. In the language of section 5.2, we want all the sample firms to be from the same typespace Θ of potential issuers. For this reason, market reactions to the announcement of private placements should be studied both including and excluding mergers and acquisitions. However, data on the use of proceeds from private placements are limited. Fortunately, data on all mergers and acquisitions on OSE is available. Private placements that potentially have been made in a merger or an acquisition are identified

[10] I thank the authors for making their data available. Of the 200 announcements in the Bøhren-Eckbo-Michalsen data, 2 were modified for use in this paper.

[11] The newspapers searched for announcements are Dagens Næringsliv and Aftenposten. Aftenposten is searched using the computer searchable full-text database Atekst. Some announcement dates for private placements 1996-1995 are from Garberg, Gifstad and Henriksen (1998). The private placements are collected in cooperation with Øyvind Bøhren and Ståle Legreid.

Table 5.2
Number of seasoned equity offerings by OSE-listed firms in the sample, 1980–1996

The sample of public offerings for the years 1980–1993 is the sample used by Bøhren, Eckbo and Michalsen (1997). The issues during the remaining years of the public offering sample period and the sample of private placements are identified from annual reports published by the OSE.

				Public equity offerings			Private placements	
Year	Number of offerings in percent of number of firms	Number of firms	Total number of offerings	Total	Standby rights	Uninsured rights	Including mergers	Excluding mergers
1980	3.2	124	4	4	0	4	–	–
1981	11.2	116	13	13	4	9	–	–
1982	16.9	118	20	20	2	18	–	–
1983	24.4	119	29	29	9	20	–	–
1984	18.9	148	28	24	14	10	4	2
1985	19.0	163	31	24	13	11	7	3
1986	14.2	155	22	14	10	4	8	7
1987	8.9	146	13	8	7	1	5	4
1988	16.4	134	22	10	10	0	12	12
1989	30.2	129	39	21	19	2	18	14
1990	24.0	121	29	9	9	0	20	12
1991	17.0	112	19	6	6	0	13	11
1992	11.4	123	14	9	9	0	5	4
1993	23.7	131	31	14	13	1	17	16
1994	16.4	134	22	11	11	0	11	9
1995	12.3	163	20	4	4	0	16	16
1996	14.5	173	25	8	8	0	17	16
Sum	16.5	2309	381	228	148	80	153	126

as those with an announcement date closer than one year to the date of a merger or an acquisition. Excluding these reduces the private placements sample to 126 observations.

The second column of Table 5.2 shows the number of issues as a percent of the number of firms listed on the OSE. In 1980 the number of issues relative to the number of firms listed is 3.2%, while in 1986 this fraction is as high as 30.2%. In most years, however, the fraction is between 10% and 20%—which is very high compared to the U.S. Even in the "hot issue" market during the first part of the 1980s, the fraction of issues relative to the number of firms listed on New York Stock Exchange and American Stock Exchange is only about 10%. For the public offering sample, there is a tendency towards the exclusive use of standby offerings towards the end of the sample period. In light of Eckbo and Masulis (1992), this could be explained by a reduction in the expected current shareholder takeup towards the end of the sample period. Consistent with this view, Bøhren, Eckbo and Michalsen (1997) report lower average takeup in the subperiod 1985–1993 than in the subperiod 1980–1984.

Panel (a) in table 5.3 provides descriptive statistics for issuers and issues. The table reports a total of 228 public offerings and 153 private placements over the sample period. These offerings are made by 116 and 91 different firms respectively. Thus, firms that float new equity using rights offers, have on average issued new equity twice during our sample period. For private placements the average firm issues 1.7 times over the sample period. However, the average firm is not representative in that there are firms that issue new equity with some consistency, while the majority of firms issue only once. For both the public offering and private placement sample, about 60 firms issue only one time during the sample period. The largest number of issues are 7 for the public offering sample and 6 for the private placement sample.

For public offerings the size of both the issuer and the issue is considerably larger for standby offerings than it is for uninsured rights offerings. However, the numbers are in nominal terms, so part of the difference reflects the fact that most uninsured offerings are made during the first part of the sample period while most of the standby offerings are from the later years. The average offersize in private placements is NOK 120 million, which is about 20 million less than the average offersize for rights offerings. Looking at the average equity market value of the offering firms, the average size of firms offering equity through standby rights offers and private placements is about the same. Note that firms floating new equity through a public offering have a larger book-to-market ratio than firms placing private equity. To the extent that the book-to-market ratio is a measure of the value of assets in place relative to the value of the growth opportunities, this is not consistent with the idea that private placements are used to resolve the underinvestment of Myers and

Table 5.3
Descriptive statistics for seasoned equity offerings by OSE-listed firms in the sample, 1980–1996

		Public equity offerings			
	Total	Total	Standby rights	Uninsured rights	Private placements
(a) Issue and issuer characteristics					
Number of observations	381	228	148	80	153
Amount offered (NOK millions)	136	146	190	65	120
Market value of equity (NOK millions)	749	667	856	316	870
Book-to-market ratio	1.01	1.15	1.22	1.04	0.79
(b) Ownership characteristics[a]					
Percent holding by the ten largest owners at the beginning of the issue year	–	–	57.42 (81)	56.77 (19)	64.96 (77)
Percent holding by the ten largest owners at the end of the issue year	–	–	57.95 (98)	47.89 (28)	65.81 (66)
Percent holding by insiders[b] at the beginning of the issue year	–	–	8.31 (115)	6.35 (66)	18.13 (88)
Percent holding by insiders at the end of the issue year	–	–	6.16 (117)	4.99 (74)	15.44 (74)

[a]Ownership information is from the companies annual reports. Numbers in parentheses are number of observations

[b]Insiders include the CEO and members of the board of directors

Majluf (1984). Underinvestment arises because the wealth loss incurred by old shareholders from selling undervalued equity outweighs the gain from realizing a project with positive net present value. The underinvestment problem, and the subsequent need to resolve it using private placement, is more likely to occur if the value of assets in place is large relative to the value of growth opportunities (i.e., when the book-to-market ratio is high).

Panel (b) in table 5.3 provides information about the ownership structure of the issuing firm. The aggregated holdings of the ten largest shareholders reveal that Norwegian companies have a relatively concentrated ownership structure. It is interesting to see that there is only a very small increase in average concentration for firms that do private placements of equity. Together with the high concentration, this suggests that monitoring of management hardly increases as a result of the private placement.

Table 5.4 describes takeup characteristics for the public offering sample. The takeup for an issue is defined as the fraction of the issue acquired by the old shareholders of the issuing company. In a rights offering, the current shareholders are given a short-term warrant for each stock they own. Following Bøhren, Eckbo and Michalsen (1997), it is assumed that each warrant is traded only once, so that takeup can be proxied by one minus the fraction of warrants traded in the secondary market.

Consistent with the model developed in section 5.2, panel (a) of table 5.4 shows that current shareholder takeup tends to be higher for uninsured rights offerings than for standby offerings. The mean takeup for uninsured rights is 94.4% while the mean takeup for standby rights is 84.8%. The corresponding numbers using median to measure differences are 96.4% and 91.5%. Panel (b) and (c) show takeup in two subsamples created by splitting the public offering sample according to whether or not the issue had a takeup above or below the median takeup. As predicted by the model, a disproportionately large number of uninsured rights offerings are in the subsample with above median takeup, and a disproportionately large number of standby rights offerings are in the subsample with below median takeup. The subsamples also reveal a "non-linearity" in the data. Panel (b) of table 5.4 shows that issues with below median takeup exhibit the same tendency as the overall sample—uninsured rights have a higher takeup than standby rights. However, panel (c) shows that the takeup in the above-median subsample is similar for both offer types.

