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LEAPS introductions and the value of the underlying stocks

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Abstract

We examine the change in the value of the underlying stock associated with long-term option introduction. Analysis of the abnormal returns associated with LEAPS (Long-Term Equity Anticipation Security) introductions indicates a decline in firm value even after we control for the endogenous nature of the listing decision. However, the evidence does not support previously-offered explanations for the price change associated with option introductions. In particular, we do not find the predicted relations between the cumulative abnormal returns and variables associated with loosening of short sale constraints such as beta, proxies for the dispersion in investor beliefs, and change in relative short interest.

JEL classification: G12; G13; G14

Keywords: LEAPS; Option introductions; Short-sale constraints

1. Introduction

Previous studies have investigated the effect of short-term option introduction on the value of the underlying stock. These investigations have generally found that short-term option introductions are associated with a stock price increase for options introduced pre-1981 and then a price decline for options introduced in the post-1981 period. Two broad views have been put forth as an explanation for the observed price declines in the post-1981 period. One view is that the option market does not substantively alter the return characteristics of the underlying asset but

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rather the abnormal returns can be explained by selection bias. This line of reasoning argues that it is not clear why options should impact the underlying stock given that the return state space generated by the option can be replicated by other securities and therefore, options can be viewed as redundant.[1] Moreover, the choice to list an option is based on endogenous firm characteristics that must be controlled for in order to avoid finding a spurious result.

An alternate view argues that there is binding short-selling constraints that can be relaxed with the introduction of an option. The loosening of short sale constraints explains the negative abnormal returns in the post-1981 period. This view is supported by a theoretical model provided by Danielson and Sorescu (2001) as well as from empirical evidence from Figlewski and Webb (1993) and Danielson and Sorescu (2001). In the presence of a downward sloping demand curves for stocks and asymmetric costs associated with long and short positions, short sale constraints can shift the price of a stock above investors' mean expectation of its value. Previous theory (Miller, 1977) shows that binding short sale constraints can increase the equilibrium price of the stock with heterogeneous investor beliefs in a partial equilibrium framework.

In this paper, we empirically investigate whether LEAPS[2] (Long-Term Equity Anticipation Security) introductions impact the underlying stocks and whether the observed impact conforms to previous theory and evidence on the impact of the introduction of short-term options on the underlying stocks. There is very little empirical evidence on LEAPS introductions despite their popularity in the market. Short-term options are traded on all LEAPS firms prior to the introduction of LEAPS but in contrast to short-term options that expire in less than one year, LEAPS expire in as many as three years. The effects of LEAPS introductions on the underlying stock have been largely untested and can be thought of as "out of sample" relative to the collection of short-term option introductions examined by others.

We examine the change in the value of the underlying stock associated with the introduction of 378 long-term options (LEAPS) introduced during the 1990–2003 period. We control for the endogenous nature of the decision to list an option on the firm by selecting control samples matched on firm characteristics that Mayhew and Mihov (2004a) identified as important in predicting short-term option introduction. The raw results also show a statistically significant value decline between one and two percent around LEAPS introductions, depending on the window selected and asset pricing model employed. The market model abnormal returns provide larger estimates of value loss than using the Fama–French three-factor model or as compared to market-adjusted returns.

Following the insights of Mayhew and Mihov (2004a, 2004b), we investigate whether selection bias explains the negative abnormal returns. We control for endogeneity in the listing decision by selecting control samples matched on firm characteristics that Mayhew and Mihov (2004a) identified as important in predicting short-term option introductions. We find that neither of our two sets of control firms has significantly negative abnormal returns around its peer firms' LEAPS introductions. The differences in the cumulative abnormal returns for the LEAPS firms as compared to the control firms are statistically significant using multiple windows and asset pricing models, with the exception of the smallest window investigated. The totality of the evidence suggests that LEAPS introductions are associated with price declines even after controlling for the endogenous nature of the decision.

[1] For example, see Mayhew and Mihov (2004a, 2004b). See Back (1993) for a counter-view.

[2] LEAPS are options that have an expiration of at least nine months and are available to retail investors.

We investigate the predictions of Danielson and Sorescu (2001) as a plausible explanation for the observed negative abnormal returns. Given that LEAPS are popular as evidenced by the number of stocks that have LEAPS traded on them, it seems reasonable that LEAPS are not redundant securities but rather create a less expensive mechanism to occupy certain state spaces. Danielson and Sorescu (2001) predict that firms that have greater dispersion of investor beliefs and higher systematic risk will have greater price declines associated with option introductions. Using the dispersion of analysts' forecasts, stock trade volume, and return volatility as proxies for dispersion for investor beliefs, we fail to find any evidence that suggests that greater dispersion of investor beliefs are negatively related with abnormal returns around LEAPS introductions. We also do not find a relation between ex-ante systematic risk and abnormal returns. Collectively, this evidence does not support the loosening of short sale constraints explanation for the observed decline in value.

In our analysis we also investigate changes in risk characteristics of the underlying stocks as there has been a large literature on this topic for short-term option introductions. We would expect that any change in firm risk that occurs subsequent to LEAPS introductions to influence the value of the underlying stock. Previous findings on the effects on the risk of the underlying stock are mixed.[3] We find a significant decline in total return volatility and systematic risk, but the evidence indicates that this is due to selection bias as the control group firms also exhibit significant declines in risk measures. We also do not find any relation between the changes in systematic risk and the abnormal returns associated with LEAPS introductions.

The paper proceeds as follows. Section 2 discusses theory and empirical results for stock option introductions and the associated price effects on the underlying stock. Section 3 describes the data. Section 4 presents our empirical results and discusses our findings. Section 5 concludes.

