Options on Initial Public Offerings

Abstract

We analyze the determinants and consequences of option listing on IPO firm stock. We find that options are listed earlier on venture-backed and lower-reputation underwriter IPOs. We find a significant decrease in stock returns immediately after option listing, persisting for a year. Analyzing the determinants of this equity underperformance, we find a permanent threefold increase in short-interest ratio and aggressive insider selling in IPO equity following option listing. Further, buying newly listed put options on IPO stock yields up to 6.3% monthly excess returns. Overall, we find that option listing relaxes short-sales constraints on IPO equity, thereby correcting short-run overvaluation.

JEL Classification: G14; G12; G24.

Keywords: IPOs; Option listing; Long-run stock returns; Short-sale constraint

1 Introduction

A large percentage of firms do not have options listed on their equity. On the other hand, for some firms, options are listed on their equity as early as one week after the initial public offering (IPO). The objective of this paper is to study the determinants and consequences of option listing on IPOs. A number of theoretical papers have argued that the listing of options on a firm's stock can have important economic effects on its equity, and consequently, on the payoffs of firm insiders and outside investors. For example, Danielsen and Sorescu (2001) make use of an extension of the model of Jarrow (1980) to argue that the listing of options on a firm's equity relaxes the short sale constraint on its stock, leading to a fall in the firm's share price in a setting with heterogeneous beliefs among investors (see also Miller (1977)). Similarly, Easley, O'Hara, and Srinivas (1998) develop a model in which the options market provides an important avenue, in addition to the stock market, through which those with private information can exploit this information. Their model shows that there are important linkages between the option and stock markets. It predicts that option demand will contain information about the subsequent stock returns on the underlying stock.

The above hypothesized effects of option listing have three important implications in the context of IPOs. First, while there is some degree of short-selling in the equity of newly public firms, it is likely that short-sale constraints are more tightly binding on the equity of newly public firms in the months immediately following their IPOs than on the equity of seasoned firms.[1] Second, it has been argued that the widely documented long-run underperformance of IPOs is partly due to the presence of short-sale constraints (see, e.g., Ritter (1991)). This means that the relaxation of short-sale constraints due to the listing of options is likely to have much larger effects on the stock returns of newly public firms.[2] Third, investors in the equity of recently public firms are likely to suffer from a much greater extent of information asymmetry during the first few quarters after their IPO compared to the information asymmetry facing investors in the equity of seasoned firms. This is due to the fact that very little credible information about their performance and future prospects (such as audited financial statements) is available prior to these firms going public. This again means that the listing of options will have a greater effect on the stock returns of newly public firms compared to the corresponding effect for seasoned firms. Surprisingly, however, the economic effects of option listing on the IPO firm's equity has not been studied prior to this paper.

[1] As explained by Geczy, Musto, and Reed (2002), short selling stocks in the IPO aftermarket is difficult because there is a restricted supply of shares in the immediate aftermarket, the lockup agreement, the prohibition on banks in the IPO syndicate from lending shares in the first 30 days, and the relatively small public float in most IPOs.

[2] Banerjee and Graveline (2014) show that when a security available for trading is scarce, i.e. an equity of recent IPO firms, the derivatives market can reduce the price distortion of the underlying security by relaxing its scarcity.

We address four important questions surrounding the listing of options on the equity of firms that have recently gone public. First, what factors influence the time-to-list options after the IPO? In addition to variables that have been documented to impact exchanges' option listing decision in the case of seasoned firms (see Mayhew and Mihov (2004)), we examine whether characteristics related to the IPO such as venture backing and underwriter reputation have an effect on the time to option listing. Second, what is the short-term and long-term effect of option listing on underlying stock returns of IPO firms? A related question is how various IPO characteristics influence the magnitude of price effect that option listing on the IPO firm's long-run stock returns.

The third question is how does the listing of options affect the returns on the IPO firm's stock? We address this question by examining the following three variables. The first is the change in the short interest ratio in the IPO firm's equity around option listing, which allows us to examine whether option listing indeed increases the extent of short selling on an IPO's stock by effectively relaxing its short sale constraint, as suggested by Danielsen and Sorescu (2001). Next, we study whether insiders aggressively sell their shares in the IPO firms immediately after options have been listed in anticipation of a potential slide in the firm's share price.[3] The last variable that we study is the relative put-call demand of newly listed options on IPO stocks, which allows us to examine whether investors with private information are making use of the options market to exploit their private information. The model of Easley, O'Hara, and Srinivas (1998) predicts that, if informed traders use the options market to exploit their private information on the expected future performance of IPO firm's stock, put prices should be expensive relative to call prices once options are listed.

The fourth question we address is whether investors can make abnormal profits after accounting for transaction costs by investing in newly listed options on the IPO firm's stock? This question is important in its own right, since the existing literature seems to indicate that strategies such as short-selling the stock soon after IPO do not yield abnormal profits: see, e.g., Edwards and Hanley (2010). Answering this question allows us to examine why there may be a significant demand for option listing on the IPO firm's equity: if trading strategies in the firm's options yield abnormal profits, the demand for such options is likely to be greater.

We address the above four questions by making use of a sample of IPOs from January 1996 to December 2008 from the Securities Data Corporation's (SDC) new issues database. We obtain option pricing and other option-related information from the Ivey OptionMetrics database which, at the time of this writing, has options data till December 2011. We exclude

[3]The "demand curve" hypothesis assumes a downward sloping demand curve for equity. Assuming such a demand curve for the IPO firm's stock, continued insider selling in the firm's shares subsequent to option listing will be associated with a permanent decline in the firm's share price, leading to negative long-run stock returns upon option listing. Further, if this insider selling persists in the months after option listing, the negative stock returns to the IPO firm's equity will persist in these months as well.

IPOs after 2008 to allow at least 3 years to follow their option listing history.

We now summarize the results of our empirical analysis, organizing them under the four questions that we mentioned above. Our first set of results regard how various IPO characteristics influence when options are listed after the firm has gone public. We apply an extended Cox hazard rate model to analyze the time-to-list options after an IPO. We find that after controlling for the various equity characteristics such as trading volume, volatility, and firm size shown in Mayhew and Mihov (2004) to affect exchanges' listing decision, there are four IPO characteristics that significantly affect the time to option listing: venture backing, underwriter reputation, initial return (IPO underpricing), and IPO proceeds. We find that venture backed firms have options listed significantly sooner than non-venture backed firms in the post-IPO period (i.e., time-to-list is almost twice shorter), and the larger are the IPO proceeds, the shorter is the time-to-list. In contrast, IPOs underwritten by higher reputation underwriters and those with larger initial returns are associated with a significantly longer time to option listing, though the economic effect of the latter variable is rather small.[4]

Our second set of results concerns the effects of option listing on an IPO firm's stock. We find that equity returns of IPO firms decrease in the short-run as well as in the long-run following option listing. For firms that have options listed on their equity within three years after IPO, we find the cumulative average abnormal returns (CAR) in the event window of [+2,+12] weeks after option listing is – 14.51% lower than the CAR in the event window of [-12 to -2] weeks before option listing, which is statistically significant at the 1% level. The IPO firm's equity continually underperforms in the longer run resulting in the CAR of – 18.14% in the event window [+13,+52] after options have been listed. This finding is evident from Figure 2, which plots the weekly CARs from 12 weeks before to 52 weeks after option listing; the option listing week is roughly when the cumulative abnormal returns to the IPO firm's equity is the greatest. We further split our sample of IPO firms based on when options are listed after their IPO issuance date and find that the above stock return underperformance is greatest for firms which had options listed on their stock within the first six months of going public. Nevertheless, we observe statistically significant equity return underperformance in all subsamples during the [+13,+52] week event window subsequent to option listing.

We also examine the effect of option listing on stock returns using a regression analysis. Our dependent variable for this analysis is the weekly return on the IPO firm's stock around option listing, measured over two event windows: a shorter event window of [-12, +12] weeks, and a longer event window of [-12, +52] weeks around the option listing date. After controlling for time varying firm-specific variables such as momentum, size, volume, and volatility, we find that option listing has a significantly negative impact on the weekly return on the underlying

[4]We find that the time to first lock-up expiration does not have a significant effect on the time to option listing; neither does a dummy variable capturing whether or not a firm's IPO has a lock-up agreement.

IPO firm's stock returns, regardless of whether these are measured over the shorter or longer event window. More specifically, the baseline regression results suggest that relative to the weeks before options are listed, IPO's stock returns, on average, decrease by 134-160 basis points per week after options have been listed. Further, two IPO characteristics have a statistically and economically significant effect on the magnitude of this negative option listing effect: venture backing, which results in a more negative option listing return, and IPO underwriter reputation, which results in a less negative option listing return.

In order to control for the possibility that the negative stock returns that we observe following option listing are driven by exchanges' decision to list options on IPO firms with certain characteristics, we apply a propensity score matching analysis. We construct a control sample of propensity-score matched IPO firms that have no options listed. Each option-listed IPO firm (Treated group) is matched to an IPO firm that does not yet have options listed (Control group) based on their propensity score. The propensity-score matching is implemented along various IPO characteristics as well as time-varying dimensions such as size, trading volume, volatility, and momentum. Using the matched sample, we estimate the difference-in-difference regressions. We find that our results are robust to controlling for any potential endogeneity arising from the exchanges' choice of firms with certain characteristics on which to list options.

Our third set of results deal with the three hypotheses regarding the possible causes for IPO underperformance following option listing. We find a significant and permanent increase of about 200% in the short-interest ratio for IPO firms' equity in the month of option listing relative to control-group firms, indicating that option listing leads to a significant relaxation of the short sale constraint on IPO firms' equity. We find that, while there is no difference in the relative change in insider holdings in the IPO firm and the matched control firm in the months prior to option listing, the difference between insider holdings in the two groups become significant after option listing. Thus, option-listed IPO firms experience a statistically and economically significant decline in insider equity holdings during the months after option listing. Under the downward sloping demand curve hypothesis, such a decline is consistent with the long-run IPO underperformance that we observe following option listing. Finally, to analyze relative put-call demand, we use two measures: the adjusted implied volatility spread (IV spread) measure of Cremers and Weinbaum (2010), and the adjusted implied volatility skewness (IV skewness) measure of Xing, Zhang, and Zhao (2010). We find that both the IV spread and IV skewness indicate that put prices are significantly more expensive than call prices for several months after options listing. This suggests that informed traders use put contracts to place bearish bets on IPO stocks.

Our final set of results relates to the profitability of trading in newly listed options on IPO firms' equity. The two option trading strategies that we examine are writing call options or buying put options when they are newly listed on IPO stocks and holding them to expiration. While we find significant excess returns from buying long-maturity newly listed put options, average excess returns from writing call options are either insignificant or negative. Further, excess put option returns increase as their "moneyness" decreases: i.e., when they are more out-of-the money at the time of purchase. We also make use of a regression analysis to study the effect of various IPO characteristics on the excess returns from buying newly listed long-maturity (6 to 12 months) put options. We find that buying long-maturity puts on venture backed firms and those with higher initial return (IPO underpricing) yield significantly higher excess returns, while buying such puts on firms with longer time to option listing after IPO yield significantly lower excess returns.

The rest of this paper is organized as follows. Section 2 discusses how our paper relates to the existing literature. Section 3 describes our data, sample selection, and provides some summary statistics regarding option listing on IPO firm equity. Section 4 tests for the determinants of time-to-list options after IPOs. Section 5 describes our empirical tests and results on the effect of option listing on IPO stock returns. Section 6 presents our empirical analyses on the possible causes of IPO equity underperformance following option listing. Section 7 presents results on the profitability of trading strategies in the newly listed options on IPO firm equity. Section 8 concludes.

2 Relation to the existing literature and contribution

Our paper is related to two strands of literature. The first is the theoretical and empirical literature on the pricing and performance of IPOs in general, and on the long-run stock return underperformance of IPOs in particular. The first paper to document the long-run stock return underperformance of IPO firms was Ritter (1991). Subsequently, several papers examined various aspects of this underperformance; see Ritter and Welch (2002) for a review. However, the academic literature has been conflicted regarding the reasons underlying this long-run underperformance of IPO firm equity. For example, a number of authors have argued that the long-run underperformance of IPOs is due to market timing, where firms and investment bankers take advantage of the periodic deviation of stock prices from fundamental values by selling overpriced equity to investors (see, e.g., Ritter (1991) or Loughran and Ritter (1995)). Another theory of long-run IPO equity underperformance is based on the assumption of an IPO market characterized by heterogeneity in investor beliefs and short-sale constraints (see Miller (1977) for informal arguments and Morris (1996) for a formal theoretical model). The pricing of IPOs in the above setting reflects only the beliefs of the most optimistic investors (since pessimists' views are not reflected in the pricing of IPOs due to short-sale constraints), with the valuation of equity decreasing over time as the short-sale constraints on the IPO stock are relaxed (or alternatively, as the heterogeneity in equity investor beliefs is reduced over time with the arrival of additional "hard" information about the firm's post-IPO operating performance). While the focus of our paper is not on the long-run underperformance of IPOs in general, our paper provides support for the heterogeneous beliefs hypothesis by showing that the effective relaxation of short-sale constraint on an IPO firm through option listing is associated with long-run equity underperformance.

The second literature that our paper relates to is the empirical and theoretical literature on the determinants of option listing and its effects on the underlying stock return. The early empirical literature on the effects of option listing (see, e.g., Conrad (1989)) showed a positive abnormal return of about 2% around option listing. However, a more recent paper by Sorescu (2000) shows that positive listing returns are confined to the period from 1973 to 1980, and that negative abnormal listing returns become more frequent after 1980. Danielsen and Sorescu (2001) use an extension of the model of Jarrow (1980), and show that, since investors can establish short positions using options (and market makers hedge their exposure by shorting the underlying stock), the relaxation of short sale constraints due to option listing will drive down the price of the underlying stock. Cremers and Weinbaum (2010) and Ofek, Richardson, and Whitelaw (2004) study put-call parity violations to determine if informed investors can profit from mispricing in the stock market. Both of these studies find that such violations predict subsequent stock returns. Xing, Zhang, and Zhao (2010) study whether investors can profit from misvaluation in the underlying stock with more complex options strategies. They find that stocks exhibiting the steepest smirks in their traded options underperform stocks with the least pronounced volatility smirks. Finally, Mayhew and Mihov (2004) study the determinants of option listing for seasoned firms, and find that stocks with higher volume and volatility in the 250-day period prior to listing are most likely to be selected by options exchanges for listing.

Because our paper is the first to study the economic effects of option listing on IPOs, it contributes to the literature in several ways. First, we document the long-run equity underperformance of firms following option listing. While existing studies have documented short-run listing effects, to the best of our knowledge, there is no paper in the literature documenting a long-run listing effect, even in the context of seasoned firms. Second, while Mayhew and Mihov (2004) study the determinants of option listing for firms in general, we analyze the IPO characteristics associated with early option listing. Third, while several papers starting with Figlewski and Webb (1993) have documented that option introductions coincide with increased short-selling in the underlying stock, ours show that option listing relaxes the short-sale constraint on IPO firm equity. Fourth, by testing new hypotheses, we are able to develop additional insight into the causes underlying the stock return underperformance of firms following option listing. Finally, we are also the first to analyze the profitability of investment strategies in the options written on IPO firm equity.

3 Data and sample

3.1 Sample selection and descriptive

Information on option prices is collected from Ivy OptionMetrics, which provides end-of-day information on all US exchange-listed equities and market indices starting from January 1996. We obtain US initial public offerings data on ordinary common shares in the period from January 1996 to December 2008 from the Securities Data Company (SDC) New Issues database. At the time of writing this paper, OptionMetrics data ends in December 2011. We exclude IPOs issued after 2008 in order to allow sufficient time (at least 3 years) to follow their option listing history. We exclude closed-end funds, real estate investment trusts (REIT), dutch auction offerings, and issuances that are flagged with "shelf registration" according to SEC Rule 415. We obtain security prices and trading information from the Center for Research in Security Prices (CRSP). Issues that have a trading history on the CRSP earlier than that reported on the SDC database are deleted. Finally, we correct for remaining errors in the SDC database as suggested by Jay Ritter.[5] The final sample consists of 2,503 firms across 29 industries.