Table 5.4
Current shareholder takeup in seasoned equity offerings by OSE-listed firms in the sample, 1980–1996

Takeup is defined as the fraction of the issue acquired by the old shareholders. Assuming that each warrants issued in a rights offering only is traded once, takeup is proxied by one minus the fraction of warrants sold in the secondary market.

	N	Mean	STD	90%	75%	50%	25%	10%
(a) All issues								
Uninsured rights	80	94.38	7.03	98.89	97.99	96.43	93.44	89.63
Standby rights	148	84.81	17.14	99.27	96.34	91.47	79.68	61.17
Total	228	88.17	15.11	99.08	97.34	93.93	84.77	69.40
(b) Issues with below median takeup								
Uninsured rights	24	87.68	9.91	93.60	93.19	91.78	87.74	74.64
Standby rights	90	76.76	17.76	92.48	89.68	81.52	70.82	52.65
Total	114	79.06	16.98	93.43	91.51	84.77	73.11	55.27
(c) Issues with above median takeup								
Uninsured rights	56	97.24	1.41	99.08	98.24	97.32	96.28	95.19
Standby rights	58	97.31	2.05	1.00	99.27	97.40	95.33	94.51
Total	114	97.28	1.76	1.00	98.00	97.34	95.80	94.98

5.4 Announcement day abnormal returns

Six month abnormal stock price run-up and announcement day abnormal returns for issuer i are computed using the following market model:

$$r_{it} = \alpha_i + \beta_i r_{mt} + \sum_{j=1}^{2} \gamma_{ij} d_{jt} + \epsilon_{it},$$

where r_{it} is continuously compounded daily return on firm i and r_{mt} is continuously compounded daily return on a value weighted market portfolio of all OSE-listed firms. The estimation period starts on trading day -310 relative to the announcement date and ends on trading day $+160$ relative to the announcement date. Issues with less than four months (82 trading dates) of data prior to the announcement are excluded from the analysis. This reduces the standby- and rights offering samples with 5 and 4 observations respectively. The private placement sample is reduced with 17 observations (14 observations if mergers are excluded). The resulting estimation period is 471 trading dates (the maximum) for about 88% of the issues, and the shortest estimation period is 244 trading dates. Two dummy variables are included to capture abnormal stock price run-up prior to the announcement and the announcement period return. The first dummy variable, d_{1t}, takes on a value of one on event dates days -130 through -4, and zero otherwise.[12] Announcement period abnormal returns are estimated using both a two-day and a four-day window ending on the announcement day. In the former specification, the dummy variable d_{2t} takes on a value of one on event dates -1 and 0, and zero otherwise. The percentage abnormal return for a ω-day event window is $\omega \times \gamma_{ij} \times 100$. Under the null hypothesis of zero abnormal return, the following test statistic converges in distribution to the standard normal distribution

$$z_j = \frac{1}{\sqrt{N}} \sum_{i=1}^{N} \frac{\hat{\gamma}_{ij}}{\hat{\sigma}_{ij}},$$

where $\hat{\gamma}_{ij}$ is the OLS estimate of γ_{ij}, and $\hat{\sigma}_{ij}$ is the estimated standard error of this estimate.

Table 5.5 reports the abnormal stock price run-up and announcement day abnormal returns for the public offering and private placement samples. Event studies of U.S. seasoned equity offerings have typically used a two-day event window to measure the announcement effect. However, event studies of private placements (e.g., Wruck (1989) and Hertzel and Smith (1993)) use a four-day announcement window. The results for the present Norwegian

[12]Event dates are trading dates relative to the announcement date, which is defined to be event date 0.

samples are sensitive to this choice. Thus, table 5.5 report announcement period abnormal return using both a two-day and a four-day event window.

The results for the public offering sample are very similar to the findings of Bøhren, Eckbo and Michalsen (1997). The two-day announcement period abnormal return for standby rights offerings are negative but statistically insignificant, while uninsured rights offerings show a positive abnormal return of about 1.0% (statistically significant at the 10% level). The results for standby rights offerings remain unchanged when using a four-day announcement window. Uninsured rights offerings, on the other hand, show a highly statistically significant four-day announcement period return of about 2.1%. That is, uninsured rights offerings outperform the market by on average 50 basis point each day in the four-day window ending at the issue announcement date.[13] The announcement period abnormal returns for public offerings reported in table 5.5 are consistent with the model predictions from section 5.2. Based on this model, the pool of issues that uses standby rights in equilibrium has lower average value than the post issue pool of potential issuers. This cause a negative market reaction when the standby rights issue is announced.

The market reaction to uninsured rights offers are on average positive. This is also consistent with the predictions of the model. However, the model also predicts that uninsured rights offers by firms with low takeup should be met with a negative market reaction that is larger than the negative market reaction for standby rights offers. Using observed takeup as a proxy for the expected takeup, there is only weak evidence of this in the data. A regression of two-day abnormal returns for uninsured rights offerings on current shareholder takeup produces a negative and statistically significant coefficient (regression not reported.) When the sample is divided in two using the median takeup, the high-takeup sample shows an average two-day abnormal return of 1.33% and the low-takeup sample an average abnormal return of 0.82%. However, the difference in average abnormal return between the groups is not statistically different from zero using a standard t-test (t-test not reported.)

The announcement period abnormal returns for private placements resembles those for uninsured rights. Using a two-day announcement period, the abnormal return is positive but statistically insignificant. Adding another two days to the event window increases the announcement day abnormal return to a statistically significant 2.7%—or almost 68 basis

[13]This effect ends with the four-day window. Increasing the window beyond four days decreases the daily average abnormal return. There are several possible explanations for the finding that a four-day window produces larger abnormal returns than the usual two-day window. The most likely explanation is that some of the sample issuers experience infrequent trading such that there exists no price in the two-day event window. However, it could be caused by either imprecise announcement dates or information leakage prior to the official announcements.

Table 5.5
Average pre-announcement date run-up and announcement-day abnormal returns to OSE-listed firms making public and private equity offerings,1980–1996

Six month run-up and announcement day abnormal returns for issuer i are computed using the following market model:

$$r_{it} = \alpha_i + \beta_i r_{mt} + \sum_{j=1}^{2} \gamma_{ij} d_{jt} + \epsilon_{it},$$

where r_{it} is daily return on firm i, r_{mt} is daily return on a value weighted market portfolio of all OSE-listed firms. The estimation period is a total of 471 trading dates, starting on trading day -310 relative to the announcement date and ending on trading day +160 relative to the announcement date. The dummy variable d_{1t} takes on a value of one on trading days -130 through -4. The dummy variable d_{2t} takes on a value of one in either a two-day or a four-day window ending at the announcement day. The percentage abnormal return for a ω-day event window is $\omega \times \gamma_{ij} \times 100$. Under the null hypothesis of zero abnormal return, the following test statistic converge in distribution to the standard normal

$$z_j = \frac{1}{\sqrt{N}} \sum_{i=1}^{N} \frac{\hat{\gamma}_{ij}}{\hat{\sigma}_{ij}}.$$

Where $\hat{\gamma}_{ij}$ is the OLS estimate of γ_{ij}, and $\hat{\sigma}_{ij}$ is the standard error of this estimate. Returns are continuously compounded. The parentheses contain number of observations and p-values. The p-values are for two-sided tests.