2. Option introductions

Black and Scholes (1973) assume frictionless capital markets and argue that options are redundant because they can be dynamically replicated through cash and stock ownership. If true, the introduction of an option should not affect the underlying stock. Other theorists relax the frictionless capital markets assumption and argue that options represent a less costly method to achieve a desired payout vector. In this case options theoretically can have a real effect on the underlying asset. As an example, Back (1993) describes how the presence of an option can affect information flow regarding the underlying stock price dynamics even if the option can already be dynamically replicated and even if there were no previous short-sale constraints.

In this section, we discuss three potential sources for value change associated with option introduction. First, a hypothesized source of the increased value is the market completion role that options provide. Consistent with this hypothesis, there is evidence of a regime change in the early 1980s. From 1973, when options were introduced, until 1981–1982, there is evidence of positive abnormal returns around the introduction of short-term options.[4] There appears to be a regime change in the post-1981 period where negative abnormal returns are associated with option listings (Sorescu, 2000). One argument for the decline of positive abnormal returns is that index options were introduced in 1982, which reduced the market completion role provided by any individual stock option.[5]

[3] See Mayhew (2000) for a review of the literature.

[4] Conrad (1989), Detemple and Jorion (1990) and Sorescu (2000).

[5] Sorescu (2000). Detemple and Jorion (1990) find insignificant results for 1982–1986.

A second hypothesized effect has been put forward to explain the negative abnormal returns in the post-1981 period. Danielson and Sorescu (2001) argue and provide evidence consistent with the conjecture that option introduction results in a relatively greater decrease in the cost of taking a short position relative to the decrease in the cost of taking a long position. This asymmetric cost reduction results in increased selling pressure by pessimistic investors, who can now, in the presence of options, take a short position less costly. The selling pressure results in a supply shock on the market and ultimately a lower equilibrium price results.[6] They speculate that the positive effect of market completion dominated in the pre-1981 period while their evidence is consistent with the effect of relaxing short sale constraints dominating the post-1981 period. Danielson and Sorescu's (2001) model suggests that the effect of option introductions on decreasing short sale constraints might not be exhausted after the introduction of short-term options as it seems plausible that LEAPS can further decrease short sale constraints as they can provide a less expensive mechanism to occupy certain state spaces analogously to short-term options.

In order for increased short selling to have an impact on equity prices, either one of two, or both, facts must be true. One, new negative information about the value of the stock is associated with the option introduction or, second, the demand curve for equities is downward sloping. Evidence for and against downward sloping demand curves have focused on changes in indexes, particularly the S&P 500. Whether there are downward sloping demand curves for stocks remains unresolved in the literature.[7]

Danielson and Sorescu (2001) argue that high-beta stocks are the most efficient method to hedge a net long portfolio and thus, the demand to take short positions is positively related to a stock's pre-listing beta. Additionally, they argue that when investor beliefs are more disperse, the firm will experience greater negative abnormal returns around option listing. They use stock return idiosyncratic volatility, stock trade volume, and dispersion of analysts' forecasts as measures of the dispersion in investor beliefs. Their empirical results for short-term options are consistent with their predicted relations. Figlewski and Webb (1993) also provide evidence that stocks that do not have exchange traded options underweight negative information.

Finally, much of the literature on short-term option introduction focuses on the impact of option introductions on the risk of the underlying stock. To the extent that systematic risk, or unsystematic risk, is priced and is influenced by option listings, we should observe a change in price associated with option listings if option listing influences the risk level of the underlying stock. The extant literature on the change in risk associated with a short-term option introduction has generally reported a decrease in total return volatility and no change in systematic risk.[8] These findings are generally consistent through 1981. Mayhew (2000) suggests that the findings reported in earlier studies are an artifact of the methodology as the listing of an option is an endogenous event.

The exchange's choice of the firm on whose shares to introduce an option is not random, but rather based on the potential profitability for the exchange. Mayhew and Mihov (2004a) provide

[6] See Miller (1977) for the theoretical basis for this line of reasoning.

[7] See Shleifer (1986) and Wurgler and Zhuravskaya (2002), among others, for evidence that suggests that the demand shock from passive index investors is the causal factor in the observed positive abnormal returns around inclusion into the S&P 500. Denis et al. (2003) and Chen et al. (2004) for alternative explanations for the observed increase in value and thus the evidence from S&P 500 additions may not reflect a downward sloping demand curve.

[8] See Mayhew (2000) for a more comprehensive list of studies regarding the effect of short-term options on risk of the underlying stock. Mayhew and Mihov (2004a) find some weak evidence that volatility rises after introduction if the effects of endogeneity are included.

evidence that the decision to list options is endogenous with respect to firm characteristics even though the decision is made outside the firm.

3. Data

The Chicago Board Options Exchange was the first organized exchange to introduce long-term options, which they named "LEAPS" in 1990. We begin with a sample of 378 firms upon which the Chicago Board Options Exchange has introduced these long-term options during the period 1990–2003. Daily shareholder returns, the equally weighted market returns and monthly trade volume are from the Center for Research on Security Prices (CRSP). Market value of equity and balance sheet data are from Compustat.

For an introduction to be included in the sample, daily returns must be available for at least 200 of the 250 trade days preceding the introduction. We use a matching technique to control for potential endogeneity in the decision to introduce LEAPS trade on the firm. We explicitly tie the matching to market-driven factors that previous research suggests predicts the selection of a firm for ordinary option listing.

LEAPS firms are matched two separate ways. First, a control group, control group one, is defined as the group of match firms selected based upon daily stock return volatility, market value of equity, lagged annual share trade volume, abnormal volatility, and abnormal volume. For each LEAPS firm a single match firm is identified for inclusion into control group one. Control group two are firms selected based upon the same variables except the exclusion of market value of equity. Before selecting a match firm for inclusion in a control sample, some preliminary values are calculated. Each LEAPS firm is paired with all other potential match firms before one best match is selected for inclusion in the respective control sample.