Ivey OptionMetrics report end-of-day option prices as well as standardized option prices calculated using raw information provided to them by Spryware, LLC. We find that information on newly listed options often first appears in the standardized option database, sometimes as early as two weeks, before it is reported in the end-of-day prices database. We hand checked a small sample of 25 new listings on the CBOE website and confirm that the standardized option database is the more accurate source. The date of option listing that we use in this paper therefore corresponds to the date that the firm's option information first appears in the Ivey OptionMetrics' standardized option database.

Panel A of Table 1 summarizes the distribution of initial public offerings by issuance year. In each year, we group firms according to when they have options listed relative to their IPO issuance date. As discussed in Ritter, Gao, and Zhu (2012), the number of firms that went public decline significantly after the tech bubble burst in 2000. Out of the 2,503 initial public offerings in our sample, we observe 1,448 option listings. We find that slightly more than half of the listings occur within the first year after the firms go public. On the other hand, if the IPO firms do not receive option listing within the first year, it is more likely to be listed very late (after 36 months), or never.

[5] http://bear.warrington.ufl.edu/ritter/SDC%20corrections.pdf

Panel B of Table 1 describes the sample. Conditional on observing an option listing, *days-to-list* is the number of days between IPO issuance and when option is introduced. The mean of days-to-list is 672 days while its median is only 324 suggesting that the distribution of timing of days-to-list is heavily left skewed. The shortest time to option listing is 9 days after the IPO issuance. In total, we observe 12 initial public offerings with days-to-list of 9 days. The longest days-to-list is 5670 days for Landec Corp (LNDC), a manufacturing firm that went public in February 1996.

Venture backed is an indicator variable equal to 1 if the IPO firm is financed by venture capital, and zero otherwise. Panel B of Table 1 shows that about half of the firms in our sample are financed by venture capital. *Underwriter rep* is the reputational score of the IPO's lead underwriter. We obtain underwriter reputation ranking from Jay's Ritter website (see also Loughran and Ritter (2004)) and map the lead underwriter of each offering to the reputational scale of 0 (lowest) to 9 (highest). The mean and median underwriter reputation score in our sample is 7.55 and 8, respectively. *First day ret*, a commonly used measure for short-run underpricing, is calculated as the log of return (%) on the offer price to the end of the first trading day. VA Linux system (LNUX) has the highest first day return in our sample of approximately 200% while Echapman Inc (ECMN) has the lowest first day return of -56%.

Proceeds amount is calculated as the log of the total proceeds amount (in $) from the offering including overallotment options, if exercised. *Total shares offered* is the number of shares issued at the offering in thousands of units. *Lock up agreement* is an indicator variable equal to 1 if the IPO has a share lockup agreement, and zero otherwise. Panel B of Table 1 shows that three-quarters of firms in our sample have shares lockup provision. If a firm has a share lockup agreement, we calculate *days to lockup expiration,* which is the number of days from IPO issuance to the first lockup expiration date. The median number of days to lockup expiration is 180.

3.2 Option listing date and IPO short and long-run performance

We evaluate the performance of initial public offerings using cumulative average adjusted returns (CAR). Following Ritter (1991), our benchmark is the return of a firm matched by industry and market capitalization to each IPO in our sample. The initial return period is defined as month 0, and the aftermarket period includes the subsequent 36 months. We define months as successive 21-trading-day periods relative to the IPO date. For instance, month 1 consists of event day 2-22, and month 2 consists of event day 23-43, and so on. Daily returns data and securities' information are downloaded from the CRSP database.

The matching firm-adjusted return for stock i in event month t is defined as

$$ar_{it} = r_i(t) - r_{im}(t),$$

where $r_i(t)$ is the logarithmic raw return of stock i in event month t, and $r_{im}(t)$ is the logarithmic raw return in event month t of a firm matched to stock i by industry and market value. Matching firms are selected from firms listed on the AMEX, NYSE, and NASDAQ stock exchanges. The matching procedure for each IPO stock i is as follows. We rank firms that have the same three-digit SIC code as firm i according to their market values calculated on the last trading day of the month prior to the IPO issuance date of stock i. Among them, the firm with the closest market value to that of the IPO firm at issuance is selected. If a matching firm in the same three-digit SIC code was not available, we match using the two-digit SIC code. Out of initial 2,503 IPO firms, we are able to find a matching firm for 2,499 of them without replacement.

The average adjusted return on a portfolio of n stocks for event month t is the equally-weighted average of the matching firm-adjusted returns. That is, $AR_t = \frac{1}{n}\sum_{i=1}^{n} ar_{it}$. IPOs that are delisted on event month t are excluded in the calculation of AR_t. The cumulative adjusted return from event month 0 to s is then calculated as the sum of average adjusted returns:

$$CAR\,(0, s) = \sum_{t=0}^{s} AR_t.$$

Figure 1 plots monthly CARs for six different IPO groups categorized by when they have options listed. For instance, the top-left and top-right panels plot CAR for the IPO stocks that have options listed in 0-6, and 25-36 months after the IPO issuance date, respectively. The darkened solid line in each panel indicates when options are listed relative to the IPO issuance date. Figure 1 shows that option listings usually coincide with the period when IPO stocks are performing well. Particularly, these strong performers with options listed relatiely early, i.e., within 24 months after their IPO, appear to persistently underperform after option listing. We do not observe a monotonic relationship between an IPO's first day return and time-to-list. The largest first day return belongs to the IPO group that have options listed in 7-12 months after the issuance. In fact, these firms exhibit the strongest short-run aftermarket performance (i.e. 6-month CAR) until options are listed.Thus, we find that option listing occurs near when the IPO firm equity is peak performing.

We further evaluate the long-run performance of initial public offerings using the price-to-value (P/V) ratio following the method in Purnanandam and Swaminathan (2004). Figure 2 plots the median P/V ratios at the monthly interval for the six IPO groups categorized by when options are listed. The P/V ratio measures the intrinsic value of the IPO equity against a comparable firm based on similar industry, sales, and EBIDTA/sales ratio. To calculate the P/V ratio, we require the IPO firms to have non-missing sales and EBITDA values in COMPUSTAT for the fiscal year prior to the initial public offering. Each IPO firm is matched to a non-IPO firm from the same Fama-French 48 industry group that has the closest prior fiscal year sales and prior fiscal year EBITDA/sales ratio to the IPO firm. Purnanandam and Swaminathan (2004) and Chemmanur and Krishnan (2012) describe details of the matching procedure. The resulting sample consists of 1,975 IPO firms. The P/V ratio for IPO firm i at time t is calculated at the monthly starting from the offering date (month 0) as:

$$PV_{it} = \left(\frac{P_i(t) \times Shrout_i(t)}{Prior\ year\ Sales_i}\right)\left(\frac{Prior\ year\ Sales_{iComp}}{P_{iComp}(t) \times Shrout_{iComp}(t)}\right).$$

$P_i(t)$ and $P_{iComp}(t)$ are equity values of the IPO firm i and its comparable firm at event month t, respectively. $Shrout_i(t)$ and $Shrout_{iComp}(t)$ are the number of shares outstanding for the IPO firm i and its comparable firm in event month t, respectively. $Prior\ year\ Sales_i$ and $Prior\ year\ Sales_{iComp}$ are the total sales of the IPO firm and its comprable firm in the fiscal year before the IPO firm i went public. The calculated P/V ratios are highly non-normal, and therefore, similar to Purnanandam and Swaminathan (2004), we plot their median values across firms for each event month. The median P/V ratio across firms at the offer date is 1.94 which isconsistent with Chemmanur and Krishnan (2012) who study IPO valuation over a simiar time period. This suggests that the intrinsic offering values of IPO firms in our sample are, on average, twice relative to their comparables.

Figure 2 shows the options are often introduced on the IPO firm equity when its P/V ratio is at the higest. However, subsequent to option listing, the intrinsic values of strong-performing IPO firms decline steadily over the long run. IPO firms that have high valuation at the offer date tend to recieve option listing within a year. The highest P/V ratio at the offer date correspond to IPO firms that have options listed in 7-12 months after the IPO issuance date. These findings are highly consistent with the long-run CAR results shown in Figure 1. Interestingly, we find that for IPO firms that never have options listed, their median P/V ratios steadily decline from approximately 1.97 at the offer date to about 1 within two years after these firms went public.

Overall, our findings in Figures 1 and 2 suggest that the timing of options listing significantly and negatively affects the long-run performance of the IPO firm equity.

4 Determinants of time-to-list

This section examines the factors influencing when IPOs are selected for option listings relative to their issuance date. Mayhew and Mihov (2004) study the determinants affecting exchange's decision of which options to list by estimating a logit model. They find that the stock's volume, volatility and size are positively related to the increase in the firm's likelihood of having options listed. While Mayhew and Mihov (2004) examine which seasoned stocks are selected for option listing, our focus is on when options are listed after the firm goes public. In keeping with this objective, our dependent variable is the *time-to-list* in unit months. Time-to-list is calculated by dividing the number of days between the IPO issuance date and the option introduction date by 30 and rounded to the nearest integer.

We estimate the extended Cox model commonly used for survival analysis in epidemiological studies (e.g. Platt, Joseph, Ananth, Grondines, Abrahamowicz, and Kramer (2004)). Let t be the current time period, and $T \geq t$ be the time when option is listed, the conditional probability that option is listed on the IPO firm between time t and $t+y$, $\mathbb{P}(t \leq T < t+y|T \geq t)$, is related to the following hazard-rate function:

$$h(t) = \lim_{y \to 0} \frac{\mathbb{P}(t \leq T < t+y|T \geq t)}{y}.$$

In the extended Cox model, the hazard function is represented by

$$h(t, \mathbf{x}, \mathbf{z}(t)) = h_0(t) \exp\left(\sum_{i=1}^{p_1} \beta_i x_i + \sum_{j=1}^{p_2} \delta_j z_j(t)\right), \tag{1}$$

where $\mathbf{x} = (x_1, x_2, \ldots, x_{p_1})'$ is a time-independent vector of IPO variables, and $\mathbf{z}(t) = (z_1(t), z_2(t), \ldots, z_{p_2}(t))'$ is a time-dependent vector of covariates affecting the hazard rate of having options listed. When $\delta_j = 0$ for all j's, the extended Cox model is known as the Cox proportional hazard (Cox PH) model, where $h_0(t)$ is the baseline hazard function. In both the extended Cox model and the Cox PH model, the baseline function is semi–parametric and hence we do not need to define the functional form for $h_0(t)$. Nevertheless, the extended Cox model is more flexible than the Cox PH model in that it allows for time-varying covariates to affect the hazard rate.

We estimate the hazard rate of time-to-list on various IPO characteristics and time-varying covariates. We truncate time-to-list beyond 36 months to focus on exchanges' listing decision in the IPO aftermarket.[6] The time-varying covariates include volume, volatility, and size which to varying degrees have been shown in Mayhew and Mihov (2004) to influence exchanges' option-listing decision. We use one-month lagged covariates to avoid biases induced by contemporaneous relationship between option listing and the time-varying covariates. These lagged covariates are calculated at the monthly frequency to be consistent with the time-to-list variables which is expressed in unit months. *Market cap* is the log of market capitalization observed at the end of the previous month and captures the firm size. *Shares*

[6] Efron (1977) show the Cox model is well suited for handling censored data. For robustness, we extend the truncation up to 60 months and obtain the same conclusions.

turnover represents the stock's trading volume and is calculated as the average of daily shares turnover (%) over the previous month. *Volatility* is calculated as the average daily standard deviation (%) of raw returns over the previous month. Besides size, volume, and volatility, we include *Cumulative ret* to capture the stock's momentum; it is calculated as the cumulative daily raw returns over the previous month.

We include IPO characteristics described in Panel B of Table 1. *Venture backed* is an indicator variable equal to 1 if the IPO firm is financed by a venture capitalist; *Time to lockup expiration* is the number of months from IPO issuance date to the first shares lockup expiration; *High underwriter rep* is a binary variable equal to 1 if the firm's lead underwriter is in the highest prestige category (rank=9) according to the ranking published on Jay Ritter's website; and *First day ret* is the logarithmic return (%) of the offered price over the end-of-first-trading-day price. We control for the IPO size in all the estimations by including *Proceeds amount* calculated as the log of total proceeds amount (in \$) from the offering including overallotment options, if exercised.

Table 3 reports the results from estimating the extended Cox model on time-to-list. Models (1)-(4) examine the impact of each IPO characteristic on time-to-list, while model (5) looks at the relative importance of IPO characteristics in explaining time-to-list. We include time-varying covariates in all the regressions. All models are estimated with the industry and year fixed effects. Table 3 shows that the coefficients on all time-varying covariates are positive and statistically significant. Consistent with Mayhew and Mihov (2004), we find that increases in the firm's market cap, shares turnover, and volatility positively increase the firm's likelihood of having option listed. The positive coefficient on cumulative returns supports the results in Figure 1 that IPOs experiencing positive return momentum are more likely to be selected for option listing.

In all models, the coefficient on *Proceeds Amount* is positive and significant at the 1% indicating that larger IPO firms are more likely to be listed early in the aftermarket. IPOs with large offering proceeds are usually associated with large institutional buyers and high media coverage (Liu, Sherman, and Zhang (2011)), resulting in active aftermarket trading and a substantial level of volatility. The finding that IPO proceeds positively correlate with shorter time-to-list supports our previous results that firm's size, trading volume, and volatility are important determinants of option listing.

Besides *Proceeds Amount*, the other variable that loads significantly positively on the likelihood of time-to-list is *Venture backed.* IPOs financed by venture capitalists therefore receive option listing earlier than non-venture backed IPOs. Table 3 shows the coefficients in models (1) and (5) on *Venture backed* are highly significant and economically large. A more intuitive way to interpret the coefficients in Table 3 is through the hazard ratio. According to equation (1), the hazard ratio of factor x_i is calculated as $\exp(\beta_i)$ and defined as the change in hazard rates conditional on one unit increase in x_i. Following this interpretation, the hazard ratios for *Venture backed* in model (1) is $\exp(0.506) = 1.66$ suggesting that the time-to-list for venture backed IPO firms is, on average, $\frac{1}{1.66} = 0.60$ times shorter than non-venture backed IPOs. Furthermore, we find the coefficient on *Venture backed* remains strongly significant and relatively stable once the other IPO characteristics are included in model (5). This suggests a robust finding that venture capital backed IPOs are likely to have option listing relatively early once they go public.

The coefficient on *High underwriter rep* is negative and statistically significant suggesting firms that undergo initial public offering with a prestigious underwriter receive option listing later than those that do not use a prestigious underwriter. Inferring the hazard ratio from the estimates in model (3), we find firms using underwriter with a high reputational ranking (=9) have options listed later by a factor of $1/\exp(-0.161) = 1.17$. Carter and Manaster (1990) show that IPOs that use prestigious underwriters are associated with lower risk offerings. They argue that with less information asymmetry, there is less incentive to acquire information and fewer informed trading. Following their argument, we conjecture the relatively lower risk offerings and informed trading level associated with firms using prestigious underwriter result in lower hedging demand and hence the lesser demand for option listing.