	Public equity offerings		Private placements	
	Standby rights	Uninsured rights	Including mergers	Excluding mergers
Six-months runup	4.60 (143; .004)	9.15 (76; .028)	18.52 (136; .000)	20.32 (112 : .000)
Two-day announcement return	−0.58 (143; .234)	0.95 (76; .088)	1.39 (136; .058)	1.00 (112 : .318)
Four-day announcement return	−0.55 (143; .502)	2.11 (76; .000)	2.66 (136; .022)	2.66 (112; .050)

point per day. These results are consistent with the findings of Wruck (1989) and Hertzel and Smith (1993) on U.S. data and with Kang and Stulz (1996) on Japanese data.[14] As recognized by Wruck (1989), the announcement day abnormal return measures the combined wealth effect for old shareholders of the new information and the effect of selling a fraction of the company at a discount or premium. In other words, the announcement effect is the information effect net of issue cost. The return to old shareholders due to the information effect can be measured using an adjusted abnormal return:

$$\text{AAR} = \left(\frac{1}{1-\delta}\right)\text{AR} - \left(\frac{\delta}{1-\delta}\right)\left(\frac{P_{\text{Offer}} - P_{a-4}}{P_{a-4}}\right), \qquad (5.2)$$

where δ is the fraction of the company owned by the new private investors after the private placements, P_{Offer} is the offer price, P_{a-4} is the market price four days prior to the announcement day, and AR is the four-day abnormal stock return. To look at an example: Suppose the private investor will own 20% of the shares after the issue ($\delta = .2$), and the average four-day announcement period abnormal return is 2.6% (AR= 0.026). In order for "issue costs" to explain the positive announcement day effect, the issue must be placed at a premium of 13%. Table 5.6 reports AAR using the observations for which the placement prices are available.

Looking at the first row of table 5.6, the average private equity placement gives the new investors a holding of approximately 24% of the post issue company. Surprisingly, the average holding is acquired at a premium of 3.5% when mergers and acquisitions are included, and at a 0.9% premium when mergers and acquisitions are excluded. However, neither of these numbers are significantly different from zero at conventional levels using a standard t-test. As a robustness check, the sample is trimmed by excluding extreme observations. The weak trimming excludes observations with a discount less than 60% or a premium higher than 60%, the medium and strong trimming use 40% and 20% as cutoff percentages, respectively. The trimmed samples also exhibit a positive premium, although smaller and still not statistically different from zero. Even though the average premium is positive, the average adjusted abnormal returns (AAR) are positive and *larger* than the average unadjusted abnormal returns (AR). This can only happen if the private placements with a positive premium typically give new investors a smaller fraction of the post issue company (i.e., δ is low) than placements that are done at a discount. Computing

[14] Using a sample of 128 private placements, Wruck (1989) reports a four-day announcement period abnormal return of about 4.4%, while Hertzel and Smith (1993) report a slightly lower abnormal return. Kang and Stulz (1996) find a cumulated abnormal return of 2.13% in a sample of 69 private equity offerings in Japan.

Table 5.6
Average adjusted four-day announcement-day abnormal returns to OSE-listed firms making private equity offerings,1984–1996

The return to old shareholders due to the information effect can be measured using an adjusted abnormal return:

$$\text{AAR} = \left(\frac{1}{1-\delta}\right)\text{AR} - \left(\frac{\delta}{1-\delta}\right)\left(\frac{P_{\text{Offer}} - P_{a-4}}{P_{a-4}}\right),$$

where δ is the fraction of the company owned by the new private investors after the private placements, P_{Offer} is the offer price, P_{a-4} is the market price four days prior to the announcement day, and AR is the four-day abnormal stock return. The sample trimming excludes extreme observations: weak trimming exclude observations with a discount less than 60% or a premium higher than 60%, the medium and strong trimming use 40% and 20% as cutoff percentages, respectively.

	Including mergers					Excluding mergers				
Sample trimming	N	Prem	δ	AR	AAR	N	Prem	δ	AR	AAR
None	133	3.56	23.96	2.67	37.19	109	0.92	23.65	2.67	42.34
Weak [−60%, +60%]	115	0.41	20.79	1.81	5.85	94	0.81	20.13	1.49	5.58
Medium [−40%, +40%]	109	1.39	20.30	1.38	4.85	90	1.95	19.96	1.46	4.91
Strong [−20%, +20%]	87	0.70	17.99	1.07	2.29	72	1.06	16.60	1.37	1.79

an average premium by weighting each observation with δ confirms this. Under the weak sample trimming, the weighted average premium is −2.7% (i.e., a discount). The full sample shows huge AARs, but looking at the trimmed samples, these seem to be driven by outliers in the data. Nevertheless, the trimmed samples show AARs that are more than twice the size of the unadjusted abnormal returns—implying a large positive information effect from the announcement of private placements.

An alternative explanation ot the positive announcement period return for private placements is that higher shareholder concentration increases monitoring of management. Panel (a) in table 5.7 explores this hypothesis by dividing the sample in two groups based on whether shareholder concentration increased or decreased over the year of the private placement. There is no evidence of a differences in abnormal return in these groups. In fact, the group of issuers that experienced decreased shareholder concentration in the offering year have the largest announcement period returns. Panel (b) in table 5.7 reports the coefficients from a regression of abnormal returns on shareholder concentration and percent insider holding. Under the monitoring hypothesis we should expect less abnormal performance for highly concentrated firms, resulting in a negative γ_1. The regression provides no evidence to support this. Also note that the two-day announcement period returns are positively related to the insider holdings at the beginning of the offering year. This is not consistent with the hypothesis that private investors form an alliance with management to extract corporate control benefits.

Table 5.7
Ownership structure and abnormal announcement period returns for OSE-listed firms making private equity offerings,1984–1996

The regression variables CONC and INSIDE are shareholder concentration and insider holdings respectively. Shareholder concentration is measured as the percentage total shareholdings of the ten largest stockholders. Insider holdings include the shareholdings of the CEO and the board of directors. The regression in panel (b) uses the concentration and insider holdings from the beginning of the offering year.

(a) Change in shareholder concentration and abnormal announcement period return

	Including mergers			Excluding mergers		
	Decreased shareholder concentration	Increased shareholder concentration	Test of difference (p-value)	Decreased shareholder concentration	Increased shareholder concentration	Test of difference (p-value)
Two-day announcement return	3.02	1.16	.403	3.34	1.62	.491
Four-day announcement return	4.23	2.75	.654	5.01	3.63	.696

(b) Regression: $AR = \gamma_0 + \gamma_1\text{CONC} + \gamma_2\text{INSIDE}$

Dependent variable	γ_0	γ_1	γ_2	γ_0	γ_1	γ_2
Two-day announcement return	0.03 (.918)	−0.05 (.333)	0.097 (0.041)	0.04 (.260)	−0.06 (.287)	0.09 (.082)
Four-day announcement return	−0.01 (.831)	0.05 (.520)	0.04 (.532)	−0.00 (.999)	0.04 (.607)	0.05 (.537)

The announcement period abnormal returns for private placements reported in tables 5.5 and 5.6 are to some extent as predicted by the model developed in section 5.2. Announcing a private placement causes a positive market reaction. This is consistent with the idea that (imperfect) inspection by the private investor certifies that the issuer is a high-value company. The lack of any statistically significant discounts for the private investors may indicate that the issuer has most of the bargaining power when the offerprice is determined. This could be explained by competition among private investors to participate in private placements. The model in section 5.2 also predicts that the announcement period abnormal returns for private placements should be greater than the announcement period abnormal returns for uninsured rights offerings. Based on a two-day announcement period, the data gives no support for this prediction. At conventional levels of statistical significance, the 1.0% abnormal return for private placements (excluding mergers) is not different from the 0.95% abnormal return for uninsured rights offering. On the other hand, the adjusted four-day abnormal returns reported in table 5.6 give some indications that the announcement effect is larger for private placements than for uninsured rights.