The parameters used as criteria to identify the two control samples are based upon the empirical work of Mayhew and Mihov (2004a). They identify firm characteristics that appear to influence the listing decision for short-term options. We utilize their findings to identify matched firms that have similar characteristics in an attempt to mitigate any endogeneity in the selection of LEAPS firms in order to reduce the possibility of finding a spurious result. Mayhew and Mihov's out-of-sample evidence applies to our study to the extent that the firm characteristics that predict short-term option listing are also useful for predicting LEAPS listings. The same parameters seem to be reasonable given the lack of evidence suggesting that the CBOE's profit function for a LEAPS listing is any different than that of a short-term option. It is plausible that CBOE has a different profit function for short-term options as compared to LEAPS, which possibly could bias our findings; however, it is not clear why this might be true.

Mayhew and Mihov (2004a) identify two regimes for short-term options introduction, "early" and "late." During the "early" regime 1973–1980, market value of equity, share trade volume, stock return volatility, abnormal share trade volume, and abnormal stock return volatility are important characteristics for predicting ordinary option listing. In the "late" regime, the post-1981 period, firm size is no longer an important predictor of ordinary option listing. These two regimes correspond to our control groups one and two.

The Huang and Stoll (1996) matching technique is used to select match firms for inclusion in the control sample. For each LEAPS introduction a single matching firm is identified by considering all possible match firms and choosing the match firm that most closely resembles the LEAPS firm, given the match parameters. Firms upon which LEAPS were introduced are not considered as match firms. For each pairing of LEAPS firm j, with potential match firm i, we calculate the following match score, $S_{j,i}$:

$$S_{j,i} = \sum_{k=1}^{K} \left(\frac{x_{k,j}^{LEAPS} - x_{k,i}^{Match}}{(x_{k,j}^{LEAPS} + x_{k,i}^{Match})/2} \right)^2, \tag{1}$$

where $S_{j,i}$ is the match score for LEAPS firm j when paired with potential match firm i. Matching parameters x are indexed on k, where k matching variables are used, $k = 3$ and $k = 2$ when identifying control one and control two, respectively. We define $x_{k,j}^{LEAPS}$ as the value of the matching parameter k for the LEAPS firm j and $x_{k,i}^{Match}$ is the value of the matching parameter k for potential match firm i. After pairing the LEAPS firm with all potential match firms and calculating $S_{j,i}$ for each pair of firms, we select the match firm to be included in the control sample. For each LEAPS firm j we select the matching firm i for which $S_{j,i}$ is smallest and then delete all other potential match pairs.

Descriptive statistics appear in Table 1. All statistics listed are for the year ending the day immediately preceding LEAPS introduction. LEAPS firms are very large as shown by their average market value of equity being slightly over \$9 billion. We use book equity relative to market value of equity as a measure of relative value. Volatility is the standard deviation of daily returns over the last 250 trading days. Abnormal volatility is the ratio of the standard deviation for the prior 100 days relative to the standard deviation for the prior 250 days. The LEAPS statistic of 0.989 reflects a slight decline in volatility as the LEAPS introduction becomes closer.

Volume is the share trade volume in the previous year and the abnormal volume is the ratio of the previous month's trading relative to the previous year's trading. The LEAPS statistic of 0.106 is very similar to the 0.103 that Mayhew and Mihov (2004a) report for the 1973–1977 time period for short-term option introductions. Mayhew and Mihov report that this ratio increases for short-term options to 0.251 for the 1991–1996 time period. Finally, beta is estimated using a one factor market model over 250 days prior to the LEAPS introductions as is discussed in Section 4.

Table 1
Descriptive statistics

	LEAPS	Match firms	
		Control 1	Control 2
Market value of equity (000,000s)	\$9072.42	\$8810.16	\$9403.34
Book-to-market	0.476	0.449	0.459
Volatility	3.142%	3.089%	3.062%
Abnormal volatility	0.989	0.979*	0.980
Volume (000s)	3251.54	3222.18	3172.87
Abnormal volume	0.106	0.103*	0.103*
Beta	1.596	1.558	1.563

Notes. The LEAPS sample is 378 firms from 1990 to 2003. Balance sheet data is from Compustat and return data is from CRSP. All values are for the year, immediately preceding introduction. *Book-to-market* is the book value of equity divided by market value of equity. *Volatility* is the standard deviation of daily returns for the year preceding LEAPS listing. *Abnormal volatility* is the ratio of the daily standard deviation for the 100 trade days preceding LEAPS listing divided by the daily standard deviation for the year preceding LEAPS listing. *Volume* is the share trade volume for the year preceding LEAPS listing. *Abnormal volume* is the share trade volume for the month preceding LEAPS listing relative to the year preceding LEAPS listing. Systematic risk (*Beta*) is estimated using a market model with daily returns for the year preceding LEAPS listing. *Control 1* is a sample matched on volatility, abnormal volatility, volume, abnormal volume, and market value of equity. *Control 2* is matched on volatility, abnormal volatility, volume, and abnormal volume. Paired comparison T-tests are used to test the null hypothesis that the means for the LEAPS and control sample are equal.

* Indicates that the means are significantly different at the 10 percent level.

Paired comparison T-tests are used to test the null hypothesis that the means for the LEAPS and control sample are equal. As shown in the table, the differences in the statistics for three of the five parameters used for matching between the LEAPS firms and the two control groups; market value of equity, return volatility, and annual share trade volume are economically similar and are not statistically distinguishable. The means for the abnormal volatility and abnormal volume are statistically different between the LEAPS and control group one, and abnormal volume is statistically different for control group two. However, the differences do not appear to be large in economic terms. Additionally, we show that the relative valuations are similar (book-to-market) as well as the level of systematic risk (beta) are similar across the LEAPS firms relative to the two matched control groups. We conclude that the matching procedure provides two control sets of firms that have characteristics that are similar to the LEAPS sample mitigating the possibility of finding a spurious result.