We find that IPO performance on the first trading day is weakly related to time-to-list. The coefficient on *First day ret* in model (4) is negative and significant at the 10 percent level, although with a relatively small magnitude. Translating this into the hazard ratio shows that one percent increase in the first day return at offering increases the time-to-list by only a factor of $1/\exp(-0.002) = 1.002$. We offer a potential explanation why high IPO's first day return may delay option listing from the point of view of the market maker. IPOs that experience extremely high first day return are considered severely underpriced and are very volatile. The market making cost associated with extremely underpriced IPO stocks is therefore expected to be very high which may consequently delay the listing. All in all, given the small economic magnitude on *First day ret*, we conclude that IPO's performance on the first trading day has little impact on when options are listed.

We do not find that the time from IPO issuance date to the first shares lockup expiration date affects when options are introduced. We also experiment with other shares lockup variables instead of *Time to lockup expiration* and arrive at the same conclusion. For instance, we use a binary variable equal to 1 if the firm has shares lockup agreement and zero otherwise. We also test whether the percent of shares held under the lockup agreement matter for option listing. Overall, we fail to find that time-to-list are associated with the aforementioned characteristics of shares lockup agreement. Finally, as a robustness check, we repeated our estimations on time-to-list using a logit model instead of the extended Cox model and obtain similar conclusions.

5 Consequences of option listing

5.1 Main results

5.1.1 Abnormal returns around option listing

In this section, we study the performance of IPO stocks around their option listing dates. We center our event windows around the date of option introduction. We compute abnormal returns at weekly intervals. Each event week return is calculated using successive five trading days. Return on event week 0 is the sum of log daily raw returns over the [-2,+2] trading days relative to the option introduction date. Returns on event weeks -1 and 1 are the sum of log daily raw returns over event days [-7,-3] and [+3,+7], respectively. Similarly, returns on event weeks -2 and 2 are the sum of log daily raw returns over event days [-12,-8] and [+8,+12], respectively, etc. We calculate weekly returns over a long horizon starting on event week -12 and ending on event week +52 relative to the option listing date.

We measure abnormal returns of stock i on event week t using benchmark-adjusted returns, ar_{it}, calculated as the raw return of stock i minus the return of a benchmark for the same event week t. The benchmarks used are: (1) the CRSP value-weighted NASDAQ index, (2) an index of the smallest decile of the New York Stock Exchange, and (3) a listed firm matched by industry and size. The data for the two benchmark indices are downloaded from CRSP and the matched firm is selected using the method discussed in Section 3.2. The calculation of weekly average adjusted return and weekly cumulative adjusted return are identical to those in Section 3.2, with the exception of using weekly benchmark-adjusted returns instead of monthly matching firm-adjusted returns.

Figure 3 plots weekly cumulative adjusted returns (CAR) around option listing for the four IPO groups categorized by when options are listed relative to their IPO issuance date. Weekly cumulative adjusted returns are plotted starting from event week -12. We plot CAR series calculated from the three benchmark-adjusted returns as well as the raw returns. The bottom-right panel in Figure 3 plots CAR for the full sample consisting of IPO stocks that have options listed within 3 years after going public. The other panels in Figure 3 plot CAR series for different subsamples. We include a solid vertical line in each panel to visually indicate when option listing occurs. Overall, these plots convincingly show that IPO firms, on average, experience their peak benchmark-adjusted performance in event week 0, exactly when options are listed. More importantly, Figure 3 shows that IPO stocks continue to underperform once options have been introduced. The results are consistently robust across the four IPO groups although the long-run underperformance are stronger for firms that have options listed within 1 year after going public.

Table 3 reports CAR calculated using matching firm-adjusted returns around option-listing week. We compute CAR for four non-overlapping event windows. The first event window is designed to capture abnormal returns before option is introduced and covers event weeks -12 to -2. We exclude week -1 from the first event window because from hand-searching 25 randomly selected listing announcements on the CBOE website, we find that the announcement date is usually one week prior to the option introduction date. The second event window covers event weeks -1 to 1 and is designed to capture any price effects associated with the announcement of option listing as well as option introduction. The third event window measures the short-run price impact after options have been introduced; it covers event weeks 2 to 12. Finally, the fourth event window measures the price impact of option listing over event weeks 13 to 52.

Panel A of Table 3 reports CARs around option listing weeks for all observations. In Panels B and C, we report results grouped by when options are listed after the IPO issuance (Panel B) and the year that the firms go public (Panel C). Clearly, Table 3 shows that IPO stock returns strongly outperform in the weeks prior to option listing. The results for the full sample (Panel A) show that IPOs earn a significant 12.29% CAR in the event window [-12,-2]. The strong positive performance quickly disappears after options are listed resulting in a statistically insignificant CAR value of -2.05% in the event window [+2,+12]. Looking at a longer horizon, the CAR value falls to -17.63% in the event window [+13,+52] and is statistically significant at the one percent level. We find similar abnormal return patterns when looking at Panels B, suggesting the impact of option listing applies to all IPO groups. However, the biggest underperformance after option listing belongs to the IPO group that have option introduced in the first six months.

We next test the hypothesis that IPO stock returns decrease after option listing by calculating the difference in CAR between post-event window [+2,+12] and pre-event window [-12,-2]. The last row of each panel in Table 3 reports the result. Overall, we find that CAR significantly decrease after options have been introduced on the firm equity. The difference is largest for IPOs that have options listed in 0-6 months and 13-36 months, respectively. Interestingly, Table 3 shows the difference in CAR is smallest for IPOs that have options listed in 7-12 months (Panel B). We conjecture that the result is partly caused by the expiration of IPO shares lockup, which is usually 180 days after the issuance date and is associated with a significant negative abnormal return (see Field and Hanka (2001)).[7] Since options are listed after the lockup expiration date, the CAR value in the pre-option-listing period is affected by the negative price impact of shares lockup expiration. We confirm this conjecture by looking

[7] Seventy six percent of IPOs in our sample feature share lockup agreements. Field and Hanka (2001) report a three-day abnormal return of -1.5% and a permanent 40% increase in trading volume around the unlock days. They further show that insiders sell their shares aggressively after lockups expiration. Similarly, Bradley, Jordan, Ha-Chin, and Roten (2001) find a significant negative price impact around lockup expirations and that the loss appear to be concentrated among firms financed by venture capital. Figure A1 in the Appendix shows that abnormal IPO performance around option listing that we observe is not an artifact of shares lockup expiration.

at the CAR value in the [-12,-2] window in Panel B for the 7-12 months IPO group. The CAR for this group in the [-12,-2] window is 7.48% which is the smallest relative to the other IPO groups. As a result, the difference in CAR between the pre- and post-option listing periods for this IPO group also has the smallest magnitude. We further examine the robustness of our results to shares lockup expiration. Figure A1 in the Appendix reports CAR around option listing date using only listing events that are not contaminated by the IPO lockup expiration. Overall, we find that our main finding on the impact of option listing is unaffected by shares lockup expiration.

Existing studies find that the IPO market follows a cycle with dramatic change in the number of issuance and volume per year.[8] In Panel C of Table 3, we present CARs calculated for four subperiods grouped by the IPO issuance year. Overall, we find that our main conclusion remains robust to the year of firms' IPOs.

5.1.2 Regression analysis

In this subsection, we use panel regressions to examine IPO performance around option listing. Table 4 reports the results. The dependent variable is weekly matching firm-adjusted return, ar_{it}, calculated relative to the option listing dates, which is event week 0. We run panel regressions on two intervals of weekly adjusted returns. Panel A reports regression results for weekly adjusted returns for a balanced panel starting in event week -12 and ending in event week 12. Panel B reports regression results for the long-run window from event weeks -12 to 52.

We test the impact of option listing on IPO stock performance by regressing weekly adjusted returns on *Optionstat,* an indicator variable equal to 1 in event week 0 or greater, and zero otherwise. We estimate the following regression model for each firm i:

$$\begin{aligned} ar_{it} &= \alpha + \beta\, Optionstat + \textstyle\sum_{i=1}^{4} \gamma_i \left(Optionstat \times IPO\ variable_i\right) + \delta_1 IPO\ variable_i + \\ &\quad + \delta_2 Lockup\ expiration(t) + \delta_3\, Weekly\ Controls_{it-1} + \varepsilon_{it}, \end{aligned} \tag{2}$$

where *IPO variable*$_i$ is the IPO variable under examination. Models (1) and (3) report the baseline regression results where $\gamma = 0$ in equation (2). In models (2) and (4), we estimate the coefficient on *Optionstat* as well as its interaction with the other IPO variables. The objective of running regression models (2) and (4) is to examine whether the underperformance that we observe after option listing is associated with certain IPO characteristics. Most of the IPO variables that we use in the regressions model are identical to those in Table 3. *Time-to-list* is the number of months from IPO issuance date to the date of option introduction.

[8] See for examples, Ibbotson and Jaffe (1975), Ritter (1984) and Loughran and Ritter (1995).

In addition to the year fixed effects, we include lagged weekly variables, *Weekly Controls*$_{it-1}$, to control for time-varying changes in the firm's momentum, size, volume, and volatility. We control for momentum using the matching firm-adjusted return, ar_{it-1}, calculated in event week $t-1$. We control for size and volume using the average daily log market capitalization and daily shares turnover of firm i in event week $t-1$, respectively. We control for changes in volatility using the standard deviation of stock i's daily raw returns calculated in event week $t-1$. Finally, we include the dummy variable *Lockup expiration (t)* to remove the well-documented price impact of shares lockup expiration. This variable takes the value of 1 if the firm has share lockup expiration in event week t, and zero otherwise.

The regression model in equation (2) is useful for interpreting the magnitude of IPO performance around option listing. The intercept α provides an estimate of average weekly adjusted returns prior to option listing. When γ_i's= 0, equation (2) provides estimates for the baseline regression, where β represents an average change in weekly adjusted returns after option introduction. Panel A in Table 4 shows the coefficient estimates on *Optionstat* are significant and negative for all models reinforcing our main findings that IPO stocks underperform once options have been introduced. Looking at the baseline estimates in model (1), the coefficient on *Optionstat* is -1.234% which is statistically significant with the t-stat of 6.35. This finding suggests that IPO stocks' weekly adjusted returns decrease by 123 basis points, on average, once options have been listed. Similar to Field and Hanka (2001), we find an economically significant price impact of IPO shares lockup expiration. The coefficient on *Lockup expiration (t)* is -3.16% suggesting that the IPO equity price drops by about 3 percent in the week that its shares lockup agreement expires.

Next we look at the results from interacting *Optionstat* with IPO characteristics. Model (2) shows the estimate on *Optionstat*$\times$*Venture backed* is -1.20% suggesting the impact of option listing is more severe for venture capital backed firms. This estimate implies that weekly adjusted returns of venture backed IPOs decrease further than non-venture backed IPOs by 120 basis points. Another IPO characteristic that strongly impacts post-option listing performance is the lead underwriter's reputation. Model (2) shows a positively significant coefficient on *Optionstat*$\times$*High underwriter rep* of 0.406%. In other words, weekly adjusted returns in the post-option listing period for firms using a prestigious underwriter is 46 basis points higher than those using a non-prestigious underwriter. Finally, we do not find that the first trading day return as well as the time-to-list options significantly impact post-option listing performance.

Panel B of Table 4 reports regression results over the [-12,+52] window. We find the coefficients on *Optionstat* in Panel B are strongly significant and negative indicating that our previous findings hold at a longer horizon. In fact, results from model (4) shows the impact of option listing on IPO underperformance is highly persistence and slightly stronger over the long run. The coefficient on *Optionstat* is -1.486% and is statistically significant at the one percent level (t-stat is -8.77). This translates to lower post-listing weekly adjusted returns of 149 basis points relative to when options have not yet been listed. Finally, the results regression model (4) confirm our previous findings on the significance of *Venture backed* and *High underwriter rep.*

Overall, we find very strong results that weekly returns of IPO firms are significantly lower once options have been listed on their equity. Further, the magnitude of option listing on IPO equity appears to be robustly determined by their venture-capital backed status, and the reputation of their underwriter.

5.2 Controlling for exchanges' listing decision

5.2.1 Propensity score matching method

Our results in Tables 3-4 support the conclusion that IPO equity returns decrease subsequent to option listing. However, Table 3 shows there is a selection bias because exchanges tend to list options on a stock following its period of high trading volume, volatility, market capitalization, and upward momentum. Further, certain IPO characteristics such as venture capital backing, and the underwriter's reputation affect when exchanges will list options after the firms go public. Therefore, the alternative hypothesis is that option exchanges may have superior stock picking ability such that they time their listing to coincide with when the IPO stocks are performing at their peak. To rule out this alternative hypothesis, we address the selection bias using a propensity score-matched control sample and show that our main finding is not due to exchanges' ability to time the market.

We construct a control group of eligible IPO firms based on their likelihood of having options listed. Our approach is similar to Mayhew and Mihov (2004) who uses the propensity-score matched control sample of seasoned firms to examine the impact of option listing on firm's equity volatility. However, instead of using seasoned firms, we require that firms in our control group are relatively recent IPO firms facilitating their comparisons with options listed-IPO firms in the treated group. The first step in constructing the control sample is to determine the universe of IPO firms that are eligible for option listing but do not yet have options listed. We use firms that went public in 1996-2008. ADRs, country funds, REIT, and closed-end funds are excluded since they are not considered in our IPO sample. For firms to be considered eligible for matching, they must not have become seasoned for more than 48 months. We also require that eligible control firms must have option listed within 48 months after their IPO issuance date. This requirement ensures that these control-group firms are, at some point, eligible for option listing.

In the next step, we estimate the following logit model to select eligible IPO firms that do not have options listed:

$$L\left(List_{i,t}\right) = \alpha_0 + \sum_{i=1}^{4} \alpha_i IPO\ variable_i + \sum_{j=1}^{4} \beta_j Monthly\ Controls_{jt-1}, \tag{3}$$

where $L\left(List_{i,t}\right)$ is the log-odds ratio that firm i will be selected for option listing during month t. The independent variables we use are IPO characteristics, $IPO\ variable_i$, and one-month lagged time-varying covariates, $Monthly\ Controls_{jt-1}$ obtained from Table 3. Details of the matching procedure as well as the matching results are described in the Appendix A.

After estimating the logit model in equation (3) on all eligible firm-month observations from January 1996 through December 2012, we use the nearest-neighbourhood caliper matching approach of Cochran and Rubin (1973). We exclude observations with propensity score below 5 percent and require that the propensity score of the control observation is within $\pm 5\%$ of that of the treated observation. Importantly, for each matched pair, we require that the treated observation is matched to an IPO firm that *will not* have options listed on their equity for at least 12 months. This last requirement is critical because our objective is to measure the difference in long-run returns (up to one year) between IPO firms that experience and do not experience options listing. Each firm-month observation is matched without replacement. However, due to the small number of IPO firms eligible for the control group, we allow each control firm to be matched with up to two treated-group firms. The final matching yields 541 matched pairs, where 541 unique treated-group firms are matched to 349 unique control-group firms. Table A1 in the Appendix reports logistic regression results for the samples before and after matching.[9]

5.2.2 Propensity score-matched sample results

Using the propensity score-matched sample, we perform a difference-in-difference regression analysis. We estimate a panel regression model similar to that in equation (2) for the matched sample and separately for the treated and control samples. Table 5 reports the results. The dependent variable is the weekly matching firm-adjusted returns. We report the results for three observation windows. The first is the pre-event window corresponding to the weeks [-12,-1] prior to when option is listed. The second and third observation windows correspond to weeks [-12,+12] and weeks [-12,+52] relative to the option listing week.

[9] In an earlier version of this paper, we followed the matching method identical to Mayhew and Mihov (2004) by running logistic regressions for each month and matching the treated and control firms on a rolling month-by-month basis. Such method relies on eligible *seasoned* firms in the control sample and hence is straightforward due to a large number of control-group firms available for matching. The results from the rolling month-by-month matching can be made available upon request. Overall, our results remain qualitatively the same when using either IPO firms or seasoned firms in the control group.