In sum, announcing a standby rights offering seems to produce a negative market reaction, while announcements of uninsured rights and private placements are regarded as good news. Moreover, there is weak evidence of abnormal returns induced by the private placement information effect that is larger than the abnormal returns for uninsured rights offerings. These results are largely consistent with the model developed in section 5.2.

5.5 Conclusions

This chapter develops a model of equity flotation method choice by extending the theoretical developments in Eckbo and Masulis (1992). The model is specifically design to explain the choice among private placements, standby rights offerings, and uninsured rights offerings—which represent the complete range of flotation methods used on the Oslo Stock Exchange. In the model, private investors and investment banks (underwriters) perform a private inspection of the companies that seek project financing. The knowledge about firm quality acquired during the inspection allows a credible quality certification of high-value firms. Thus, these firms can raise funds at more favorable conditions than would have been possible without the inspection. Firms that are rejected by a private investor, may withdraw the issue or use uninsured rights or rights with standby underwriting. Since the rejection by the private investor is based on an informative inspection, the expected issuer-value for private placements exceeds that for uninsured rights and standby rights.

An event study of the market reactions to announcements of private placements, standby rights, and uninsured rights largely produces results that are consistent with the model. As in Bøhren, Eckbo and Michalsen (1997), announcing a standby rights offering gives a negative market reaction, while announcements of uninsured rights are regarded as good news. Consistent with the findings of Wruck (1989) and Hertzel and Smith (1993) on U.S. private equity offerings, and Kang and Stulz (1996) on Japanese offerings, the announcement of private placements on the OSE produces positive abnormal announcement period returns.

Since private placements have lower direct costs than standby rights and since the model assumes that private investors and underwriters have equal inspection ability, the model implies that in equilibrium the pool of firms using private placements have higher average value than the pool of firms that use standby rights. This prediction receives support by the evidence, in that two-day abnormal returns following announcements of private placements are positive and larger than the abnormal returns for standby rights offerings. An alternative hypothesis for the positive announcement period abnormal returns observed for private placements holds that higher shareholder concentration increases management monitoring. However, there is no evidence that the announcement effect is greater for the subsample of private placements with the greater increase in shareholder concentration over the year of the issue. Overall, these results indicate that information asymmetry between firms and investors is an important consideration when firms decide on cost efficient ways to float new equity.

Chapter 6

SEO performance: The Norwegian evidence

Chapter 4 looked at insufficient control for risk as a potential explanation for the "new issues puzzle". This chapter explores the timing hypothesis (or windows-of-opportunity hypothesis) and the underreaction hypothesis as alternative explanations for the "new issues puzzle". Empirical predictions of these hypotheses are tested using data on Norwegian private and public equity offerings. The prediction that issuers with less timing incentive should have better long-run performance than issuers with strong timing incentive and the prediction that the announcement period abnormal return should be positively correlated with the long-run abnormal return, are tested using several long-run performance measures. Neither the timing-hypothesis nor the underreaction hypothesis receive support by the reported evidence.

6.1 Introduction

This chapter looks at the market reaction to equity offerings in light of issue-timing and investor overconfidence. The timing-hypothesis builds on the notion that investors are overly optimistic about the prospects of issuing firms, and as a consequence do not fully incorporate into prices managers incentive to time an equity issue. This results in initial overpricing of issuing firms and a subsequent long-run underperformance when investors correct this initial mispricing over time. Apart from the long-run underperformance prediction, this argument implies that equity offerings where managers have less incentive to time the issue, should have better long-run stock return performance than offerings where managers have a strong incentive to time the issue. If managers behave in the interest of current shareholders, the incentive to time an issue should be less the larger the current shareholders takeup.

The overconfidence hypothesis of Daniel, Hirshleifer and Subrahmanyam (1998) is closely related, but is derived in a formal model and carries some explicit empirical predictions. The overconfidence hypothesis is based on the assumption that investors are overconfident about the precision of their private information, but not about the precision of public information. Overweighting private information relative to public information causes underreaction to new *public* information. Thus, the theory predicts that discretionary corporate events (such as equity issues) associated with abnormal announcement period returns, on average should be followed by long-run abnormal performance of the same sign as the average announcement period abnormal return, and there should be a positive correlation between the announcement period abnormal return and the long-run abnormal return.

Several empirical papers have explored different aspects of the timing- and overconfidence hypotheses. Rangan (1995) and Teoh, Welch and Wong (1998*b*) look at discretionary accruals in the years around an equity offering. The idea is that if investors are overly optimistic about the prospect of firms that issue equity, the optimism could potentially have been induced by the issuing firms through reporting of inflated earnings. Both papers find evidence of earnings management prior to seasoned equity offerings. For example, Teoh, Welch and Wong (1998*b*) find that although cash flows from operations are declining prior to the issue, the reporting of discretionary accruals leads earnings to peak around the issue. Moreover, the amount of discretionary accruals prior to the seasoned equity offering is negatively related to the post-issue long-run stock return performance. The authors view this as evidence in favor of timing and overly optimistic investors.

Cornett, Mehran and Tehranian (1998) employ a direct test of the relationship between the incentive to time an issue and the subsequent stock return performance. They study voluntary and involuntary SEOs by commercial banks. Capital regulations in the banking industry state that banks are not allowed to have total capital ratios below a certain level. If the total capital ratio falls below the regulated lower bound, a bank may need to issue new equity to raise their capital ratio. Cornett, Mehran and Tehranian (1998) define an involuntary SEO as an issue by a bank with capital ratio close to or below the required minimum ratio. If timing is driving the long-run underperformance of SEOs, we should expect to see less or no underperformance for involuntary issues. The results support the timing hypothesis, showing no abnormal three-year post issue stock return performance for the involuntary issues, while the voluntary issues show significant underperformance.

Brous, Datar and Kini (1998) perform another test of the timing- and overconfidence hypotheses. They argue that if managers are timing equity issues and investors systematically underreact to the issue announcements, we should expect to see that investors get disappointed when firms convey their post-issue earnings. That is, post-issue earnings announcement should on average be associated with negative stock price reactions. The results show no evidence of abnormal stock price reactions to the earnings announcements.

Kang, Kim and Stulz (1997) tests the overconfidence hypothesis using data on Japanese public and private equity offerings. The non-negative announcement period abnormal return to Japanese equity offerings supports the view that equity offerings are regarded as good news in Japan. Nonetheless, they document post-issue negative long run abnormal performance. Taken at face value, this is evidence against the overconfidence hypothesis.