4. LEAPS price effects

In this section we begin by examining the abnormal returns for the underlying stocks surrounding the introduction of a LEAPS. We then examine alternative explanations for the abnormal returns.

4.1. Event study results

The introduction date is defined as day 0. The trading days immediately preceding and following are referred to as days -1 and $+1$, respectively. The previously-documented value change associated with options has been found on the introduction date rather than the announcement date. More recently, the announcement date for the introduction of LEAPS on a new stock precedes the introduction of trading generally by two days. The only LEAPS introductions in our sample where the announcement is greater than five days preceding the introduction of trading are the first LEAPS issued on October 5, 1990 that were announced over one month in advance. Since most firms' announcements are so close to the introduction, we do not attempt to separate any announcement from introduction effect. Several alternative windows surrounding the introduction date are examined to fully capture any introduction effect and any potential announcement effect.

Expected returns are first calculated using a one-factor market model. Market model parameters are estimated using daily returns over the period defined by the days (-256, -6). We estimate the model

$$R_{i,t} = \alpha_i + \beta_i R_{m,t} + e_{i,t} \tag{2}$$

where $R_{i,t}$ is the return on firm i for day t, $R_{m,t}$ is the equally weighted market index return for day t, and $e_{i,t}$ is a normally distributed error term. Regression coefficients are estimated using Ordinary Least Squares. The abnormal return $AR_{i,t}$ is defined as:

$$AR_{i,t} = R_{i,t} - \hat{\alpha}_i - \hat{\beta}_i R_{m,t} \tag{3}$$

where $\hat{\alpha}_i$ and $\hat{\beta}_i$ are the Ordinary Least Squares estimates of the market model parameters for firm i. The abnormal returns are cumulated to define an abnormal return over a window surrounding day zero as:

$$CAR_{i,t} = \sum_{t=t_1}^{t_2} AR_{i,t} \tag{4}$$

Table 2
Cumulative abnormal returns around introduction date using multiple asset pricing models

Days	Market model CAR			Fama–French CAR			Cumulative MAR		
	LEAPS	Control 1	Control 2	LEAPS	Control 1	Control 2	LEAPS	Control 1	Control 2
(−5, +5)	−1.950%	−0.183%[a]	−0.390%[b]	−1.469%	−0.079%[b]	−0.357%[b]	−1.286%	0.119%[b]	−0.181%[c]
	(0.0001)	(0.7350)	(0.3892)	(0.0049)	(0.8850)	(0.4345)	(0.0084)	(0.8280)	(0.6976)
(−3, +3)	−1.313%	0.496%[a]	0.008%[a]	−0.913%	0.704%[b]	0.200%[a]	−0.884%	0.717%[a]	0.214%[b]
	(0.0012)	(0.2656)	(0.9830)	(0.0264)	(0.1092)	(0.5900)	(0.0263)	(0.1108)	(0.5767)
(−2, +2)	−1.046%	0.630%[a]	0.146%[a]	−0.748%	0.838%[a]	0.348%[a]	−0.742%	0.807%[a]	0.336%[b]
	(0.0026)	(0.1062)	(0.6634)	(0.0299)	(0.0300)	(0.2978)	(0.0322)	(0.0441)	(0.3237)
(−1, +1)	−0.906%	0.300%[b]	−0.069%	−0.652%	0.432%[c]	0.067%	−0.656%	0.405%	0.041%
	(0.0013)	(0.3170)	(0.7740)	(0.0196)	(0.1412)	(0.7751)	(0.0202)	(0.1924)	(0.8692)

Notes. The sample is 378 firms from 1990–2003. *Market model CAR* (Cumulative Abnormal Return) is the cumulative daily return net of a market model. Market model parameters are estimated over days (−256, −6) using an equally weighted market index. *Fama–French CAR* is the mean cumulative daily return net of a three-factor market model, using the daily risk factors from Kenneth French's website, daily firm returns are from CRSP. Model parameters are estimated over the (−256, −6) period. *Cumulative MAR* (Market Adjusted Return) is the cumulative market-adjusted daily return calculated by subtracting from the firm's return the equally-weighted market index return. Day 0 is the date of the LEAPS introduction and day −6 is the sixth trading day preceding the introduction. *Days* denotes the window of days, over which mean abnormal returns were calculated. P-value is for the two-tailed test of the null hypothesis that the mean abnormal return (CAR or MAR) is equal to zero. *Control 1* is a sample matched on volatility, abnormal volatility, volume, abnormal volume, and market value of equity. *Control 2* is matched on volatility, abnormal volatility, volume, and abnormal volume. Paired-comparison T-tests are used to test the null hypothesis that the abnormal return of the LEAPS firm less the abnormal return of the control firm is equal to zero.

[a] LEAPS and control sample abnormal return are significantly different at the 1% level.

[b] Idem., 5%.

[c] Idem., 10%.

where t_1 is the beginning day of the window and t_2 is the end day of the window.

Cumulative abnormal returns, CARs, are used as an estimate of the value change associated with LEAPS introduction (Table 2).

For the LEAPS firms, the market model abnormal returns are significantly negative at the 1% confidence level for all windows investigated. The larger windows that are more likely to capture any potential announcement effect have larger negative abnormal returns than the more narrow windows. Similar to findings on short-term option introductions, the returns directly around the listing date are significantly negative as well.