The general regression model that we estimate is

$$
\begin{aligned}
ar_{it} \;=\;& \alpha + \beta_1 \mathit{Treated} + \beta_2 \mathit{Optionstat} + \gamma\,(\mathit{Optionstat} \times \mathit{Treated}) + \cdots \qquad (4)\\
& + \delta_1 \mathit{IPO\ variables}_i + \delta_2 \mathit{Lockup\ expiration}(t) + \delta_3 \mathit{Weekly\ Controls}_{i,t-1} + \varepsilon_{it},
\end{aligned}
$$

where *Optionstat* is a binary variable (0 or 1) indicating whether options have been introduced. Option listing week is defined as event week 0 and therefore, *Optionstat* takes the value of 0 from weeks -12 through -1, and the value of 1 from weeks 0 through +52.[10] In the regression model in equation (4), we introduce a new variable, *Treated* which is an indicator variable equal to 1 if the firm is in the treated group and zero otherwise. The coefficient on the interaction term $\mathit{Optionstat} \times \mathit{Treated}$ is the DID estimator, which is our variable of interest. The DID estimator measures the difference in adjusted returns between before and after options have been listed for the treated group relative to that of the control group. We include IPO characteristics used in the propensity score matching in the above regression. We also include lagged weekly variables, $\mathit{Weekly\ Controls}_{it-1}$, in (4) to control for time-varying changes in the firm's momentum, size, volume, and volatility.

Regression model (1) reports regression results for the matched sample prior to option listing. For this sample, *Optionstat* takes the value of 0 for all observations and hence the DID estimator is not available. In this case, the *Treated* variable directly measures the difference in CARs between the treated ($\mathit{Treated} = 1$) and control ($\mathit{Treated} = 0$) groups in the pre-event weeks. Model (1) shows the coefficient estimate on *Treated* is not statistically significant suggesting that the treated and control groups are reasonably well matched along the cumulative returns dimension prior to the event week.

Regression models (2) and (3) report results based on the regression model equation (4) estimated using the matched sample over the observation windows [-12,+12] and [-12,+52], respectively. We find the coefficients on $\mathit{Optionstat} \times \mathit{Treated}$ are negative and highly significant. Importantly, their values are economically large. For instance, the DID estimator regression specification (2), which is estimated over the event weeks [-12,+12] is -0.750 and significant at the five percent level. This magnitude can be interpreted as follows. Controlling for the price effect associated with exchanges' listing decision, weekly adjusted returns of IPO stocks, on average, decrease by 75 basis points after options have been introduced. Further, the coefficient on *Treated* is statistically insignificant, confirming that our matched samples consists of stocks that perform similarly before option is introduced.

Another variable of interest in regression models (2) and (3) is *Optionstat.* The coefficient

[10]For the control-group firms, option listing month corresponds to when the non-option listed IPO firms have been matched to option listed IPO firms (Treated group). The control-group firms are assigned option listing dates identical to their matched pairs from the treated group.

on *Optionstat* is negative but insignificant in regression specification (2). It is, however, negative and statistically significant in regression specification (3) which measures post-listing performance over a longer horizon. We therefore find some evidence supporting exchanges' ability to time their option listing on stocks before they underperform.

6 Explaining IPO underperformance post option listing

We evaluate three hypotheses related to the underperformance of IPO stocks after options have been listed. First, we test whether option listing alleviates short sale constraints for IPO stocks. Second, we test whether insiders sell their shares aggressively for months after options have been listed, which could explain why IPO stocks continually underperform for months after options have been listed. Our third hypothesis tests whether put prices are abnormally more expensive than call prices when option are newly listed on the IPO stocks. If informed traders use the options market to trade on negative future IPO performance, put options should be in relatively higher demand than calls. .

6.1 Short interest

One of the long standing arguments for short-run IPO underpricing is the difficulty with selling IPO stocks immediately after the firms go public. Under short sale constraints, Miller (1977) argues that prices will reflect a more optimistic valuation of the security because pessimistic investors are kept out of the market by high short-sale costs. D'Avolio (2002) provides institutional details of the short-sale market and shows that an important cost related to short selling is the loan fees. An investor who wishes to short sell a stock must find an existing owner who is willing to lend her the shares. Having negotiated a loan cost, she may short sell the borrowed share to any willing buyer. Relying on a proprietary database, D'Avolio (2002) reports the value-weighted mean loan fee for general collateral stocks of 17 bps per annum. However, the fee for stocks that are "special" is significantly much higher (4.69%). Geczy, Musto, and Reed (2002) argue that IPO stocks are "special" for the following reasons: the restricted supply of shares in the immediate aftermarket as share allocations are incomplete, the lockup agreement which restricts insiders from selling and lending their shares, the prohibition on banks in the IPO syndicate from lending shares in the first 30 days, and the relatively small public float in most IPOs.

More recently, Edwards and Hanley (2010) question the assumption that short-run IPO underpricing is due to short sale constraints. Using short selling transaction data in 2005-2006, they show that short selling is an integral part of the IPO aftermarket, therefore concluding that other factors must be at play. Even though short selling in the aftermarket is not as inactive as previous studies suggested, its cost may still be too high such that several investors who hold negative views on the IPO stocks are deterred from participating in their price discovery. In fact, Edwards and Hanley (2010) show that short sellers, on average, do not appear to earn abnormal profits suggesting insignificant rewards from taking such short positions. This is in contrast to when options are introduced as we will later show in Section 7 that investors who take synthetic short positions by buying long-maturity put options, on average, earn significant positive excess returns.

An important cost advantage of synthetic shorting by purchasing puts is the indirect avoidance of the loan fee, which can be substantially high when stocks are "special". Essentially, investors who buy put contracts transfer the burden of locating the borrowed shares as well as the loan fees to the market makers who can hedge their position by short selling the underlying security at a lower cost.[11] Danielsen and Sorescu (2001) provide several reasons why option market makers can avoid most of the costs faced by other short sellers. Among those reasons is the market makers' exemption from locating shares to borrow before shorting.[12] In fact, Evans, Geczy, Musto, and Reed (2009) provide evidence showing that in most hard-to-borrow stocks, the market maker chooses not to borrow and instead fails to deliver stock to their buyers. Following the above argument, the level of short interest should increase significantly when options are introduced, reflecting the participation of pessimistic investors in IPO price discovery.

We obtain data for adjusted end-of-month short interest from Shortsqueeze and COMPUSTAT North America. Short interests information for newly public firms in COMPUSTAT are often missing. We supplement short interests obtained from COMPUSTAT with commercially purchased data from Shortsqueeze. When end-of-month values are not available, we use the reported mid-month adjusted short interest. COMPUSTAT's short interest coverage begins in 2003 (Shortsqueeze coverage begins in 2004), the sample consists of IPOs in 2003-2008. We exclude option listings that overlap with IPO unlock days within +/- 180 days, because Ofek and Richardson (2003) show that short interests also increase significantly when IPO lockups expire. The columns under *Rel change* in Table 7 report relative changes in short interest ratio from event months -3 to 6, where event month 0 is when option is introduced. Short interest ratio is calculated as the total number of adjusted short interest divided by the number of shares outstanding. Relative change in short interest ratio is then calculated as a change in short interest ratio relative to event month -4 in percentage terms. We exclude firms that do

[11] The option delta of a put contract is negative. Therefore, hedging a put contract involves short selling the underlying security.

[12] Regulation SHO, passed in July 2004, requires the broker facilitating a short sale must locate the stock prior to the trade. However, those engaged in bona-fide hedging of market making activity are exempt from this rule.

not have available short interest data on event month -4.[13] Table 7 reports across-firm average values of *Rel change* the treated and the control group. Firms in the treated group are IPOs that experience option listing while control-group firms are IPO firms that do not yet have options listed. The matching is done based on the propensity that the firms will have option listing, as discussed in Section 5.2.1. The reported average values are winsorized at the 1st and 99th percentile in order to mitigate outliers; this practice does not qualitatively affect our conclusions.

The results in Table 7 and the top panel of Figure 3 clearly show that *Rel change* in the treated group permanently increase by about 300% after options have been listed. More importantly, the largest increase in *Rel change* appears to be on the option listing month. On the other hand, firms in the control group experience about 135% increase in relative short interest ratio on event month 0 but gradually increase to catch up to the treated group after about 6 months. The difference in *Rel change* between the treated and control groups between event months 0 and 4 is also significant as suggested by the two nonparametric tests. We use nonparametric tests instead of the conventional t-test due to the small sample size as well as the non-normal distributions of *Rel change*. Overall, our results strongly show that option listing is associated with an economically large increase in short interest. Danielsen and Sorescu (2001) also find short interest ratios increase significantly around option listing. Our results further support their finding by showing that short interest ratios significantly increase around option listing after controlling for the endogeneity associated with exchanges' listing decisions.

Finally, in order to link the above documented increase in short interests to options' market making activities, we examine the number of failed-to-deliver short positions around option listing week. When an option market maker writes a put option, she can hedge her position by short selling the underlying security. However, Evans, Geczy, Musto, and Reed (2009) finds that in most hard-to-borrow stocks, i.e. IPOs, the market maker chooses not to borrow and instead fails to deliver stock to their buyers. We hand-collect daily failed-to-deliver quantities from the U.S. Securities and Exchange Commission's (SEC) website. The SEC website reports failed-to-deliver short positions on the daily basis starting in mid-2004. We merge failed-to-deliver data with our propensity-score matched IPO samples and find that data available on 46 treated-group firms and 69 control-group firms. We then calculate relative change in failed-to-deliver quantities around option listing weeks (relative to event week -1) and find that it increases by 73.86% for the treated-group IPO firms, and 19.09% for the control-group firms. Interestingly, the quantity of failed-to-deliver short positions continually increased

[13] This requirement is likely to understate the relative changes in short interest ratio because our sample represents IPOs that have option listed relatively late (i.e., at least after 6 months and on average after one year relative to the IPO issuance date).

after options have been introduced. After two weeks since options have been introduced, we find that failed-to-deliver quantities of the treated-group firms have increased by about 186% relative to the week prior to option listing, while the same value for the control-group firms is only 55%. Importantly, the difference in these values are statistically significant at the 5 percent confidence level.

Overall, the results in this section allow us to conclude that short interests increase sharply after options have been introduced on the IPO firms' equity, and that such increase is related to market making activities of options' dealers.

6.2 Insider selling

Evidence for downward-sloping demand curves for equity has been shown by Shleifer (1986), Bagwell (1992), and Kaul, Mehrotra, and Morck (2000). So, when insiders sell their shares, the share price will fall. Unlike the price pressure hypothesis, the effect of price changes due to the downward demand curve hypothesis is permanent. Thus, the long-run underperformance of IPO stocks after option listing may be associated with changes in insider ownership. We test this hypothesis here.

We obtain insider filing data from Thomson Reuter's Insider Filing Data Feed (IFDF). For each month, we compute the total direct and indirect holding of conventional stock (Table 1 in IFDF) held by the company's insiders with relationship hierarchy level between 1 to 3. This hierarchy cutoff includes 10% block holders, executive officers, board members, and members of the various operational committees. Indirect holdings refer to shares held by insiders' immediate family members while direct holdings are shares held by the insiders.

Table 8 reports changes in insider holdings from event months -3 to 6 relative to event month -4. The bottom panel of Figure 3 also illustrates our results. The procedure for calculating the relative change in insider holdings is similar to that for the relative change in short interest ratios, except the variable of interest is the total insiders' shares held. Firms in the treated group consist of IPOs in 1996-2008 that have options listed within 36 months after the issuance date. Firms in the treated group are matched to IPO firms that do not yet have options listed (Control group) based on their propensity of having options listed. Field and Hanka (2001) find that insiders, especially those that finance venture backed IPOs, sell their shares aggressively after the lockups expire. We therefore exclude listings that overlap with unlock days within one calender year.

Table 8 shows the relative changes in insider holdings between the treated and control groups do not significantly differ in event months -3 to -1. However, the difference between the treated and control group becomes remarkably clear after options have been listed; this is also evident in the bottom panel of Figure 3. Firms in the treated group continually experience a decrease in insider holding for months after options have been listed. For instance, we find that the total shares held by insiders decrease by 20.8% after options have been listed for 6 months relative to event month -4. The number of shares sold to public by insiders after option listing is economically large. According to the downward sloping demand curve hypothesis, such unloading of shares in the months after option listing is consistent with the IPO underperformance that we observe.

6.3 Relative put-call demand

According to the asymmetric information model of Easley, O'Hara, and Srinivas (1998), informed traders may choose to trade using their private signals in the option market resulting in a predictable relationship between option demand and subsequent underlying stock returns. In their model, buying a put or selling a call increases the prices of put options relative to call options conveying negative information about the underlying future stock prices. If informed trades use the option market to trade on negative future performance of the IPO stocks, put prices should be expensive relative to call prices when options are introduced.

We use two measures of relative put-call pricing. The first is the adjusted implied volatility spread (IV spread) measure of Cremers and Weinbaum (2010) which represents the deviation to the put-call parity relation. The second measure that we use is the adjusted implied volatility skewness (IV skewness) of Xing, Zhang, and Zhao (2010) which represents the difference between implied volatilities calculated using at-the-money calls and implied volatilities calculated using out-of-the-money puts. Table 9 reports the results. We report the averages of daily adjusted IV spread and daily adjusted IV skewness calculated over each event month. We define each event month as a successive 21-trading-day periods relative to the option listing date. Month 0 is when option is listed and includes event day 0-20 relative to option listing. Month 1 includes event day 21-41 and month 2 includes event day 42-62 relative to option listing, etc.

An IV spread is the difference between implied volatilities of a call and a put options with identical strike price and maturity. For each stock i on day t, the IV spread is calculated as

$$IVspread_{i,t} = \sum_{j=1}^{N_t} w_{j,t} \left(IV_{j,t}^{call} - IV_{j,t}^{put} \right), \tag{5}$$

where $IV_{j,t}^{call}$ and $IV_{j,t}^{put}$ refers to the implied volatilities for the j^{th} pair of call and put option with identical strike price and maturity, N_t is the number of valid put-call pairs, and $w_{j,t}$ is the weight. Note that we drop the i subscript on the variables in the right-hand side of equation (5) for brevity. Following Cremers and Weinbaum (2010), we use the average open interest of each put-call pair as the weight $w_{j,t}$ in (5). Finally, the adjusted IV spread is calculated by subtracting $IVspread_{i,t}$ with $\overline{IVspread_t}$ calculated as an equally weighted average of $IVspread_{i,t}$ across all stocks on day t. Under this definition, adjusted IV spread represents stock i put-call mispricing relative to the market average. A negative value of adjusted IV spread would suggest that the put price is abnormally more expensive than the call price, while a positive value of adjusted IV would suggest the opposite.

The calculation of daily adjusted IV skewness is as follows. For each stock i on day t, we calculate an IV skewness as

$$IVskewness_{i,t} = IV_t^{ATMC} - IV_t^{OTMP}, \tag{6}$$

where $IV_{i,t}^{ATMC}$ is the implied volatility of an at-the-money call, and $IV_{i,t}^{OTMP}$ is the implied volatility of an out-of-the-money put.[14] We use a put contract with moneyness (F/K) closest to 0.95 and a call with moneyness closest to 1. Finally, the adjusted IV skewness is calculated by subtracting $IVskewness_{i,t}$ with $\overline{IVskewness_t}$ calculated as an equally weighted average of $IVskewness_{i,t}$ across all stocks on day t. Similar to the measure of adjusted IV spread, a negative value of adjusted IV skewness suggests that the put price is abnormally more expensive than the call price, while a positive value of adjusted IV suggests the opposite.