Using data on Norwegian public and private equity offerings, this chapter provides several contributions to the long-run performance literature. First, although the Norwegian stock market is of marginal importance for international investors, equity offerings in Norway have some interesting institutional features that allow us to conduct experiments not possible on U.S. data. In particular, almost all public seasoned equity offerings in Norway are conducted using rights offers. This will allow a direct test of the timing hypothesis. Assume that managers have approximate knowledge of the fraction of an issue that is going to be taken up by current shareholders. If current shareholders will subscribe to a large portion of the issue, managers have less incentive to time the issue in order to exploit temporary overvaluation. Therefore, under the timing hypothesis, there should be less or no long-run underperformance for these issues.

Second, this chapter provides the first evidence in the literature on the correlation (or lack thereof) between announcement period abnormal returns and long-run abnormal performance. This enables a direct test of one of the predictions of the overconfidence hypothesis.

Third, as pointed out by Kang, Kim and Stulz (1997), knowledge of whether the negative long-run performance found in the U.S. also holds in other countries may shed light on whether the U.S. results are due to chance (as argued by Fama (1998).) If the underperformance found in the U.S. data is a result of chance, we should expect to find normal performance for SEOs in other countries.

The rest of the chapter proceeds as follows. The next section discusses the relationship between equity offering methods and the timing-incentive. Section 6.3.1 and 6.3.2 estimate long-run performance for private and public equity offerings using the matching firm procedure and the Jensen's alpha approach respectively. Section 6.4 concludes.

6.2 Offertype and timing incentive

Assuming that managers behave in the interest of current shareholders, they have incentive to time an equity issue with low expected current shareholder takeup to a period where the company's stocks are overvalued.[1] On the other hand, as current shareholders are expected to take up a larger portion of the new issue, the incentive to time the issue diminishes. Thus, under the timing hypothesis, there should be less or no long-run underperformance for issues with a high current shareholder takeup.

The model of Eckbo and Masulis (1992) and the model presented in chapter 5 imply that firms are more likely to choose standby rights as the flotation method if the expected current shareholder takeup is low. This relationship between takeup and flotation method is supported by the empirical evidence in Bøhren, Eckbo and Michalsen (1997). They find that the probability of choosing standby rights as the flotation method decreases statistically significant with the expected current shareholder takeup. Since managers decision whether or not to time an issue must be based on their predictions of current shareholder takeup, an issue timed to exploit overvaluation is more likely to be floated using standby rights. Thus, the timing hypothesis predicts that the long-run abnormal performance should be worse for standby rights than for uninsured rights.

Chapter 5 documents statistically significant positive announcement period abnormal returns for private placements. This is consistent with the findings of Wruck (1989) and Hertzel and Smith (1993) for private placement of equity in the U.S., and Kang and Stulz (1996) for Japanese private placements. Wruck (1989) presents evidence in favor of the hypothesis that the abnormal returns are positive because a private placement aligns the interest of managers and shareholders through more efficient monitoring or by increasing the probability of a value-enhancing takeover. Chapter 5 and Hertzel and Smith (1993) argue that the positive announcement effect for private placements may be explained by the resolution of the "underinvestment" problem pointed out by Myers and Majluf (1984). Under either of these theories, management is not able (or have no incentive) to time an issue to a period of overvaluation. Thus, the timing hypothesis predicts that the long-run performance of firms conducting private placements should be non-negative.

Table 6.1 sums up the theoretical predictions of the timing hypothesis and the underreaction hypothesis.

[1] The portion of the new shares purchased by current shareholders will be referred to as the current shareholder "takeup".

Table 6.1
Theoretical predictions of the timing hypothesis and the underreaction hypothesis

(a) The timing hypothesis

Long-run underperformance should be less for issues with a large current shareholder takeup than for issues with low takeup.

Long-run underperformance should be less for issues floated using uninsured rights than for issues floated using standby rights offers.

Long-run underperformance should be non-negative for private placements.

(b) The underreaction hypothesis

If the announcement of an equity issues is associated with abnormal announcement period returns, the long-run abnormal performance should on average be of the same sign as the announcement period abnormal return, and there should be a positive correlation between the announcement period abnormal return and the long-run abnormal return.

6.3 Results

Two procedures are implemented in order to estimate long-run abnormal performance. First, under the *matched control firm procedure*, abnormal returns are estimated using the realized stock return to a non-issuing firm matched to have similar risk characteristics as that of the issuer. Second, under the *factor model procedure*, a multi-factor return generating process is used to form expected one-period returns for both issuing firms and non-issuing matched-firms.

6.3.1 Performance using matched control firms as benchmark

The performance measures used in this section is similar to the one used by Loughran and Ritter (1995) and Spiess and Affleck-Graves (1995). For a firm that announces an issue in month t, a size- and book-to-market matched control firm is selected as follows: (1) All firms on OSE are ranked based on market capitalization (size) in month t. Then all companies with an equity capitalization in the interval $[(v_i/1.3), 1.3v_i]$ are identified, where v_i is the market value of the issuer's common stock at the end of month t. The company within this size-interval with the closest book-to-market ratio to the issuer is then selected as the matching firm, (3) the holding period returns for the issuer and the matching firm are computed between the month after the announcement day and the earlier of the three-year anniversary date and the delisting date for the issuer, (4) if the matching firm delists a new matching firm is chosen as described above but using the delisting date instead of the announcement date, and (5) if the matching firm issues equity it is treated as if it was delisted.

The first rows of panel (a) in table 6.2 report three-year holding-period return for the sample of Norwegian public and private equity offerings when the matching firm is selected based on size only. Both for public offerings and private placements, the issuing firms have lower three-year holding period returns than their corresponding matching firms. The average abnormal performance for public offerings is a statistically insignificant −7.1%, while private placements show a negative performance of about 25 percent (statistically significant at around the 8% level). However, results change when using size/book-to-market as matching criteria. The first two rows of panel (b) of table 6.2 show that both offertypes still have negative long-run abnormal returns, but with size/book-to-market matching the underperformance of public offerings is more severe than the underperformance of private placements. The average three-year holding period return for public offerings is 42.6%, while matching firms have an average holding period return of 65.2%. The −22.6% abnormal return is significant at the 10% level using bootstrapped p-values and at the 5% level using a standard t-test. The private placement sample shows negative three-year abnormal performance (−10.4%), but this is not statistically different from zero. Thus, the evidence in table 6.2 shows that there is some evidence of a "new issues puzzle" for Norwegian public equity offerings, especially when using size/book-to-market matched firms as the benchmark. Given the difference in results using size and size/book-to-market matching, the rest of the discussion of table 6.2 will mainly be based on the results using size/book-to-market matching. There are two reasons for this. First, since we know that both size and book-to-market ratio are related to the cross-section of asset returns, size/book-to-market matching should give a better match than size matching only. Second, the findings of Jegadeesh (1997) indicate that as long as one control for differences in book-to-market ratio, adding other matching criteria only affects long-run abnormal performance marginally.