The results presented here for LEAPS are actually more pronounced then the previously reported results for the value decline around short-term option introduction. Danielson and Sorescu (2001) document a −1.464% market model abnormal return for an 11-day window around the listing date. This compares to a −1.950% abnormal return that we find for an 11-day window around LEAPS listings. Moreover, the decline in value is economically significant as a 1.950% price decline represents an average value loss of $177 million in equity value per firm.[9]

Match firm returns are also examined around the date upon which the firm with which they were matched had LEAPS introduced. The control firm's cumulative abnormal returns are not different from zero for either control group for any of the five windows and the signs are not con-

[9] Holland and Wingender (1997) find positive abnormal returns for the period 1990–1991. Their results for 3- and 7-day windows are marginally significant with cumulative abnormal returns less than 1%. Our results suggest that this regime of positive abnormal returns was short lived.

sistently positive or negative. The cumulative abnormal returns for the control firms, presented in Table 2, range from −0.183 to 0.630% for control one and from −0.390 to 0.146% for control two.

Paired comparison T-tests are used to test the null hypothesis that the abnormal return of the LEAPS firm less the abnormal return of the control firm is equal to zero. As indicated in Table 2, the abnormal returns for LEAPS firms are significantly lower for the 5-, 7-, and 11-day windows. For the 3-day window, we find a significant difference between LEAPS firms and control group one, but not a significant difference for control group two. These results suggest that LEAPS firms experience a value loss around the listing date and there is no value loss associated with the introduction for the match firms.

Returns in excess of a three-factor Fama–French model are examined for additional robustness. Fama–French abnormal returns are calculated in a similar manner to the market model abnormal returns, except the expected returns are calculated using the Fama–French three-factor model. The daily risk factors are from Kenneth French's website. The following model is estimated:

$$R_{i,t} = \gamma_{0i} + \gamma_{1i} RMRF_t + \gamma_{2i} SMB_t + \gamma_{3i} HML_t + \varepsilon_{i,t} \tag{5}$$

where γ_{0i} is a constant, RMRF is the excess return on value-weighted aggregate market proxy, SMB is the return on a value-weighted zero-investment factor mimicking portfolio for size and HML is the return on a value-weighted zero-investment factor mimicking portfolio for market-to-book, where t indexes trade days. The Fama–French abnormal return is defined as:

$$FFAR_{i,t} = R_{i,t} - \hat{\gamma}_{0i} - \hat{\gamma}_{1i} RMRF_t - \hat{\gamma}_{2i} SMB_t - \hat{\gamma}_{3i} HML_t \tag{6}$$

where $\hat{\gamma}_{1i}$, $\hat{\gamma}_{2i}$ and $\hat{\gamma}_{3i}$ are the estimated coefficients, the model is estimated over days (−256, −6) using ordinary least squares. The cumulative Fama–French abnormal return is defined:

$$FFCAR_{i,t} = \sum_{t=t_1}^{t_2} FFAR_{i,t} \tag{7}$$

where t_1 is the beginning day of the window and t_2 is the end day of the window.

LEAPS firms experience non-zero Fama–French cumulative abnormal returns for all of the four introduction windows examined and are significant at least at the 5% confidence level. The magnitudes of the negative abnormal returns are consistently smaller than for the market model. The −1.469% negative abnormal return for the 11-day window is very similar to previous findings for short-term option introductions and is significant at the 1% confidence interval. In summary, Fama–French abnormal returns are between five and eight basis points higher per day than are the market model abnormal returns.

Similarly to the results for the market model, the abnormal returns for LEAPS firms are significantly lower at least the 10% confidence level than both control groups for the 5-, 7-, and 11-day window and significantly lower than control group one for the 3-day window. The three-factor results provides consistent evidence with the market model results that negative abnormal returns are associated with LEAPS listings.

We also include market-adjusted abnormal returns. The Market Adjusted Return (MAR) is calculated:

$$MAR_{i,t} = R_{i,t} - R_{m,t}, \tag{8}$$

where $R_{i,t}$ and $R_{m,t}$ are defined as previously, and MAR is the market-adjusted return and

$$CMAR_{i,t} = \sum_{t=t_1}^{t_2} MAR_{i,t} \tag{9}$$

is the cumulative market-adjusted return over days t_1 through and including day t_2.

The results are generally consistent with the market model and three-factor results. LEAPS firms have significantly negative abnormal returns with p-values below 5% for all windows investigated. Interestingly, the market-adjusted abnormal return for the 11-day window of negative 1.587% for a LEAPS listing is three times the magnitude of the negative 0.500% market-adjusted abnormal return that Danielson and Sorescu (2001) report is associated with the listing of a short-term option. As in the case of market model and Fama–French abnormal returns, the 5-, 7-, and 11-day abnormal returns for LEAPS firms are significantly smaller as compared to each control group.[10]

4.2. Risk

In this section we examine the change in total return volatility and systematic risk subsequent to the LEAPS listing since any potential relation between LEAPS introductions and risk changes for the underlying stocks should have a direct influence on the stocks' value. We test the null hypothesis that the change in volatility and beta is equal to zero, and report p-values for the two-tailed tests, in Table 3. The change in risk is examined from year 0 to years +1, and the p-values for the T-test of the change in the means are shown in column (3).

LEAPS firms exhibit a significant decline in volatility consistent with previous findings on short-term option introductions (Skinner, 1989 and Conrad, 1989). However, both control groups also show a significant decline in volatility. Paired-comparison T-tests are used to test the null hypothesis that the change for the LEAPS and the change for the control sample are equal. P-values are reported in column (4). The decrease in volatility for the LEAPS firms is not statistically different for the LEAPS than for either control sample. The evidence suggests that LEAPS introductions are not associated with a change in volatility for the underlying stock.

We examine beta changes during the period surrounding LEAPS introduction. LEAPS firms exhibit beta declines significant with a p-value of 0.016. In unreported results, this decline remains statistically significant through at least year 3 and the inferences are not changed by examining the median decline in beta. The short-term option literature has generally found no relation between listings of short-term options and changes in beta. Our results suggest that this is not the case for LEAPS firms. Betas for LEAPS firms are large at the time of listing and decline subsequent to LEAPS introductions.