Table 9 shows that both adjusted IV spread and IV skewness are significantly negative for several months after an options have been listed. This finding shows that put prices are significantly more expensive than call prices when they are newly listed. Looking at the adjusted IV spread measure (Panel A) shows that the relative mispricing persists for 9 months after options have been listed. We find a similar result when looking at the adjusted IV skewness, although the relative mispricing disappears after 5 months. Overall, we find strong evidence to support the hypothesis that informed traders, and perhaps speculators, use put contracts to take synthetic short positions on IPO stocks for several months after options have been listed.

6.4 Explaining post-option-listing stock returns

We now study how the variables that we examined in the previous subsections, i.e. change in short interest ratios; change in the level of insider shares holdings; and relative put to call option prices, are associated with subsequent IPO underperformance once options are listed.

First, we examine the factors that explain the magnitude of IPO stock underperformance *when* options are listed. We address this question by estimating the following panel regression

[14] We define the IVskewness variable in equation (6) slightly differently from Xing, Zhang, and Zhao (2010). They calculate IV skewness by subtracting the IV of out-of-the-money put with the IV of at-the-money call. Therefore, our IV skewness measure will have an opposite sign to theirs.

of monthly adjusted returns from event months -3 to +12 relative to when options are listed:

$$\begin{aligned} ar_{it} &= \alpha + \beta\, Optionstat + \gamma\, (Optionstat \times Q_i) + \ \delta_1\ Q_i \ + \ \ldots \\ & \ldots\ + \delta_2 Fixed\ effects_i + \delta_3 Monthly\ Controls_{it-1} + \varepsilon_{it}, \end{aligned} \tag{7}$$

where Q_i is the variable under examination. Table 10 reports the results. Monthly adjusted return, $ar_{i,t}$, for firm i in month t is calculated using the return of an industry-matched firm as a benchmark. *Optionstat* is the monthly indicator variable equal to one when options have been listed, i.e. $t \geq 0$, and zero otherwise. We include, in the regression model, standard IPO variables identical to those reported in Table 3 to control for any IPO fixed effects. For the purpose of our analysis, we are interested in the coefficient estimate on $Optionstat \times Q_i$ which measures the magnitude of IPO underperformance in relation to Q_i when option is listed. We study the magnitude of post-option-listing underperformance in relation to four variables: (i) *IV spread*(1), defined as the average implied volatility spread in event month 1; (ii) *IV skewness*(1), defined as the average implied volatility skewness in event month (1); (iii) Δ*Short interest ratio*(1), defined as the relative change in short interest ratios in event month 1; and (iv) Δ*Insider holding*(1), defined as the relative change in total insider shares holdings in event month 1. The calculation of these variables are explained in the previous subsections.

The results in Table 10 show the coefficient on *Optionstat*$\times\Delta$*Short interest ratio*(1) is negative and statistically significant. In other words, we find stronger IPO underperformance for firms that experience larger increase in short interest ratios when options are listed supporting the hypothesis that option listing significantly relaxes the short sale constraint on the IPO firm equity. Table 10 also shows that the coefficients on *Optionstat*$\times$*IV spread*(1) and *Optionstat*$\times$*IV skewness*(1) are positive and significant at the 10 and 5 percent levels, respectively. These results suggest the post-option-listing underperformance is stronger for firms that have relatively higher put-to-call option prices; note that the more negative *IV spread* and *IV skewness* are, the more expensive is put price relative to call price. We therefore find evidence supporting the hypothesis that informed traders use put options to trade on negative future performance of the IPO stocks. We however do not find that the change in total shares held by insiders around option listing month predicts the magnitude of subsequent IPO underperformance. This finding is not surprising because IPO lockup periods often restrict insiders from selling their shares freely to the public. Therefore, the effect of insider selling may not be visible until the lock up period expires, which can be several months after options are listed.

We next examine the variables that are associated with the observed long-run underperformance in the months *after* options have been listed. Table 11 reports results from the following regression model over event months +2 to +24:

$$ar_{it} = \alpha + \beta Q_{i,t-1} + \delta_1 Fixed\ effects_i + \delta_2 Monthly\ Controls_{it-1} + \varepsilon_{it}, \tag{8}$$

where $Q_{i,t-1}$ is the monthly lagged variable under examination. The use of regression window [+2,12] allows us to focus on the long-run IPO performance after options have been listed. We regress monthly adjusted returns on one-month lagged variables representing the level of IV spread, IV skewness, short interest ratio, and percentage shares held by insiders. Table 11 shows that the coefficients on *IV spread*(t-1) and *IV skewness*(t-1) are positive and significant at the 10 and 5 percent levels, respectively, indicating that IPO stock prices usually decline following the period when there is a relatively higher demand for puts than calls. This finding is consistent with Cremers and Weinbaum (2010) and Xing, Zhang, and Zhao (2010), who show that the relative mispricing between put and call options have significant explanatory power on future returns of their underlying stocks.

Table 11 shows that the coefficient on *Insider holding*(t-1) is 0.022 and is statistically significant. The magnitude of this coefficient estimate is economically large. It suggests that one percent decrease in the number of shares held by insiders relative to the total shares outstanding is followed by a 2.2% decrease in adjusted returns over the next month. We therefore find strong evidence that IPO stock continually underperforms in response to decreasing insider shares holding in the months after options have been listed. We do not find that the level of short interest ratio affects IPO performance in the months after options have been listed. This finding is consistent with Figure 3 (top panel) which shows the change in short interest ratios that is due to option listing occurs mostly in the event month 0, i.e. when options are introduced. We conclude that the effect of option listing on the IPO firm equity's short sale constraint is confined to the months around the option listing date.

7 Option trading strategies on IPO firm equity

In this section, we examine whether substantial returns can be made by trading newly listed options on IPO stocks. We examine two simple strategies of buying put and selling call options when they become newly listed and holding them to expiration. Before we proceed, we point out two reasons why option trading may not necessarily result in significant returns even though we find overwhelming evidence that IPO stocks underperform after options have been listed on them. First, the payoff of an option contract depends on the underlying stock price. Consequently, option return depends on the raw returns of the underlying and not necessarily the benchmark-adjusted returns which we have primarily used in the previous sections. Second, the newly listed options are usually traded at a premium due to their relatively low liquidity and high implied volatilities making them too expensive to hold. Further, the bid-ask spreads of newly listed options may be too large for the trading strategy to be profitable.

We downloaded end-of-day option prices and their bid-ask spreads from Ivey OptionMetrics. We keep only newly listed contracts defined as those that are quoted within 30 days relative to the option introduction date. To account for bid-ask spreads, we assume that each contract is transacted at its lowest bid and highest ask prices. Mayhew (2002) and De Fontnouvelle, Fishe, and Harris (2003) show that the effective spreads for equity options is typically less than 0.5. Therefore, our assumption that we must transact options at their full bid-ask spreads is conservative. Stocks' daily close prices, and their daily highs and lows are obtained from CRSP. We assume that stocks are transacted at their effective bids and asks calculated using the method in Corwin and Schultz (2012). When effective bid-ask spreads cannot be calculated, we take a conservative approach by assuming that the underlyings are transacted at their daily highs and lows.

Table 12 reports average excess returns over the risk free rate from buying put options (Panel A) and writing newly listed call options (Panel B) on IPO stocks. The results are grouped according to the moneyness (F/K) and maturity (τ) of the option contract. We define the moneyness of each contract as F/K where F is the futures price of the underlying and K is the strike price.[15] Each cell group in Table 12 reports the result for a particular moneyness-maturity bin.

The raw return from buying each put option is calculated as

$$R_{put} = \frac{30}{\tau}\left(\frac{\max\left(K - S_{\tau}^{bid}, 0\right)}{P^{ask}} - 1\right),$$

where S_{τ}^{bid} is the bid price of the expiring stock at maturity τ and P^{ask} is the asking price of the put option bought from the market maker. The raw return from writing each call option is calculated as

$$R_{call} = \frac{30}{\tau}\left(1 - \frac{\max\left(S_{\tau}^{ask} - K, 0\right)}{C^{bid}}\right),$$

where S_{τ}^{ask} is the asking price of the expiring stock at maturity τ, and C^{bid} is the bid price of the written call option sold to the market maker. Option returns are scaled by $30/\tau$ in order to express them as monthly returns. Excess return for each option contract is then calculated by subtracting the return of a corresponding riskfree bond from the raw option return.

For each moneyness-maturity bin, average excess return is calculated as an equally-weighted portfolio of excess option returns across firms. Only one observation from each firm is used in

[15] The futures price of the underlying is calculated as $F = S_t \exp(r\tau)$, where S_t is the current spot price, r is the risk-free rate, and τ is the maturity of the contract.

the calculation. When there are multiple contracts traded in each moneyness-maturity bin, the contract with the highest open interests is used. For each moneyness-maturity bin, we report the average excess returns, t-statistics, and the number of firms used in the calculation. Table 12 shows that most of the newly listed option contracts have relatively short maturities, i.e. less then 60 days.[16]

Panel A of Table 12 shows that there is a significant return from buying long-maturity put options with days to maturity between 6 to 12 months. The profit from buying long-maturity puts is very robust across the moneyness bins. In fact, excess put returns appear to increase as the moneyness level (F/K) increases, that is, when the contracts are more out-of-the-money. Such a finding is supported by Coval and Shumway (2001) who theoretically show that expected put returns should increase as the contract become more out-of-the-money. For instance, Panel B shows the profit from buying long-maturity puts that are deep out-the-money ($1.15 < F/K \leq 1.20$) yields an average excess return of 6.35% per month. While we find a significant profit from buying long-maturity puts written on newly listed IPO stocks, Panel B shows that the average excess returns from writing call options are either insignificant or negative. We conclude that there is no profit from writing newly listed call options on IPO stocks. The inability of investors to make a profit from writing calls is important for reconciling the profit/loss from a market making perspective. That is, while market makers lose money by selling long-maturity put options, they do make profits from purchasing call options from the investors. Therefore, because the long and short positions of a derivative contract clear in the aggregate, market makers are likely to make profits for providing their services.

We next examine the cross-sectional determinants of returns generated from buying newly listed put options on IPO stocks. We focus our analysis only on long-maturity puts, defined as those with 6-12 months to maturity. Table 13 reports the results. The dependent variable is the monthly excess return from buying long-maturity puts. We control for the cross-sectional differences in returns across moneyness by including *Moneyness bin* variable in the regressions. *Moneyness bin* is a discrete variable with a value of 1 to 9 indicating the moneyness bin (from lowest to highest) associated with each option contract. The classification of each moneyness bin is shown in Table 12. For instance, *Moneyness bin* is equal to 1 if the option contract's moneyness falls in the $(0.80, 0.85]$ interval, and *Moneyness bin* is equal to 2 if the option contract's moneyness falls in the $(0.85, 0.90]$ interval, etc.

Regression model (1) in Table 13 reports the baseline estimates. The intercept in model (1) corresponds to the average excess returns from buying long-maturity puts in the lowest moneyness bin, i.e., $(0.80, 0.85]$. Its value is 2.29% which is roughly consistent with the results in Panel A of Table 12. Model (1) in Table 13 shows that the coefficient on *Moneyness bin*

[16] We exclude newly listed options with maturity greater than one year from the analysis because only two firms in our sample have such options traded.

is 0.427 and is highly significant. This finding suggests that monthly excess returns from our strategy is increasing in each moneyness bin by about 42.7 basis points.

In models (2)-(6), each IPO variable is added to the regression. We find positive and highly significant coefficients on *Venture backed* and *First day ret.* On the other hand, the coefficient on *Time-to-list* is negative and significant. Our simple strategy of buying long-term puts are therefore related to three IPO characteristics. This finding suggests that we can increase profits by selectively buying put options on certain IPO stocks. More specifically, we find that the profits from our trading strategy are more specific to venture backed IPOs that experience large first day returns and have options listed early in the IPO aftermarket.

8 Conclusion

Using a sample of IPOs from 1996 to 2008, we study four important questions regarding the listing of options on the equity of firms that have recently gone public.

First, we analyze the effects of option listing on an IPO firm's stock. We find that the equity underperforms in the short- as well as the long-run in the weeks following option listing. The significant underperformance subsequent to option listing is robustly observed throughout the sample and holds regardless of when options are listed (up to 3 years) after the IPO issuance. We also show that our findings are not an artifact of exchanges' ability to time the market using a propensity score matching analysis.

Second, we examine the determinants of exchanges' decision regarding when to list options on a firm's stock after the IPO. In addition to variables that have been documented (in the case of seasoned firms by Mayhew and Mihov (2004)) to have an impact on option listing, we examine whether IPO characteristics have an effect on the time to option listing. The results show that venture backed firms have options listed significantly earlier. Another important factor that significantly shortens the time to option listing is the size of IPO proceeds: greater the IPO proceeds, shorter the time to list.

Third, we test three possible hypotheses that can explain the strong underperformance of IPO equity following options listing. Our results show that: (i) The listing of options significantly relaxes short-sale constraints on the firm equity by providing investors with a relatively cheap channel to short sell the underlying shares; (ii) Insiders sell their shares on IPO firms more aggressively after options have been listed further exacerbating the long-run IPO performance; and (iii) Put options remain significantly in greater demand than call options for several months after option listing suggesting that investors are synthetically shorting the underlying IPO stocks through put option contracts.

Finally, we examine the ability to make abnormal profits net of bid ask spreads by investing in newly listed options on the IPO firm's stock. The two option trading strategies that we examine are writing call options or buying put options on IPO firm stock when they are newly listed and holding them to expiration. We find significant excess returns from buying long-maturity newly listed puts on IPO firm stock.

To the best of our knowledge, this paper is the first to study the economic effects of option listing on IPOs. Consequently, our work contributes to the literature in several ways. First, we document the long-run equity underperformance of firms following option listing. While existing studies have documented short-run listing effects, to the best of our knowledge, there is no paper in the literature documenting a long-run listing effect even in the context of seasoned firms. Second, this paper is the first to analyze the IPO characteristics associated with early option listing. Third, we document that option listing relaxes the short-sale constraint on IPO firm equity. Fourth, by testing new hypotheses, we are able to develop additional insight into the causes underlying the stock return underperformance of firms following option listing. Finally, we are the first to analyze the profitability of option trading strategies on IPO firm equity.

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Appendix

A: Propensity score matching method

This section discusses the propensity score matching of options listed-IPO firms to non-optioned IPO firms in the control group. We estimate the firm i's ex ante probabilities of options listed in month t using the following logit model

$$\begin{aligned} L\left(List_{i,t}\right) &= \alpha_0 + \alpha_1 \textit{Venture backed}_i + \alpha_2 \textit{High underwriter rep}_i + ... \\ &\quad \alpha_3 \textit{First dat ret}_i + \alpha_4 \textit{Proceeds amount}_i + \beta_1 \textit{Market cap}_{i,t-1} + ... \\ &\quad \beta_2 \textit{Shares turnover}_{i,t-1} + .\beta_3 \textit{Volatility}_{i,t-1} + \beta_4 \textit{Cumulative returns}_{i,t-1}, \end{aligned} \tag{A.1}$$

where $L\left(List_{i,t}\right)$ is the log-odds ratio that the i stock will be selected for option listing in month t. The independent variables that we use are IPO characteristics and time-varying covariates that significantly determine time-to-list options on IPO stocks (see Table 3). The IPO characteristics that we use for propensity score matching are: *Venture backed*, *High underwriter rep*, *First day ret*, and *Proceeds amount*. We do not consider *Time to lockup expiration* in the matching because Table 3 shows that they are irrelevant for time-to-list options. Following Mayhew and Mihov (2004), we include one-month lagged time-varying covariates which are firm's size, trading volume, volatility, and momentum. Section 4 discusses the construction of these variables.