Next the results of table 6.2 is viewed in light of the timing hypothesis. As argued in section 6.2, the incentive to time an issue is increasing in the current shareholder takeup. Thus, under the timing hypothesis, the long-run performance should be worse for rights offerings with below median takeup than for offerings with above median takeup. The results in the last two rows of table 6.2 lend no support in favor of this argument. The three-year holding period return for issuers with below median current shareholder takeup is 38.8%,

Table 6.2
Three-year buy-and-hold returns (%) to OSE-listed firms making public and private equity offerings, and their matched control sample, 1980–1993

The matches are chosen using size- and size/book-to-market matching. The size-matching is done using the equity market value of the issuer in the month prior to the issue announcement. Size/book-to-market matching is done by first selecting all companies that have an equity market value within 30% of that of the issuer. Then the company with the closest book-to-market value is chosen as the matching firm. Monthly book-to-market rankings in year t are created by dividing the end-of-year book-value from year $t-1$ with monthly market capitalizations for year t. Numbers in the columns marked "issuer" and "match" are computed using:

$$\frac{1}{N}\sum_{i=1}^{N}\left[\prod_{t=\tau_i}^{T_i}(1+R_{it})-1\right]\times 100$$

The p-values in the column marked $p(t)$ are p-values of the t-statistic using a two-sided test of no difference in average five-year buy-and-hold return for issuer and matching firms. The p-values in the column marked $p(N)$ are bootstrapped $p-$values of a two-sided test.

Issue type	N	Issuer	Match	Difference	$p(N)$	$p(t)$
(a) Size matching						
All public offerings	227	44.5	51.6	-7.1	0.640	0.458
All private placements	150	20.4	45.3	-24.9	0.078	0.080
Standby rights	147	31.8	45.5	-13.7	0.360	0.261
Uninsured rights	80	67.9	62.7	5.1	0.795	0.731
Below Median takeup	114	39.7	49.8	-10.1	0.548	0.461
Above Median takeup	113	49.3	53.4	-4.1	0.902	0.762
(b) Size- and book-to-market matching						
All public offerings	221	42.6	65.2	-22.6	0.080	0.033
All private placements	147	20.0	30.4	-10.4	0.302	0.390
Standby rights	143	30.9	53.1	-22.2	0.072	0.087
Uninsured rights	78	64.1	87.4	-23.3	0.296	0.193
Below Median takeup	111	38.8	56.1	-17.3	0.285	0.222
Above Median takeup	110	46.5	74.4	-27.9	0.089	0.077

while the corresponding return for matching firms is 56.1%. The −17.3% abnormal long-run return is not statistically significant at conventional levels. Issuers with above median takeup experience an even worse negative long-run performance (−27.9%), statistically significant at the 10% level using both bootstrapped and t-statistic based p-values. Moreover, using a standard t-test of difference in means, the null hypothesis of no difference in the abnormal long-run returns of the two median-divided subsamples cannot be rejected.

Under the overconfidence hypothesis, announcement period abnormal returns represent an underreaction to the release of new information. Thus, given the sign of the abnormal announcement period returns in table 5.5 and 5.6, the overconfidence hypothesis predicts that the average abnormal long-run return of standby rights should be (weakly) negative, while uninsured rights and private placements should show positive long-run abnormal returns. The standby rights offerings show a weakly negative long-run performance, which is consistent with the prediction of the overconfidence hypothesis. However, the uninsured rights and private placements do not show positive long-run abnormal returns. Thus, the results in table 6.2 lend only weak support in favor of the overconfidence hypothesis. In order to test the prediction that announcement period abnormal return and long-run abnormal return should be positively correlated, long-run abnormal returns are regressed on a constant and the announcement period abnormal returns. Table 6.3 shows that the sign of the relationship between the announcement day abnormal return and the long-run abnormal return is as predicted for standby rights offerings and for private placements, but not for uninsured rights. Given that none of the coefficients are statistically different from zero, the regressions in table 6.3 cannot support the overconfidence hypothesis.

Overall, the timing and overconfidence hypotheses receive little support from the evidence on long-run performance using the matching firm benchmark. However, as shown in chapter 4, long-run performance evaluation is very sensitive to the benchmark. Therefore, the next section explores the performance of the Norwegian issuers using a factor-model benchmark.

6.3.2 Performance using a factor-model as benchmark

Factor model procedures assume that expected returns are generated by a set of K prespecified risk factors. Following a tradition started by Jensen (1968), the abnormal return of portfolio p is estimated by regressing the returns of portfolio p on a constant term α_p and the K risk factors. The estimated constant term is "Jensen's alpha" and represents the average abnormal return on portfolio p over the estimation period. The expectation of Jensen's alpha equals zero for passively held portfolios provided the specified factor model

Table 6.3
The relationship between announcement period abnormal returns and three-year buy-and-hold returns for OSE-listed firms making public and private equity offerings, 1980–1996

The long-run abnormal returns are computed using size/book-to-market matching. Announcement period returns are computed over a four-day window ending at the announcement day. The regression model is:

$$\text{LRAR}_i = \beta_0 + \beta_1 \text{AR4}_i + \epsilon_i,$$

where LRAR is percent long-run abnormal returns for issuer i and AR4 is percent four-day announcement period abnormal return. Numbers in parenthesis are $p-$values.

Offer type	N	Average LRAR	Average AR4	Intercept (β_0)	β_1
Standby rights	135	−21.9	−0.5	−21.4 (.069)	1.12 (.488)
Uninsured rights	72	−22.3	2.1	−18.6 (.339)	−1.87 (.476)
Private placements	129	−10.8	2.9	−12.3 (.312)	0.58 (.575)

adequately captures the pervasive risk factors underlying the economy. Jensen's alpha is estimated using the K-factor model developed in chapter 3:

$$r_{pt} = \alpha_p + b'_{p0} r_{mt} + b'_{p1}(Z_{t-1} \otimes r_{mt}) + e_{pt}, \tag{6.1}$$

where r_{pt} and r_{mt} are portfolio return and returns on factor mimicking portfolios in excess of the return on the risk free asset, the K-vector b_{p0} measures average factor loadings, the KL-vector b_{p1} is designed to pick up predictable time-variation in factor loadings. The motivation for the conditional model framework is the growing evidence that expected returns are predictable using publicly available information.[2] In the presence of time-varying expected returns. an estimate of Jensen's alpha derived from an unconditional model is a biased measure of the true abnormal performance, and our conditional factor model estimation represents an attempt to correct for this bias.

The dependent variables in the factor model regressions are the monthly returns to equal- and value-weighted portfolios of security issuers. To illustrate, the value-weighted private placement portfolio is constructed as follows: Invest one dollar in the stocks of the

[2]Ferson and Harvey (1991) and Evans (1994) argue that time-variation in conditional betas for passive portfolios is economically and statistically small in the U.S. However, Ferson and Schadt (1996) find that time-varying betas are important in their measurement of the performance of managed U.S. mutual funds. Moreover, it is commonly accepted that conditional expected risk premiums tend to vary with economy-wide factors such as the business cycle.

first firm that announces a private placement. At the beginning of the first month after the second private placement announcement, the portfolio is rebalanced to include the new company using current value-weights. This process is continued as additional firms issue securities, until the first firm reaches its three-year anniversary or a firm in the portfolio is delisted, at which point it is removed and the portfolio is again rebalanced using value-weights. Dividing the sample by offering type (standby rights/uninsured rights/private placements), by the takeup (above and below median), and using equal and value weights result in ten such "issuer portfolios". Similarly, ten "match-portfolios" are constructed using the corresponding set of size/book-to-market matched firms identified in the preceding analysis. Finally, six "zero-investment" portfolios are constructed by selling issuer-portfolios short to finance a long position in the match-portfolio, and two zero-investment portfolios are constructed by going short in below median takeup issuers and long in above median takeup issuers.