However, as we have already discussed, it is necessary to investigate the differences between LEAPS firms and the control samples before any inferences can be confirmed. As shown in column (4), the changes in the means for beta also decline for each control sample although not significantly so. Paired-comparison T-tests are used to test the null hypothesis that the change for the LEAPS and the change for the control sample are equal. P-values are reported in column (4).

[10] In unreported results, we investigate two-week windows before and after the windows examined in Table 2 (-15 to -6 and $+6$ to $+15$) using market model, market-adjusted, and Fama–French, as well as using various estimation periods for the betas. We do not find any evidence of a significant change in value net of the control firms in either the -15 to -6 day window or the $+6$ to $+15$ day window.

Table 3
Changes in total return volatility and systematic risk subsequent to LEAPS introductions

	Year 0 (1)	Year 1 (2)	T-test P-value vs. 0 (3)	T-test P-value vs. LEAPS (4)
Volatility				
LEAPS	3.142%			
Control 1	3.089%			
Control 2	3.062%			
Volatility change (vs. year 0)				
LEAPS		−0.249%	0.000	
Control 1		−0.301%	0.000	0.442
Control 2		−0.264%	0.000	0.942
Beta				
LEAPS	1.596			
Control 1	1.558			
Control 2	1.563			
Beta change (vs. year 0)				
LEAPS		−0.085	0.016	
Control 1		−0.051	0.989	0.211
Control 2		−0.045	0.165	0.369

Notes. The sample is 378 firms from 1990 to 2003. Return data is from CRSP. *Volatility* is the average standard deviation of daily returns for the year. *Volatility change* is the volatility for year-one risk less year-zero volatility. Systematic risk (*Beta*) is estimated using a market model with daily returns for the year ended the day preceding LEAPS listing. *Beta change* is the beta for year one less year zero beta. *Year 0* is the year immediately preceding introduction; *Year 1* is the year immediately subsequent to the introduction. The p-value for the two-tailed T-test of the hypothesis that the change in volatility and beta are equal to zero is reported in column (3). Paired-comparison T-tests are used to test the null hypothesis that the change for the LEAPS and change for the control sample are equal. P-values are reported in column (4). *LEAPS* is the sample upon which LEAPS were introduced. *Control 1* is a sample matched on volatility, abnormal volatility, volume, abnormal volume, and market value of equity. *Control 2* is matched on volatility, abnormal volatility, volume, and abnormal volume. Returns for at least 200 trading days both before and after introduction must be available for inclusion in the sample.

We cannot reject the null hypothesis that the change in LEAPS net of the control group is equal to zero.

The pre-listing beta in our study is 1.596. It is not clear why exchanges appear to prefer high beta stocks for LEAPS listings, or similarly, prefer stocks with characteristics that are correlated with high betas. One possibility is Danielson and Sorescu's (2001) argument that investors have a higher demand for short positions in high-beta stocks. The matching firms also have relatively high betas. Thus, comparing LEAPS firms relative to the control groups has the effect of mitigating the probability of generating a spurious result due to a potential mean reversion. The evidence presented in Table 2 is consistent with the argument that LEAPS have no effect on systematic risk of the underlying stock despite the apparent decline in betas for LEAPS firms.

4.3. Value changes and the underlying stocks

Table 2 presents evidence that LEAPS listings are associated with negative abnormal stock returns. Table 4 presents the results on whether the abnormal returns are consistent with prior theory. We examine the implications of Danielson and Sorescu (2001) who argue that the demand to take short positions is positively correlated with high beta stocks and the demand is

Table 4
Regressions on cumulative market-adjusted return to test predictions of the loosening of short-sale-constraint hypothesis

	Market-adjusted return (−5 to +5)				
	(1)	(2)	(3)	(4)	(5)
Constant	−0.058	−0.059	−0.031	−0.049	−0.068
	(0.225)	(0.219)	(0.561)	(0.351)	(0.346)
Beta	−0.005	−0.004	−0.010	−0.002	0.000
	(0.437)	(0.601)	(0.211)	(0.822)	(0.978)
Volatility	0.225	0.160	−0.032	0.131	0.046
	(0.624)	(0.741)	(0.950)	(0.828)	(0.937)
Size	0.006	0.006	0.002	0.004	0.007
	(0.265)	(0.277)	(0.679)	(0.466)	(0.390)
Book-to-market	−0.004	−0.004	−0.005	−0.004	−0.001
	(0.597)	(0.592)	(0.463)	(0.622)	(0.915)
ΔBeta (0 to +1)		0.004			
		(0.675)			
Volume			1.925		
			(0.246)		
Dispersion				−0.018	
				(0.363)	
ΔRSI					0.413
					(0.178)
One-month					−0.002
					(0.902)
Observations	341	341	340	309	181
Adjusted R-squared	0.000	−0.003	−0.001	−0.009	−0.015

Notes. The sample is 378 firms from 1990 to 2003 with the necessary data. Balance sheet data is from Compustat and return data is from CRSP. The dependent variable is the cumulative market-adjusted return calculated over days (−5, +5) where day 0 is the date of the LEAPS introduction. Systematic risk (*Beta*) is estimated using a market model with daily returns for the year preceding LEAPS listing. *Volatility* is the standard deviation of daily returns for the year preceding LEAPS listing. *Size* is the natural logarithm of the firm's total book assets in year 0. *Book-to-market* is the book value of equity divided by the market value of equity. *ΔBeta* (0 to +1) is the firm's beta estimated immediately prior to introduction less the beta estimated over the year beginning the day following introduction. *Volume* is the share trade volume in the year preceding LEAPS listing. *Dispersion* is the ratio of the standard deviation of analysts' forecasts relative to the absolute value of the mean of the analysts' forecasts. *ΔRSI* is the change in relative short interest for the month subsequent to the LEAPS introduction minus the relative short interest in the month subsequent to LEAPS introduction, scaled by shares outstanding. *One-month* is an indicator variable equal to 1 if the RSI value is available within one month subsequent to the LEAPS introduction, 0 otherwise. P-values are reported in parenthesis below the coefficient.

greater when beliefs are more disperse. As discussed above, their hypothesis that loosened short sale constraints are associated with option listing implies a negative relation between abnormal returns and pre-listing beta as well as a negative relation between a dispersion of investor beliefs and abnormal returns.