We estimate the regression model (A.1) on firms that went public in 1996-2008 and have option listed within 48 months after their IPO issuance date. We remove firms once they have become seasoned for more than 48 months to ensure that we are matching among relatively recent IPO firms. We also require that firms in our matching sample eventually have option listed to ensure that they are, at some point, eligible for option listing. Overall, the pre-matching sample consists of 1,173 firms corresponding to 19,758 firm-month observations. Regression model (1) in Table A1 reports logistic regression results of (A.1) for the pre-matching, i.e. "Before matching", sample.

Table A1 shows the scaled pseudo adjusted R-squared for the pre-matching sample is 13.62%. All the IPO characteristics and time-varying covariates in the model are statistically significant. The signs on these regression coefficients are also consistent with the results in Table 3 which estimates the determinants of time-to-list using an extended Cox model on a much larger firm sample. This finding confirms that our results on the determinant of option listing are robust to different IPO sample sizes as well as using a different estimation model.

Using the estimated ex ante probabilities of options listed, we match each option-listed IPO firm (treated group) to another IPO firm that does not yet have options listed (Control group). The matching is done at the firm-month level. We use is the nearest-neighbourhood caliper matching approach of Cochran and Rubin (1973). We exclude observations with propensity score below 5 percent from the matching. We require that the propensity score of the control observation is within $\pm 5\%$ of that of the treated observation. Importantly, for each matched pair, we require that the treated observation is matched to an IPO firm that *will not* have options listed on their equity for at least 12 months. This last requirement is critical because our objective is to measure the difference in long-run returns (up to one year) between IPO firms that experience and do not experience options listing. Therefore, requiring that control-group firms do not experience options listing for at least one year ensures that their post-matching returns are not impacted by their future possibility of having option listed. Due to the small number of IPO firms eligible for the control group, we allow each control firm to be matched with up to two treated-group firms. The final matching yields 541 matched pairs, where 541 unique treated-group firms are matched to 349 unique control-group firms.

The regression model (2), titled "After matching", in Table A1 reports results from estimating the logit model in (A.1) on the propensity score-matched pairs. The intercept of the regression model is no longer statistically significant compared to the "Before matching" sample. Most of the IPO characteristics and time-varying covariates lose their statistical significance after the matching. *First day ret* and *Volatility* are two only two variables that remain weakly significant. Nevertheless, we control for any remaining imperfect matching between the treated and control group by including all the independent variables in (A.1) when estimating the difference-in-difference regression model in Table 6. Overall, the results from regression model (2) show that the treated and control firm observations are quite similar in term of IPO characteristics and their time-series properties.

Table A1. *Propensity score matched firms*

We estimate the firm's propensity of having options listed on any given month following the logistic regression model in (A.1).This table summarizes the results. The regression model (1), titled "Before matching", reports results for the propensity score model used in matching. The sample consists of firms that have gone public in 1996-2008 and have option listed within 48 months after their IPO issuance. When firms become seasoned for more than 48 months, their observations are removed from the sample to ensure that we are matching among relatively recent IPO firms. The regression model (2), titled "After matching", reports the results for propensity-score-matched firms from the treated and control groups. The treated group consists of IPO firms in 1996-2008 that have options listed within 36 months after the issuance date. Firms in the treated group are matched to IPO firms that do not yet have options listed (Control group) based on their propensity of having options listed. The matching is done at the firm-month level using relevant IPO characteristics as well as monthly lagged time-varying covariates, i.e., size, trading volume, volatility, and momentum (see Table 3 for descriptions).Robust t-statistic is reported in bracket below each estimate. Number of observations refers to the number of firm-month used in the estimation. Number of events refers to the number option listings observed in the sample. Year and industry fixed effects are included. We report the scaled pseudo adjusted R-squared for each regression model.

	Before matching	After matching
	(1)	(2)
Intercept	-23.663***	3.834
	(-27.82)	(0.05)
IPO characteristics		
Venture backed	0.302***	-0.002
	(4.13)	(-0.01)
High underwriter rep	-0.179**	0.062
	(-2.51)	(0.42)
First day ret	-0.003**	0.010*
	(-2.20)	(1.96)
Proceeds amount	0.181***	0.040
	(3.40)	(0.33)
Time-varying covariates		
Market cap	0.927***	-0.119
	(22.17)	(-0.81)
Shares turnover	0.064*	0.026
	(1.87)	(0.49)
Volatility	0.039***	0.100*
	(3.26)	(1.92)
Cumulative retuns	0.263***	0.356
	(3.93)	(1.16)
Year and Ind fixed effects	YES	YES
Number of obervations	19758	1076
Number of events	1173	541
Pseudo R^2	13.62%	9.33%

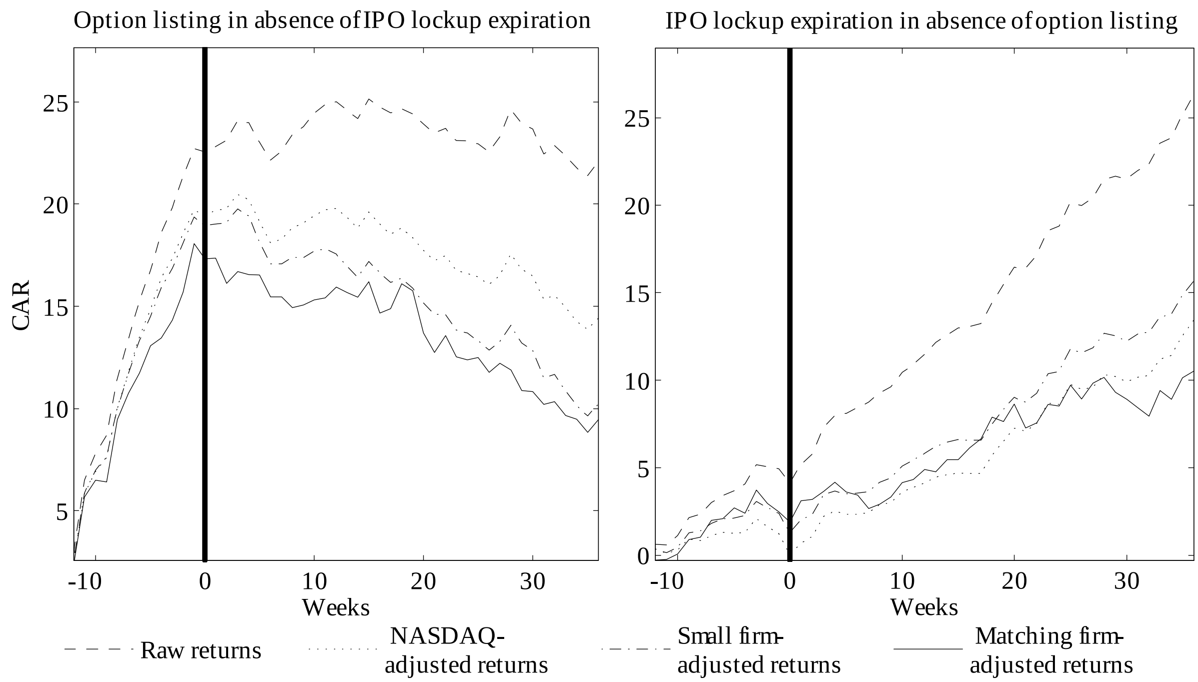

Figure A1. Option listing versus IPO lockup expiration. The initial sample contains initial public offerings in 1996-2008. We plot cumulative average adjusted returns centered on two events: IPO option listing date and IPO lock expiration date. Cumulative average adjusted returns (CAR) are computed weekly centered on the option listing date. Week 0 refers to the option-listing window and is computed from trading day -2 to day +2 relative to the option listing date. In the left panel, we consider 581 option listing events occurring within 36 months after the IPO issuance date that do not coincide with IPO lock expirations within one calender year. In the right panel, we consider 434 IPO lock expirations that do not coincide with option listing events within one calender year. In each panel, we plot CAR series starting 12 weeks before option listing dates. Four CAR series are plotted: 1) raw returns, 2) CRSP-value-weighted NASDAQ index adjustment (NASDAQ-adjusted), lowest decile of NYSE market capitalization index adjustment (small firm-adjusted), and 4) matching firm adjustment (matching firm-adjusted).

Table 1. *Summary statistics.*
The sample consists of US initial public offerings on ordinary common shares between January 1996 and December 2008, excluding closed-end funds, REIT, Dutch auction offerings, and issuances flagged with "shelf registration". We obtain option listing dates for 1996-2011 from OptionMetrics and CBOE website. Panel A summarizes the distribution of initial public offerings by issuance year grouped according to their option listing date (in months) relative to their IPO issuance date. Panel B reports descriptive statistics of the IPO variables. *Days-to-list* is the number of days between IPO issuance and when option is introduced, conditional on observing a listing. *Venture backed* is an dummy variable equal to 1 if the IPO firm is backed by a venture capitalist. *Underwriter rep* is the reputational score of the IPO's lead underwriter according to the ranking published on Jay Ritter's website. *First day ret* is computed as the log of return (%) on the offered price to the end of the the first trading day. *Proceed amount* is the log of the total proceeds amount from the offering. *Total shares offered* is the number of shares issued at the offering in thousands. *Lock up agreement* is a dummy variable equal to 1 if the IPO has a share lockup agreement, and zero otherwise. Conditional on having shares lockup agreement, *Days to lockup expiration* is the number of days from IPO issuance date to the first shares lockup expiration.

Panel A: Distribution of time-to-list options after initial public offerings

Year IPO issuance	Months relative to the IPO issuance date							Never optioned	Total
	6 or less	7 to 12	13 to 18	19 to 24	25 to 30	31 to 36	37 or more		
1996	24	53	41	34	29	13	56	278	528
1997	29	35	22	14	8	8	43	199	358
1998	21	38	7	5	7	5	25	118	226
1999	89	103	21	8	5	4	23	133	386
2000	50	57	13	11	3	3	39	123	299
2001	27	11	1	1	1	4	11	10	66
2002	24	4	8	1	2	0	7	14	60
2003	7	6	5	1	1	5	13	16	54
2004	29	19	14	8	12	10	15	47	154
2005	30	9	9	7	5	4	14	37	115
2006	31	16	15	4	3	3	13	38	123
2007	22	24	7	3	9	4	13	39	121
2008	7	1	0	0	0	1	1	3	13
Overall	390	376	163	97	85	64	273	1055	2503

Panel B: Descriptive statistics

Variable	Nobs	Mean	Median	Std Dev	Minimum	Maximum
Days-to-list	1448	672	321	862	9	5670
Venture backed	2503	0.48	0	0.50	0	1
Underwriter rep	2503	7.55	8	1.95	0	9
First day ret (%)	2503	19.87	10.95	29.23	-56.69	134.60
Offer price	2503	13.48	13.00	5.30	4.00	70.41
Proceeds amount	2503	17.83	17.76	1.01	14.58	23.09
Total shares offered (×1000)	2503	6790	4200	12878	365	360000
Lockup agreement	2503	0.76	1.00	0.43	0	1
Days to lockup expiration	1906	194	180	72	30	730

Table 2. *Determinants of time-to-list*

We report results from running the extended Cox model on time-to-list options on IPO stocks. The sample includes initial public offerings in 1996-2008. The dependent variable, *Time-to-list*, is the number of months between IPO issuance date and the first date of option introduction. Time-to-list beyond 36 months are truncated. The explanatory variables include IPO characteristics and monthly time-varying control variables. *Venture backed* is an dummy variable equal to 1 if the IPO firm is backed by a venture capitalist. *Time to lockup expiration* is the number of months from IPO issuance date to the first shares lockup expiration. *High underwriter rep* is a dummy variable equal to 1 if the firm's lead underwriter is in the highest prestige category (rank=9) according to the ranking published on Jay Ritter's website. *First day ret* is computed as the log of return on the offered price to the end of the the first trading day. *Proceed amount* is the log of the total proceeds amount from the offering. Following Mayhew and Mihov (2004), we include monthly time-varying variables that have been shown to explain Exchange's decision to list options. *Market cap* is the log of market capitalization observed at the end of the previous month. *Shares turnover* is the average of daily shares turnover over the previous month (%). *Volatility* is the average daily standard deviations of raw returns over the previous month (%). *Cumulative ret* is the cumulative daily raw returns over the previous month. Number of observations refers to the number of firm-month used in the estimation. Number of events refers to the number option listings available for the estimation. All regressions include industry and year fixed effects. We report robust t-statistics in bracket below each estimate.

	(1)	(2)	(3)	(4)	(5)
IPO Characteristics:					
Venture backed	0.506***				0.589***
	(6.78)				(7.54)
Time to lockup expiration		0.006			0.005
		(1.03)			(0.16)
High underwriter rep			-0.161**		-0.177**
			(-2.14)		(2.16)
First day ret (%)				-0.002*	-0.004***
				(1.80)	(2.71)
Proceeds amount	0.289***	0.213***	0.234***	0.197***	0.327***
	(5.04)	(3.84)	(4.13)	(3.54)	(5.40)
Time-varying covariates:					
Market cap	1.321***	1.326***	1.341***	1.348***	1.393***
	(32.42)	(32.03)	(32.21)	(31.23)	(30.05)
Shares turnover	0.102***	0.100***	0.100***	0.102***	0.108***
	(7.29)	(7.30)	(7.22)	(7.40)	(7.36)
Volatility	0.021*	0.028***	0.028***	0.030***	0.021*
	(1.93)	(2.67)	(2.66)	(2.85)	(1.83)
Cumulative retuns	1.131***	1.138***	1.140***	1.131***	1.173***
	(9.13)	(9.14)	(9.17)	(9.09)	(9.16)
Year and Ind fixed effects	Yes	Yes	Yes	Yes	Yes
Number of obervations	52569	52569	52569	52569	52569
Number of events	1087	1087	1087	1087	1087

Table 3. *Cumulative abnormal returns around option listing*
We report cumulative average adjusted returns for IPO stocks over the (-12, +52) weekly intervals centered on their option listing date. The sample contains initial public offerings in 1996-2008 that have options listed within 36 months after the firms have gone public. Panel A report results for all observations. Panel B reports results grouped by firms that have options listed in 0-6 months, 7-12 months, and 13-36 months after their IPO issuance date. Panel C reports results grouped by IPO issuance year, i.e., 1996-1998, 1999-2002, 2003-2005, and 2006-2008. The cumulative average adjusted return, CAR_{t_1,t_2}, from week t_1 to t_2 relative to the option listing date is computed as $CAR_{t_1,t_2} = \sum_{t=t_1}^{t_2} AR_t$, where AR_t is the average matching firm-adjusted return on week t across firms. Weekly adjusted return for each firm is calculated by substracting its raw return with the return of a benchmark firm matched by industry and size. We report robust t-statistic in bracket below each estimate. The t-statistic for the cumulative average adjusted return, CAR_{t_1,t_2}, is computed as $CAR_{t_1,t_2}\sqrt{n_{t_1,t_2}}/csd_{t_1,t_2}$ where n_{t_1,t_2} is the number of firms trading with non-missing observations between week t_1 and t_2. The cross section standard deviation, csd_{t_1,t_2}, is computed as $csd_{t_1,t_2} = [(t_2 - t_1) \cdot var + 2 \cdot (t_2 - t_1 - 1) \cdot cov]^{1/2}$, where var and cov are the average cross-sectional variance and the first-order auto covariance of the AR_t series over the $(-12, +104)$ intervals, respectively. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