The four risk factors used in the conditional expectation models are described in panel (a) of Table 6.4. The excess return on the market portfolio (RM) is computed as the difference between the return on the Oslo Stock Exchange (OSE) total index and the return on 1-month NIBOR. The term structure factor (ΔTerm) is computed as the change in the yield on long-maturity government bonds (from OECD) and the yield on 3-month NIBOR. The exchange-rate factor (ΔUSD/NOK) is computed as the change in the NOK/USD exchange rate. The oil-price factor (ΔBrent) is the change in the Brent Blend (crude oil) spot price. The three factors that are not return on portfolios (ΔTerm, ΔUSD/NOK, and ΔBrent) should ideally have been represented by factor mimicking portfolios. However, it turned out to be hard to accomplish this in a meaningful way. The factor mimicking procedures of Breeden, Gibbons and Litzenberger (1989) and Lehmann and Modest (1988) using either decile portfolios or the 30 largest companies each year to mimic the factor time-series, generated factor returns that had very low correlation with the original macro economic variables. Thus, the factor model uses the macro economic variables ΔTerm, ΔUSD/NOK, and ΔBrent directly as factors.[3]

The information variables are the return on the Morgan Stanley World Index in excess of the return on the 1-month NIBOR (MSWI excess return), the dividend yield on the

[3]When factors are not returns on portfolios, they enter the asset return generating process as deviations from their conditional expected value (See Shanken (1992)). In order to test the sensitivity of our results to the model misspecification that arises because ΔTerm,ΔUSD/NOK, and ΔBrent are not measured as the deviation from their respective expectations, the regressions are also run with ΔTerm,ΔUSD/NOK, and ΔBrent measured as the deviation from the time-series mean. The results using this specification are virtually identical.

Table 6.4
Summary statistics on risk factors and information variables used in the conditional factor models of expected returns

The excess return on the market portfolio (RM) is computed as the difference between the return on the Oslo Stock Exchange (OSE) total index and the return on 1-month NIBOR. The term structure factor (ΔTerm) is computed as the change in the yield on long-maturity government bonds (from OECD) and the yield on 3-month NIBOR. The exchange-rate factor (ΔUSD/NOK) is computed as the change in the NOK/USD exchange rate. The oil-price factor (ΔBrent) is the change in the Brent Blend (crude oil) spot price. The information variables are: The return on the Morgan Stanley World Index in excess of the return on the 1-month NIBOR (MSWI excess return), the dividend yield on the Morgan Stanley World Index (Dividend Yield), a dummy variable that takes on the value of one in January and zero otherwise (January Dummy), and real per capita growth rate of industrial production (Industrial Production Growth). Except for the January dummy, the numbers in panel (a) and (b) are percentages.

	Mean	Min	Max	Std.
(a) Risk factors in the conditional factor model				
RM	0.50	-26.59	20.20	6.29
ΔTerm	0.00	-0.29	0.36	0.06
ΔNOK/USD	0.23	-6.14	7.56	2.40
ΔBrent	0.37	-28.51	53.86	9.76
(b) Information variables Z_{t-1}				
MSWI excess return	0.31	-18.13	10.40	4.09
Dividend Yield	3.26	0.39	6.05	1.21
January dummy	8.06	0.00	1.00	27.28
Industrial Production Growth	-0.17	- 3.14	2.31	0.69
(c) Correlation matrix for risks in the conditional factor model				
	RM	ΔTerm	ΔNOK/USD	ΔBrent
RM	1.00			
ΔTerm	0.12	1.00		
ΔNOK/USD	0.07	0.02	1.00	
ΔBrent	0.14	-0.03	0.04	1.00

Morgan Stanley World Index (Dividend Yield), a dummy variable that takes on the value of one in January and zero otherwise (January Dummy), and real per capita growth rate of industrial production (Industrial Production Growth). Except for the January dummy, the numbers in panel (a) and (b) are percentages.

Table 6.5 reports the results for portfolios of standby rights offerings, uninsured rights offerings, and private placements. The rows labeled "EW issuer" or "VW issuer" contain results for equally- and value weighted issuer portfolios respectively. The rows labeled "EW M−I" contain the results for zero-investment portfolios that sell issuer portfolios short in order to finance a long position in the matching firm portfolios. The last column of table 6.5 reports the p-values for an F-test of the null hypothesis that none of the elements in b_{p1} are statistically different from zero. In other words, a test of the time-varying beta model. Except for the equally weighted standby rights portfolio, the null hypothesis cannot be rejected at the 5% level for any of the portfolios. That is, we cannot reject the hypothesis that betas are constant over time. The betas reported in the columns marked RM, ΔTerm, ΔUSD/NOK, and ΔBrent in table 6.5 and 6.6 are computed as $\hat{b}_{p0}+\hat{b}_{p1}\overline{Z}$, where $\overline{Z}$ contains the time series means of the information variables, where $\hat{b}_{p0}$ is the estimate of b_{p0}, and $\hat{b}_{p1}$ is the estimate of b_{p1}.[4] The p-values for the betas are from an F-test of the null that the scalar b_{j0p} and the L-vector b_{j1p} ($j = 1, \dots K$) are both zero.

Panel (a) and (b) of table 6.5 reports Jensen's alpha for the standby- and uninsured rights portfolios. The 0.7% average monthly abnormal return for value weighted standby rights issuers represents a three-year abnormal return of about 29%. The portfolios of rights issuers have positive alphas, indicating that these are overperforming relative to the benchmark. However, except for the value-weighted standby rights issuer portfolio, the alphas are not different from zero at conventional levels of statistical significance. The evidence of no abnormal performance for rights issuers are reinforced by the results for the zero-investment portfolios. The zero-investment portfolios have positive but statistically insignificant alphas. These results are consistent with the evidence in chapter 4 on long-run performance of SEOs by firms listed on the New York Stock Exchange (NYSE) and the American Stock Exchange (AMEX). Since both the timing-hypothesis and the overconfidence-hypothesis predict abnormal performance for the issuer portfolios, the evidence in panel (a) and (b) of table 6.5 do not support these theories.

Turning to the private placements portfolios, panel (c) shows underperformance for the equally-weighted issuer portfolio and the equally-weighted zero-investment portfolio. The

[4] The mean of the January dummy has no economic meaning. Replacing the mean with zero or one has virtually no effect on the results.

Table 6.5
Jensen's alpha for private and public equity issuer using a conditional multi-factor asset pricing model as an expected return benchmark

The four risk variables in r_m are the excess return on the market portfolio (RM) is computed as the difference between the return on the Oslo Stock Exchange (OSE) total index and the return on 1-month NIBOR. The term structure factor (ΔTerm) is computed as the change in the yield on long-maturity government bonds (from OECD) and the yield on 3-month NIBOR. The exchange-rate factor (ΔUSD/NOK) is computed as the change in the NOK/USD exchange rate. The oil-price factor (ΔBrent) is the change in the Brent Blend (crude oil) spot price. The information variables Z_{t-1} are listed in table 6.4. The model

$$r_{pt} = \alpha_p + b'_{p0} r_{mt} + b'_{p1}(Z_{t-1} \otimes r_{mt}) + e_{pt}$$

is estimated using OLS. Standard errors are computed using the heteroskedasticity consistent estimator of White (1980). The column labeled Adj.-R^2 contains adjusted R^2 for the regression. The numbers in parentheses are p-values.