We start our analysis by including variables in model (1) that are associated with risk and valuation of the underlying stock. We use the market-adjusted return is used as proxy for the LEAPS introduction return. The market-adjusted return is chosen to assure orthogonality between the systematic risk estimate and the introduction return. Market model abnormal returns are inappropriate as a dependent variable when beta is included as an independent variable since beta is included in the market model estimation of abnormal returns which would increase the likelihood of finding a spurious relation. Market-adjusted returns introduce a positive bias of a lesser magnitude than the negative bias associated with market model returns.[11]

Additionally, we also orthogonalize volatility relative to the firm's beta since beta and volatility are correlated. In Eq. (10), we report results of the cross-sectional regression using estimates from the 250 day pre-listing interval, with the p-value appearing in brackets below the estimated coefficients:

$$\begin{aligned} \mathit{TotalVolatility}_i = \underset{\substack{0.01352\\ [0.0001]}}{\alpha} + \underset{\substack{0.01128\\ [0.0001]}}{\gamma\hat{\beta}_i} + \upsilon_i, \end{aligned} \tag{10}$$

Adjusted R-squared $= 0.3757$.

The error from this regression is the measure of volatility not explained by the firm's beta and is the orthogonalized volatility in the CMAR regressions that are reported in Table 4. A similar regression is estimated for the beta and total return volatility estimated over the 100 day pre-listing interval. The coefficient estimates are $\alpha = 0.01309$ (p-value < 0.0001) and $\gamma = 0.01026$ (p-value < 0.0001). It appears that the estimates are not particularly sensitive to the window employed.

The dependent variable in the regressions, in Table 4, is the CMAR for the $(-5, +5)$ interval.

The independent variables in model (1) are: year 0 beta (Beta), year 0 stock return volatility (Volatility) and the natural logarithm of market value equity (Size) and Book-to-market. Arguably, all of these variables could have a relation to investors' desire to short sale. We include variables that would seem to be correlated with the degree of the short sale constraint and thus, should have the largest degree of a demand shock associated with LEAPS introductions. We include size since it seems reasonable that smaller firms may be more difficult to short and book-to-market as a measure of relative valuation as investors desire to short may be correlated with the valuation of the firm. Under markets where short sale constraints are not binding, both of these variables should have coefficients not different from 0.

In model (1) we find no evidence in support of the loosening of short-sale-constraint hypothesis. Beta and volatility both have very high p-values and volatility has the incorrect sign relative to the predictions of the loosening of short-sale-constraint hypothesis. It can be argued that since market-adjusted returns have an upward bias on a beta coefficient, the true relation is not being observed. However, Danielson and Sorescu (2001) find a significantly negative coefficient on beta when market-adjusted abnormal returns is the dependent variable. Our finding in model (1) suggest that LEAPS listings do not appear to have the same relation between abnormal returns and beta or abnormal returns and volatility as has been reported for short-term option listings. In unreported results, the coefficient of beta and volatility are not particularly sensitive to the inclusion or exclusion of the other variables in the model and remain insignificant in all cases.

The anticipated change in risk, if any, should be associated with value change. It is possible that the lack of a significant net change in beta in Table 3 hides economically important cross-sectional differences. Alternatively, if the control groups are incorrectly specified, the raw results of a decline in beta are more indicative of the true impact of LEAPS listings. In model (2), independent variables include, in addition to all those in model (1), the change in beta (year 0 less year +1 beta). If the market can correctly anticipate changes in beta in the

[11] See Danielson and Sorescu (2001, p. 466) for the results of their simulations.

cross section, then we expect to observe a positive relation between the firm's decline in beta and abnormal returns surrounding the listing. We find that the coefficient is positive but not significant.

In models (3) through (5), we include stock trade volume, dispersion in analysts' beliefs and change in relative short interest to further investigate the predictions of the loosening of short-sale-constraint hypothesis. In model (3), we start with the variables from model (1) and include share trade volume. As has been argued by previous authors such as Jones et al. (1994) and Danielson and Sorescu (2001), volume is related to the heterogeneity of investor beliefs. Therefore, a negative relation between abnormal returns and volume is consistent with the loosening of short-sale-constraint hypothesis. In model (3) we include volume and find that the coefficient for volume is insignificantly different than 0. In unreported results, we substitute abnormal volume for volume and still fail to find a relation with abnormal returns.

In model (4) we use analysts' forecast dispersion (Dispersion) as a measure of the dispersion of investor beliefs. We use First Call data for our analysts' estimates. Our measure of dispersion equals the standard deviation of analysts' forecasts divided by the absolute value of the forecast as done in Diether et al. (2002). The coefficient for dispersion is insignificantly negative. The coefficient has the sign predicted by the loosening of short-sale-constraint hypothesis, but the lack of significance does not provide any evidence that this proxy for the dispersion of investor beliefs is related to the observed abnormal returns.

Finally, in model (5) we include the change in relative short interest (ΔRSI). We calculate this variable in a manner similar to that employed by Danielson and Sorescu (2001). As discussed in Danielson and Sorescu, the relative short interest data is reported monthly. We take the change in short interest from the first observation preceding the LEAPS introduction by at least 16 days as compared to the first observation subsequent to the LEAPS introduction by at least 16 days scaled by the number of shares outstanding. Thus, using day 0 as the listing date, the pre-listing relative short interest observation is within the -46 to -16 day window and the post-listing relative short interest observation is within the $+16$ to $+46$ day window.