		Panel A: All observations		Panel B: Grouped by when options are listed after the IPO					
				0-6 months		7-12 months		13-36 months	
		Nobs	CAR(t_1,t_2) %	Nobs	CAR(t_1,t_2) %	Nobs	CAR(t_1,t_2) %	Nobs	CAR(t_1,t_2) %
(1)	[-12, -2]	1075	12.29***	345	14.46***	364	7.48**	366	15.67***
			(7.61)		(5.46)		(2.48)		(5.74)
(2)	[-1,+1]	1111	-0.17	383	-1.53	364	-1.26	364	2.33
			(-0.21)		(-1.16)		(-0.80)		(1.63)
(3)	[+2,+12]	1092	-2.05	381	-5.64**	359	1.09	352	-1.38
			(-1.28)		(-2.24)		(0.36)		(-0.50)
(4)	[+13,+52]	978	-17.63***	355	-20.70***	317	-19.33***	306	-12.61**
			(-5.46)		(-4.16)		(-3.14)		(-2.22)
(3) – (1)			-14.34***		-20.10***		-6.39*		-17.05***
			(-8.44)		(-7.66)		(-1.84)		(-6.31)

		Panel C: Grouped by IPO issuance year							
		1996–1998		1999–2002		2003–2005		2006–2008	
		Nobs	CAR(t_1,t_2) %	Nobs	CAR(t_1,t_2) %	Nobs	CAR(t_1,t_2) %	Nobs	CAR(t_1,t_2) %
(1)	[-12, -2]	361	13.36***	424	12.31***	162	10.82***	128	12.89***
			(4.76)		(4.15)		(3.64)		(3.32)
(2)	[-1, +1]	361	1.78	430	-2.74*	175	1.63	145	-0.49
			(1.22)		(-1.78)		(1.09)		(-0.26)
(3)	[+2, +12]	353	-1.22	424	-3.06	172	1.07	143	-3.62
			(-0.43)		(-1.03)		(0.37)		(-0.98)
(4)	[+12, +52]	303	-4.96	376	-37.85***	160	-2.26	136	-11.32
			(-0.85)		(-6.32)		(-0.40)		(-1.47)
(3) – (1)			-14.58***		-5.18***		-9.75***		-16.51***
			(-4.76)		(-5.18)		(-3.60)		(-4.62)

Table 4. *Regression analysis: Weekly adjusted returns around option listing*
The dependent variable is weekly adjusted return, $ar_{i,t}$, for firm i in event week t. It is calculated as the weekly raw return on stock i minus the corresponding period raw return of a benchmark firm matched by industry and size. Panel A reports results for the balanced window [-12,+12], and Panel B reports results for the long-run window [-12,+52]. The sample consists of IPO firms in 1996-2008 that have options listed within 36 months after the issuance date. $Optionstat$ is a dummy variable equal to 1 when event week $t \geq 0$ indicating that options have been listed, and zero otherwise. Models (1) and (3) report the base line regression results. Models (2) and (4) report estimates for the interactions between $Optionstat$ and various IPO characteristics. $Venture\ backed$ is an dummy variable equal to 1 if the IPO firm is backed by a venture capitalist. First day ret is computed as the log of return on the offered price to the end of the first trading day. $High\ underwriter\ rep$ is a dummy variable equal to 1 if the firm's lead underwriter is in the highest prestige category (rank=9) according to the ranking published on Jay Ritter's website. $Time\text{-}to\text{-}list$ is the number of months from IPO issuance date to when option is listed. The dummy variable *Lockup expiration(t)* controls for the impact of shares lockup expiration; it is equal to 1 if the IPO firm has a shares lockup agreement expiring on week t, and zero otherwise. IPO characteristics, industry and year fixed effects are included. See text for complete details. Weekly controls consist of weekly lagged variables: adjusted return, average turnover, average volatility, and average market capitalization. Robust t-statistic, clustered at the firm-level, are reported in bracket below each estimate. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

	Panel A: Weeks [-12, +12]		Panel B: Weeks [-12, +52]	
	(1)	(2)	(3)	(4)
Optionstat	-1.234***	-1.723***	-1.486***	-1.382***
	(-6.35)	(-3.43)	(-8.77)	(-3.06)
Optionstat ×				
Venture backed		-1.201***		-1.019***
		(-3.09)		(-3.02)
First day ret		0.010		0.005
		(1.37)		(0.82)
High underwriter rep		0.406***		0.699**
		(2.97)		(1.97)
Time-to-list		0.016		-0.006
		(0.62)		(-0.26)
Venture backed	-0.047	0.609*	-0.064	0.609*
	(-0.22)	(1.89)	(-0.48)	(1.89)
First day ret	-0.009**	-0.014**	-0.004*	-0.008
	(-2.10)	(-2.36)	(-1.65)	(-1.41)
High underwriter rep	0.169	0.170	0.146	-0.431
	(0.75)	(0.75)	(1.15)	(-1.31)
Time-to-list	0.019	0.012	0.007	0.012
	(1.48)	(0.55)	(0.98)	(0.59)
Lockup expiration (t)	-3.166***	-3.224***	-2.190***	-2.243***
	(-3.27)	(-3.32)	(-3.02)	(-3.09)
Intercept	3.156*	2.099	1.718	1.692
	(1.70)	(1.11)	(1.59)	(1.47)
Year and Ind fixed effects	Yes	Yes	Yes	Yes
Weekly controls	Yes	Yes	Yes	Yes
Number of clusters	25413	25413	66751	66751
Number of clusters	1121	1121	1121	1121
Adjusted R^2	0.79%	0.84%	0.73%	0.75%

Table 5. *Propensity score-matched sample: Difference-in-difference regression*

We report difference-in-difference (DID) regression results on weekly adjusted returns around option listing for the propensity-score matched sample, i.e., treated and control groups. Weekly adjusted return, $ar_{i,t}$, for firm i in event week t is calculated as the weekly raw return on stock i minus the corresponding period raw return of a benchmark firm matched by industry and size. We report results for three observation windows. Regression model (1) reports results for the pre-event window which includes week [-12,-1] relative to when options are listed. Regression models (2) and (3) report results over [-12,+12] and [-12, +52] weeks relative to option listing, respectively. The treated group consists of IPO firms in 1996-2008 that have options listed within 36 months after the issuance date. Firms in the treated group are matched to IPO firms that do not have options listed (Control group) based on their propensity score of having options listing. We require that firms in the control group must not have been public for more than four years and do not experience option listing for at least one year after they have been matched. The matching is done at the firm-month level based on IPO characteristics and monthly time-varying covariates (see Table A1). *Optionstat* is a dummy variable equal to 1 when event week $t \geq 0$ indicating that options have been listed, and zero otherwise. *Treated* is a dummy variable equal to 1 if the firm is in the treated group, and zero for the control group. The variable of interest is the DID estimator, *Treated* $\times$ *Optionstat*, which is reported for regression specifications (2) and (3). IPO characteristics, year and industry fixed effects, as well as weekly control variables (unreported) are inlcuded in the regressions. The dummy variable *Lockup expiration(t)* controls for the impact of shares lockup expiration; it is equal to 1 if the IPO firm has a shares lockup agreement expiring on week t, and zero otherwise. Weekly controls consist of the following one-week lagged variables: adjusted return ($ar_{i,t-1}$), average turnover, average volatility, and average market capitalization. Robust t-statistic, clustered at the matched-pair level, are reported in bracket below each estimate. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

	Weekly adjusted returns		
	Weeks [-12,-1]	Weeks [-12, +12]	Weeks [-12, +52]
	(1)	(2)	(3)
Treated × Optionstat		-0.750**	-0.994***
		(-2.19)	(-2.81)
Optionstat		-0.258	-0.537**
		(-1.00)	(-2.06)
Treated	0.117	0.235	0.217
	(0.32)	(0.71)	(0.67)
		0.00	0.00
Venture backed	-0.637*	-0.496**	-0.168
	(-1.84)	(-2.13)	(-0.93)
High underwriter rep	1.109**	1.010***	0.411**
	(2.13)	(4.16)	(2.20)
First day ret	0.002	0.005	0.000
	(0.22)	(0.94)	(0.03)
Lockup expiration (t)	-3.833***	-3.620**	-2.322***
	(-3.46)	0.000	(-2.75)
Year and Ind fixed effects	Yes	Yes	Yes
Weekly controls	Yes	Yes	Yes
Number of obervations	8998	19490	49712
Number of clusters	543	543	543
Adjusted R^2	1.40%	0.81%	0.68%

Table 6. *Changes in short interest ratio around option listing*

We report changes in short interest ratios relative to event month -4 around option listing date. Short interest ratio is the ratio of short interest to the number of shares outstanding. We obtain short interest levels from Shortsqueeze and COMPUSTAT which reports the number of shares held short at the end of each month after 2004. *Rel Change* is the relative percentage change in short interest ratio; it is calculated as $\left(\frac{\text{Short Int Ratio(t)}}{\text{Short Int Ratio(-4)}} - 1\right) \times 100$ where Short Ratio(t) is the short interest ratio observed at the end of the event month t relative to the option listing date. We exclude firms that do not have reported short interest on the event month -4 from the analysis. The treated group consists of IPO firms in 1996-2008 that have options listed within 36 months after the issuance date. Ofek and Richardson (2003) find that short interest level increases in the post lock-up period, we therefore exclude IPO firms in the treated group that have shares lockup expiration within 6 months of option listing date. Firms in the treated group are matched to IPO firms that do not have options listed (Control group) based on their propensity score of having options listed. The matching is done at the firm-month level. We require that firms in the control group must not have been public for more than four years and will not experience option listing for at least one year after they have been matched. The matching is done based on IPO characteristics as well as monthly time-varying covariates (see Table A1). *Diff Rel Change* reports the difference in *Rel Change* between the treated and the control group at each event month. For each event month, we report p-values from the two nonparametric tests: (i) Kolmogorov-Smirnov (KS) test that the distributions of the treated and control groups are equal; and (ii) One-way Wilcoxon signed-rank test that the distribution median of the treated group is greater than that of the control group. The reported values of *Rel Change* are winsorized at the 1 and 99 percentiles to remove outliers.

	Changes in short interest ratio relative to event month -4 (%)						
	Treated		Control		Treated–Control		Treated > Control
Month of seasoning	Nobs	Rel change	Nobs	Rel change	Diff Rel Change	KS Test P-value	Wilcoxon Test P-value
-3	118	27.8	99	18.23	9.57	0.064	0.088
-2	118	80.7	99	85.03	-4.33	0.534	0.609
-1	118	96.5	100	126.47	-29.95	0.798	0.485
0	115	268.2	97	135.26	132.95	0.002	0.009
1	117	285.2	101	124.82	160.33	0.012	0.012
2	117	256.5	98	143.58	112.94	0.001	0.004
3	117	290.5	98	172.12	118.34	0.015	0.015
4	118	342.6	99	203.67	138.89	0.064	0.114
5	113	318.0	97	267.92	50.10	0.112	0.171
6	115	334.2	96	313.00	21.21	0.220	0.110

Table 7. *Changes in insider holding around option listing*
We report changes in insider holding relative to event month -4 around option listing date. We obtain insider filing data from Thomson Reuter's Insider Filing Data Feed (IFDF). For each month, we compute the total direct and indirect holding of conventional stock (Table1 in IFDF) by the company's insiders with relationship hierarchy level between 1 to 3. Rel Change is the relative percentage change in total insider holding; it is calculated as $\left(\frac{\text{Insider holding(t)}}{\text{Insider holding(-4)}} - 1\right) \times 100$ where Insider holding(t) is the total insider holding at the end of the event month t relative to the option listing date. The treated group consists of IPO firms in 1996-2008 that have options listed within 36 months after the issuance date. We exclude listings that overlap with lockup expiration days within one calendar year to avoid the impact associated with the IPO lockup agreement (Field and Hanka (2001)). Firms in the treated group are matched to IPO firms that do not have options listed (Control group) based on their propensity score of having options listing. The matching is done at the firm-month level. We require that firms in the control group must not have been public for more than four years and will not experience option listing for at least one year after they have been matched. The matching is done based on IPO characteristics as well as monthly time-varying covariates (see Table A1). Diff Rel Change reports the difference in Rel Change between the treated and the control group at each event month. For each event month, we report p-values from the two nonparametric tests: (i) Kolmogorov-Smirnov (KS) test that the distributions of the treated and control groups are equal; and (ii) One-way Wilcoxon signed-rank test that the distribution median of the treated group is less than that of the control group. The reported values of Rel Change are winsorized at the 1 and 99 percentiles to remove outliers.

	Changes in insider holding relative to event month -4 (%)						
	Treated		Control		Treated - Control		Treated < Control
Month of seasoning	Nobs	Rel change	Nobs	Rel change	Diff Rel Change	KS Test P-value	Wilcoxon Test P-value
-3	263	-1.314	149	-1.16	-0.16	0.306	0.011
-2	262	-2.6521	151	-1.84	-0.81	0.637	0.178
-1	263	-4.8636	148	-2.85	-2.02	0.655	0.263
0	261	-9.3721	147	-3.65	-5.72	0.009	0.002
1	264	-12.1411	149	-4.85	-7.29	0.029	0.003
2	259	-15.2476	143	-5.56	-9.69	0.002	>0.001
3	260	-16.8135	143	-5.94	-10.87	>0.001	>0.001
4	262	-17.7859	144	-5.21	-12.57	>0.001	>0.001
5	258	-19.5066	141	-5.15	-14.36	>0.001	>0.001
6	256	-20.8188	140	-6.34	-14.48	>0.001	>0.001

Table 8. *Relative put-call demand*

We consider initial public offerings in 1996-2008 that have options listed within 36 months after the issuance date. The table reports two monthly measures of relative put-call demand: (i) Adjusted implied-volatility spread (Cremers and Weinbaum (2010)), and (ii) Adjusted implied-volatility skewness (Xing, Zhang, and Zhao (2010)). We report average daily values (in %) calculated over each monthly interval relative to when options are introduced. Event month 0 consists of the first 21 trading days following the option listing date. Adjusted IV spread for the firm i on day t is computed as $IVspread_{i,t} - \overline{IVspread_t}$, where $IVspread_{i,t}$ is the weighted difference between IVs of all put-call pairs quoted on firm i on day t. The market IV spread, $\overline{IVspread_t}$, is computed by averaging $IVspread_{i,t}$ across all available stocks on day t. Adjusted IV skewness for the firm i on day t is computed as $IVskewness_{i,t} - \overline{IVskewness_t}$, where $IVskewness_{i,t}$ is the difference between IVs of at-the-money calls and out-of-the-money puts for the firm i on day t. The market IV skewness, $\overline{IVskewness_t}$, is computed by averaging $IVskewness_{i,t}$ across all available stocks on day t. $Nobs$ reports the number of firms used in the monthly calculation. T-statistic is reported in bracket below each estimate. Superscripts ***, ** and * indicate significance at 1 , 5 and 10 percent-levels for a two-tailed test.