		Betas at mean Z_{t-1}					
Portfolio[a]	$\hat{\alpha}_p$	RM	ΔTerm	ΔUSD/NOK	ΔBrent	Adj.-R^2	$F(Z)$
(a) Standby rights							
EW issuer	0.20 (0.551)	1.03 (0.000)	34.86 (0.000)	0.16 (0.080)	−0.05 (0.702)	0.684	0.003
VW issuer	0.70 (0.030)	1.15 (0.000)	7.55 (0.490)	0.21 (0.062)	−0.09 (0.151)	0.753	0.231
EW M−I	0.65 (0.154)	−0.19 (0.195)	−15.33 (0.104)	0.14 (0.948)	−0.04 (0.590)	0.035	0.756
VW M−I	0.34 (0.416)	−0.18 (0.050)	−5.08 (0.502)	−0.04 (0.480)	0.05 (0.205)	0.061	0.239
(b) Uninsured rights							
EW issuer	0.21 (0.730)	0.62 (0.0[illegible])	−2.95 (0.994)	−0.04 (0.631)	−0.08 (0.320)	0.198	0.563
VW issuer	0.59 (0.327)	0.72 (0.000)	−2.39 (0.942)	−0.08 (0.934)	−0.03 (0.647)	0.237	0.653
EW M−I	1.06 (0.130)	0.00 (0.005)	11.99 (0.773)	0.03 (0.791)	0.02 (0.688)	0.040	0.062
VW M−I	0.86 (0.261)	−0.06 (0.022)	8.84 (0.961)	−0.08 (0.710)	0.01 (0.423)	0.022	0.135
(c) Private placements							
EW issuer	−0.91 (0.038)	0.94 (0.000)	21.50 (0.031)	−0.08 (0.406)	0.02 (0.768)	0.567	0.361
VW issuer	0.33 (0.371)	1.16 (0.000)	5.37 (0.302)	0.01 (0.313)	−0.02 (0.385)	0.692	0.290
EW M−I	1.00 (0.051)	−0.10 (0.704)	−4.75 (0.584)	0.12 (0.294)	−0.08 (0.588)	-0.027	0.794
VW M−I	0.16 (0.782)	−0.16 (0.581)	−10.86 (0.057)	−0.13 (0.090)	−0.04 (0.472)	0.039	0.230

[a]Portfolios are either equal-weighted ('EW') or value-weighted ('VW'). The 'M−I' portfolios are zero-investment portfolios with a long position in the matching firms and a short position in the issuing firms.

Table 6.6
Jensen's alpha for portfolios of issuers with below and above median takeup using a conditional multi-factor asset pricing model as an expected return benchmark

The four risk variables in r_m are the excess return on the market portfolio (RM) is computed as the difference between the return on the Oslo Stock Exchange (OSE) total index and the return on 1-month NIBOR. The term structure factor (ΔTerm) is computed as the change in the yield on long-maturity government bonds (from OECD) and the yield on 3-month NIBOR. The exchange-rate factor (ΔUSD/NOK) is computed as the change in the NOK/USD exchange rate. The oil-price factor (ΔBrent) is the change in the Brent Blend (crude oil) spot price. The information variables Z_{t-1} are listed in table 6.4. The model

$$r_{pt} = \alpha_p + b'_{p0} r_{mt} + b'_{p1}(Z_{t-1} \otimes r_{mt}) + e_{pt}$$

is estimated using OLS. Standard errors are computed using the heteroskedasticity consistent estimator of White (1980). The column labeled Adj.-R^2 contains adjusted R^2 for the regression. The numbers in parentheses are p-values.

		Betas at mean Z_{t-1}					
Portfolio[a]	$\hat{\alpha}_p$	RM	ΔTerm	ΔUSD/NOK	ΔBrent	Adj.-R^2	$F(Z)$
(a) Below median takeup							
EW issuer	0.53 (0.188)	0.88 (0.000)	25.28 (0.000)	−0.03 (0.929)	−0.04 (0.640)	0.522	0.087
VW issuer	0.81 (0.032)	1.01 (0.000)	3.83 (0.505)	0.12 (0.647)	−0.07 (0.516)	0.629	0.006
(b) Above median takeup							
EW issuer	0.23 (0.545)	0.71 (0.000)	33.40 (0.000)	0.20 (0.007)	−0.06 (0.595)	0.443	0.018
VW issuer	0.90 (0.014)	0.82 (0.000)	10.13 (0.544)	−0.01 (0.489)	−0.08 (0.166)	0.498	0.663
(c) Zero investment portfolios (Above−Below)							
EW A−B	−0.29 (0.577)	−0.18 (0.107)	7.64 (0.536)	0.23 (0.032)	−0.02 (0.955)	0.022	0.280
VW A−B	0.06 (0.909)	−0.18 (0.005)	5.36 (0.898)	−0.12 (0.626)	−0.02 (0.900)	0.032	0.255

[a]Portfolios are either equal-weighted ('EW') or value-weighted ('VW'). The 'A−B' portfolios are zero-investment portfolios with a long position in the above median takeup portfolio and a short position in the below median takeup portfolio.

equally-weighted private placement portfolio underperforms the factor model benchmark by 0.91% on average over the sample period. This is comparable to a three-year abnormal holding period return of about −39%. The 1.0% average monthly abnormal return for the equally-weighted zero investment portfolio shows that most of this abnormal performance is generated by the short position in the issuing firms. The performance of the private placements change dramatically when the portfolio returns are computed using value-weights. The issuer portfolio now shows a statistically insignificant *positive* alpha (0.33%) and the abnormal performance of the zero-investment portfolio is dramatically reduced. This implies that the abnormal performance observed for the equally-weighted portfolio is driven by the performance of small firms. The underperformance for the small private placements could be taken as evidence in favor of the timing-hypothesis. If the managers that place equity privately have the same incentives as managers that place equity using a public offering, such that the theories of private placements discussed in chapter 5 and in section 6.2 of this chapter is wrong, the timing-hypothesis predicts a negative abnormal long-term performance also for private placements. For the overconfidence hypothesis, on the other hand, the evidence of underperformance for private placements is just the opposite of the prediction.

Panel (a) and (b) of table 6.6 report the performance of portfolios constructed based on the current shareholder takeup. Firms with below median takeup is placed in one portfolio while firms with above median takeup is placed in another portfolio. Using either equal or value-weights, both below-median and above-median portfolios show positive abnormal performance (significant at the 5% level for the value weighted portfolios.) In order to test whether or not the below-median portfolio have performance significantly different from the performance of the above-median portfolio, a zero-investment portfolio is constructed by selling the above-median portfolio short and using the proceeds to finance a long position in the below-median portfolio. As shown in panel (c) of table 6.6, one cannot reject the hypothesis that the equally and value-weighted zero-investment portfolios have normal returns.

Overall, based on table 6.5 and table 6.6, it seems fair to conclude that this section does not lend any convincing support in favor of either the timing hypothesis or the overconfidence hypothesis.

6.4 Conclusion

This chapter explores the timing hypothesis and the overconfidence hypothesis as alternative views of the market reaction to equity announcements. These theories are motivated by the "new issues puzzle", and therefore predict that the long-run abnormal stock return after equity issues should be negative. The analysis indicates a possible Norwegian "new issues puzzle" when using a matching firm benchmark. The average three-year holding period return for the sample of public offerings is 42.6%, while the associated matching firms have an average holding period return of 65.2%. However, this underperformance disappears when using a factor model benchmark. The overconfidence hypothesis has the additional prediction that the abnormal announcement period returns on average should be followed by long-run abnormal performance of the same sign. The results also fail to support this prediction.

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