We also include a one-month indicator variable where we assign the indicator a value of one if the relative short interest is reported within 30 days of the listing date (i.e., from day $+16$ to day $+30$) and otherwise assign a value of zero when the relative short interest is reported during the $+31$ to $+46$ day period. This indicator variable is included to control for any potential drift in relative short interest. We find that there is a 0.405% increase in relative short interest, which is a significant increase at the 5% confidence level. Figlewski and Webb (1993) and Danielson and Sorescu (2001) both provide evidence that stocks with short-term options have greater levels of relative short interest. The authors of these studies argue that an increase in relative short interest is consistent with a loosening of short sale constraints. Our finding that relative short interest increases around LEAPS introductions is consistent with a loosening of short sale constraints.

Danielson and Sorescu (2001) further argue that change in RSI will be negatively related to abnormal returns. They do not argue causality between abnormal returns and change in RSI as the underlying causation are the dispersion of investor beliefs and the loosening of short sale constraints. Model (5) on panel B does not show a significant relation between the change in relative short interest and abnormal returns. Our setup of the model, in Table 4, is different from that in Danielson and Sorescu (2001), since we use CARs as the dependent variable whereas Danielson and Sorescu use change in RSI as the dependent variable. We replicate their findings using our

LEAPS sample in Eq. (11), with the p-values in brackets below the estimated coefficients:

$$\begin{array}{llllll} \Delta RSI_i = \alpha + & \gamma_1 MAR(-5,5)_i + & \gamma_2 Ab.Volume_i + & \gamma_3 OneMonth_i + & \upsilon_i, & \\ & 0.028 & 0.127 & -0.001 & & (11) \\ & [0.116] & [0.000] & [0.682] & & \end{array}$$

Adjusted R-squared $= 0.076$.

The coefficient for abnormal volume is significant and positive, which is consistent with Danielson and Sorescu (2001). This finding seems reasonable in that an increase in total volume is related to an increase in bearish activity. We do not find a significant relation between the 11-day market-adjusted returns (MAR(−5, 5)) but the sign of the coefficient is positive, opposite of the prediction of the loosening of short-sale-constraint hypothesis.

For additional robustness, in unreported results, we re-analyze the five models from Table 4 using a probit model where the dependent variable equals one if the 11-day market-adjusted returns are positive, zero otherwise. The results of the five probit models do not alter our inferences and do not provide any support for the loosening of short-sale-constraint hypothesis. Additionally, we also examine the 5-day window as the dependent variable for all of the models in Table 4 and the results are not substantively different than those reported using the 11-day window.

In summary, Table 4 does not present any evidence that increasing selling pressure or risk changes influences value in the underlying stock associated with the introduction of a LEAPS. We do find evidence that relative short interest increases subsequent to LEAPS introductions which does provide some evidence that short sale constraint have been loosened. However, in Table 4 we are unable to find any of the cross-sectional relations that the loosening of short-sale-constraint hypothesis predicts in order to explain the value changes that we document. This is notable, since other studies document negative abnormal returns of similar or less magnitude associated with short-term option listings relative to our findings for LEAPS listings.

There are four plausible explanations for our findings for LEAPS introductions relative to Danielson and Sorescu's (2001) findings that the arrival of pessimistic traders explains the value decline associated with the introduction of short-term options. First, is that LEAPS introductions is an out of sample test and the results suggest that the relations found for short-term options are spurious. Given that we find negative abnormal returns of larger magnitude than previous findings of short-term options and do not appear to have similar relations to variables such as beta, volume, and change in RSI, this explanation seems plausible.

Second, that LEAPS are redundant to short-term options as pessimistic investors may be taking long-term short positions with an intertemporal sequence of short-term options. Third, LEAPS may lessen short sale constraints but not offer an arbitrage opportunity to pessimistic investors if their information pertains to short-term rather than long-term mis-pricing. Fourth, since our proxies for dispersion of investor beliefs are estimated with noise, it is possible that the loosening of short-sale-constraint hypothesis explains the abnormal returns but is not clearly identifiable with our sample.

We cannot reject the second or third explanations. However, neither of these explanations explains the observed price declines that are documented in Table 2. We can only conclude that our findings do not provide evidence in support of the loosening of short sale constraints as the causal factor for the price decline associate with LEAPS introductions. We are left with an unexplained price decline that should be the subject of further research.

5. Conclusion

This paper examines the evidence on the relation between the introduction of a long-term single stock option and the associated change in value of the underlying stock. Our findings have some similarities to the evidence on the introduction effect of ordinary, short-term options; however, we find some interesting, unexplained differences. Previous evidence on short-term option introduction generally reports either a positive or negative introduction return, depending largely on whether or not the introduction precedes or follows 1981.

We find evidence that net of a market model the introduction return is negative, consistent with most of the reported results on short-term options listed after 1981. For additional robustness we examine the introduction return net of a three-factor model. We also compare the introduction return to the returns of two sets of control firms. Using a three-factor model, we find that the value decline is smaller by 25 to 30%. We conclude that a richer model of asset pricing explains some of the value decline, yet there still is evidence of a value decline net of the control samples.

We find negative abnormal returns of similar, or larger, magnitude than those previously documented for short-term options. However, we find only limited evidence that is consistent with previously offered explanations for the reported value decline for short-term options in our LEAPS sample. In particular, we find little evidence to support Danielson and Sorescu's (2001) hypothesis that the observed value decline associated with short-term option introductions occurs as pessimistic investors arrive, coinciding with a decline in short-sale constraints. Our evidence that the abnormal returns are unrelated to beta and are unrelated to multiple proxies for the dispersion of investor beliefs is inconsistent with this conjecture as an explanation for the observed price decline associated with LEAPS introductions.

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