Month of seasoning	Panel A: Adjusted IV spread (%) Cremers and Weinbaum (2009)		Panel B: Adjusted IV Skewness (%) Xing, Zhang and Zhou (2010)	
	Nobs	Mean	Nob	Mean
0	901	-1.40***	811	-1.61***
		(-6.39)		(-5.65)
1	974	-1.60***	900	-1.28***
		(-7.00)		(-4.45)
2	990	-1.45***	927	-0.91***
		(-7.57)		(-2.58)
3	1009	-0.96***	934	-0.91*
		(-5.69)		(-1.75)
4	1010	-0.73***	938	-0.98**
		(-3.43)		(-2.08)
5	984	-0.72***	917	-0.94**
		(-3.24)		(-2.14)
6	956	-0.60*	890	-0.33
		(-1.75)		(-0.66)
7	944	-0.55***	870	-0.59
		(-2.61)		(-1.25)
8	929	-0.41**	848	-0.66
		(-2.00)		(-1.49)
9	906	-0.43**	831	-0.81*
		(-2.16)		(-1.68)
10	871	-0.20	794	-0.12
		(-1.01)		(-1.43)
11	864	-0.40	776	-0.64
		(-1.58)		(-1.22)
12	848	-0.37	758	-0.47
		(-1.35)		(-0.83)

Table 9. *Explaining abnormal returns around option listing*

We report regression results on monthly adjusted returns from event month -3 to event month +12. Event month 0 consists of 21 trading days centered on the option listing date. Other event months are defined as successive 21-trading-day periods relative to event month 0. Monthly adjusted return, $ar_{i,t}$, for firm i in event month t is calculated as the monthly raw return on stock i minus the corresponding period raw return of a benchmark firm matched by industry and size. The sample consists of IPO firms in 1996-2008 that have options listed within 36 months after the issuance date. *Optionstat* is a dummy variable equal to 1 when event month $t \geq 0$ indicating that options have been listed, and zero otherwise. We study factors influencing the magnitude of option listing on IPO underperformance by interacting *Optionstat* with variables calculated in Sections 6.1−6.3. *IVspread*(1) is the average adjusted implied volatility spread observed in event month 1, and *IVskewness*(1) is the average value of adjusted implied volatility skewness observed in event month 1. Δ *Short interest ratio*(1) is the percentage change in short interest ratios in event month 1 relative to event month -4 (see Table 7). Δ *Insider holding*(1) is the percentage change in insider shares held at the end of event month 1 relative to event month -4 (see Table 8). Year fixed effects and various IPO characteristics are included in the regressions. We include one-month lagged adjusted returns and average firm volatility as monthly control variables. Robust t-statistic, clustered at the firm-level, is reported in bracket below each estimate. Superscripts ***, **, and * indicate significance at the 1, 5, and 10 percent levels for a two-tailed test

	Dependent variable: Monthly adjusted returns from event months -3 to +12			
	(1)	(2)	(3)	(4)
Optionstat	-5.588***	-5.681***	-4.087***	-6.357***
	(-6.25)	(-6.10)	(-3.80)	(-8.10)
Optionstat × IV spread(1)	0.334*			
	(1.70)			
Optionstat × IV skewness(1)		0.233**		
		(2.35)		
Optionstat × Δ Short interest ratio(1)			-0.003**	
			(-2.50)	
Optionstat × Δ Insider holding(1)				0.018
				(0.56)
IV spread(1)	-0.292			
	(-1.43)			
IV skewness(1)		-0.179		
		(1.40)		
Δ Short interest ratio(1)			0.001	
			(1.14)	
Δ Insider holding(1)				-0.007
				(-0.26)
Intercept	11.433**	12.616**	6.285	9.589*
	(2.28)	(2.39)	(0.74)	(1.65)
IPO characteristics	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Monthly controls	Yes	Yes	Yes	Yes
Numer of obervations	13124	12079	4088	11193
Number of clusters	933	859	270	734
Adjusted R^2	0.96%	1.14%	0.70%	0.70%

Table 10. *Explaining long-run IPO underperformance post-option listing*

We report regression results for long-run IPO abnormal returns in the post-option listing period. The dependent variable is monthly adjusted return from event month +2 to event month +12. We define each event month as a successive 21-trading-day period relative to when option is listed. Monthly adjusted return, $ar_{i,t}$, for firm i in event month t is calculated as the monthly raw return on stock i minus the corresponding period raw return of a benchmark firm matched by industry and size. The sample consists of IPO firms in 1996-2008 that have options listed within 36 months after the issuance date. We regress $ar_{i,t}$ on monthly lagged variables calculated in Section 6. *IVspread*(t-1) is the lagged monthly average value of adjusted implied volatility spread. *IVskewness*(t-1) is the lagged monthly average value of adjusted implied volatility skewness Section 6.3 explains the calculations of IV spread and IV skewness. *Short interest ratio*(t-1) is the lagged monthly level of short interest ratio. *Insider holding*(t-1) is the lagged monthly percentage level of total shares held by insiders. We include year fixed effects and various IPO characteristics in the regressions. Monthly controls consists of lagged monthly adjusted returns and average firm volatility. Robust t-statistic, clustered at the firm-level, is reported in bracket below each estimate. Superscripts ***, **, and * indicate significance at the 1, 5, and 10 percent levels for a two-tailed test..

	Dependent variable: Monthly adjusted returns from event months +2 to +24			
	(1)	(2)	(3)	(4)
IV spread(t-1)	6.416*			
	(1.67)			
IV skewness(t-1)		6.895***		
		(3.87)		
Short interest ratio(t-1)			-0.034	
			(-0.67)	
Insider holding(t-1)				0.022**
				(2.41)
Intercept	0.441	-1.985	-0.949	2.873
	(0.12)	(-0.52)	(-0.23)	(0.77)
IPO characteristics	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Monthly controls	Yes	Yes	Yes	Yes
Numer of obervations	16450	15534	7943	15790
Number of clusters	1007	1004	440	864
Adjusted R^2	0.60%	0.74%	0.09%	0.74%

Table 11. *Returns from buying puts and selling calls on IPO stocks*
This table summarizes the average monthly returns from trading in newly listed put and call options on IPO stocks. We consider initial public offerings in 1996-2008 that have options listed within 36 months after the issuance date. We define newly listed options as those that are quoted within 30 days after the first option contract is introduced. Panel A reports monthly average returns over the risk free rate from a trading strategy that buys put options. Panel B reports monthly average returns over the risk free rate from a trading strategy that writes call options. Option returns are expressed in monthly percentage term and are sorted by days to maturity and moneyness (F/K), where F is the futures price of the underlying spot. We include transaction costs in the return computation by assuming that options and the underlyings are traded at their bid and ask prices. We assume that options are held until their expiration. T-statistics and the number of observations are reported below each estimate. Superscripts ***, ** and * indicate significance at 1, 5 and 10 percent-levels.

	Panel A: Average monthly excess returns from buying put options (%)				Panel B: Average monthly excess returns from writing call options (%)			
Moneyness bins	Days to maturity				Days to maturity			
(F/K)	7 to 60	61 to 120	121 to 180	181 to 360	7 to 60	61 to 120	121 to 180	181 to 360
(0.80 , 0.85]	-4.15	3.14	2.48**	2.22**	1.72	2.34	0.54	1.01
t-stat	(-0.94)	(1.41)	(1.97)	(2.31)	(0.11)	(0.35)	(0.13)	(0.35)
nobs	356	171	210	175	384	329	401	358
(0.85, 0.90]	-2.33	0.61	1.8	2.53***	-6.54	-1.58	-3.79	2.05
t-stat	(-0.49)	(0.28)	(1.44)	(2.76)	(-0.52)	(-0.27)	(-1.04)	(1.06)
nobs	445	222	260	236	499	416	490	447
(0.90, 0.95]	0.6	0.25	2.37*	3.22***	0.86	2.22	-5.08	-0.64
t-stat	(0.10)	(0.12)	(1.91)	(3.59)	(0.09)	(0.55)	(-1.52)	(-0.28)
nobs	549	265	310	295	602	472	511	486
(0.95, 0.98]	3.54	5.73**	4.43***	4.66***	-1.47	2	-0.9	0.24
t-stat	(0.49)	(2.21)	(3.17)	(4.26)	(-0.17)	(0.46)	(-0.39)	(0.11)
nobs	522	263	275	271	621	420	466	415
(0.98, 1.02]	-6.12	5.39**	4.83***	4.46***	-5.39	1.4	-1.71	-0.2
t-stat	(-0.85)	(2.02)	(3.38)	(4.35)	(-0.66)	(0.38)	(-0.64)	(-0.10)
nobs	559	294	327	307	692	448	482	444
(1.02, 1.05]	2.39	5.85*	4.15**	4.4***	-5.98	0.36	-3.24	1.44
t-stat	(0.27)	(1.88)	(2.46)	(3.95)	(-0.69)	(0.09)	(-1.08)	(0.86)
nobs	506	263	311	282	634	351	433	384
(1.05, 1.10]	4.83	6.57*	3.05*	4.24***	-1.67	-1.65	-2.49	0.6
t-stat	(0.47)	(1.91)	(1.84)	(3.89)	(-0.23)	(-0.44)	(-1.02)	(0.35)
nobs	483	321	357	337	663	412	452	442
(1.10, 1.15]	1.49	6.1*	2.1	5.49***	-5.67	-0.22	-1.1	1.92
t-stat	(0.13)	(1.73)	(1.31)	(3.82)	(-0.91)	(-0.06)	(-0.51)	(1.12)
nobs	419	278	345	331	586	359	396	390
(1.15, 1.20]	26.96	12.7***	5.38**	6.35***	3.63	1.97	-5.93*	1.87
t-stat	(1.46)	(2.72)	(2.51)	(4.32)	(0.69)	(0.58)	(-1.78)	(1.06)
nobs	337	226	297	290	494	300	337	330

Table 12. *Regression analysis: Returns from buying long-maturity put options*
We report regression results on monthly excess returns over the risk free rate from buying new listed put options with maturities between 6 months and 1-year. The sample consists of initial public offerings in 1996-2008 that have options listed within 36 months after the issuance date. Returns are computed from a trading strategy that buys put options quoted within 30 days after the first option contract is introduced. We assume that put options are held until their expiration and include transaction costs in the return calculation by assuming that options and the underlyings are traded at their bid and ask prices. Monthly excess return is computed by subtracting the put option return with the corresponding riskfree rate and scaling it with $\frac{30}{\tau}$, where τ is the contract's maturity. Model (1) reports the base line regression result. Models (2) -(6) report the regression results of excess put returns on IPO variables. We include $Moneyness\ bin$ in the regressions to control for differences in excess returns across the contracts' moneyness (see text for full details). $Venture\ backed$ is a dummy variable equal to 1 if the IPO firm is backed by a venture capitalist. $First\ day\ ret$ is computed as the log of return on the offered price to the end of the first trading day. $High\ underwriter\ rep$ is a dummy variable equal to 1 if the firm's lead underwriter is in the highest prestige category (rank=9) according to the ranking published on Jay Ritter's website. $Shares\ locked\ up$ is a dummy variable that is equal to 1 if the IPO firm has an outstanding shares is locked up agreement when option is listed, and zero otherwise. *Time-to-list* is the number of months from IPO issuance date to when option is listed. Year fixed effect are included in the regressions. Robust t-statistic, clustered at the firm-level, are reported in bracket below each estimate. Superscripts ***, **, and * indicate significance at the 1, 5, and 10 percent levels for a two-tailed test.

	Dependent variable: Monthly returns from buying long-maturity put options (%)					
	(1)	(2)	(3)	(4)	(5)	(6)
Intercept	2.286***	-0.070	-0.545	2.172*	2.057**	-1.295***
	(2.69)	(-0.05)	(-0.49)	(1.69)	(2.12)	(-3.88)
Venture backed		3.549**				
		(1.97)				
First day ret			6.819***			
			(4.24)			
High underwriter rep				0.182		
				(0.10)		
Shares locked up					1.044	
					(0.47)	
Time-to-list						-0.220**
						(-2.34)
Moneyness bin	0.427***	0.430***	3.166*	0.425***	0.429***	0.431***
	(2.79)	(2.82)	(1.71)	(2.78)	(2.80)	(2.82)
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Numer of obervations	2524	2524	2524	2524	2524	2524
Number of clusters	504	504	504	504	504	504
Adjusted R^2	0.27%	1.01%	2.71%	0.23%	0.28%	1.19%

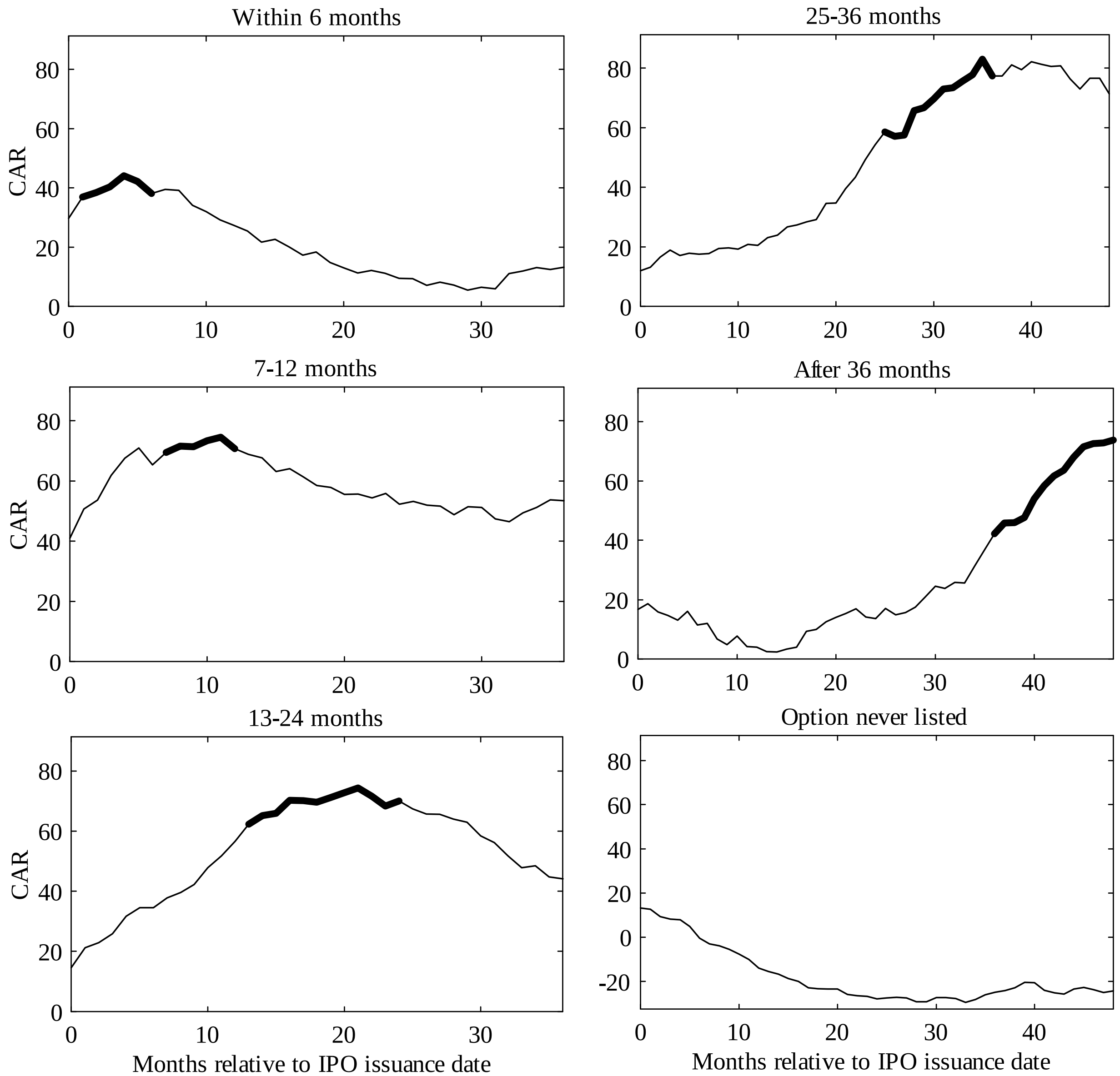

Figure 1. *Option listing and IPO long-run returns.* The sample consists of firms that go public in 1996-2008. We plot the long-run cumulative adjusted returns (CAR) of initial public offerings starting from the event month 0 which represents to the first day return. We plot the results for six IPO groups categorized by when they have options listed after the IPO issuance date. CAR series are computed using matching firm-adjusted returns. Darkened solid line in each panel indicates when options are listed relative to the IPO issuance date. Panels in the left column plot CAR for IPO firms that have options listed within 6 months (top), in 7-12 months (middle), and in 13-24 months (bottom), of the IPO issuance date. Top and middle panels in the right column plot CAR for IPO firms that have options listed in 25-36 months (top), and after 36 months (middle), respectively. The bottom-right panel plots CAR for IPO firms that never have options listed.

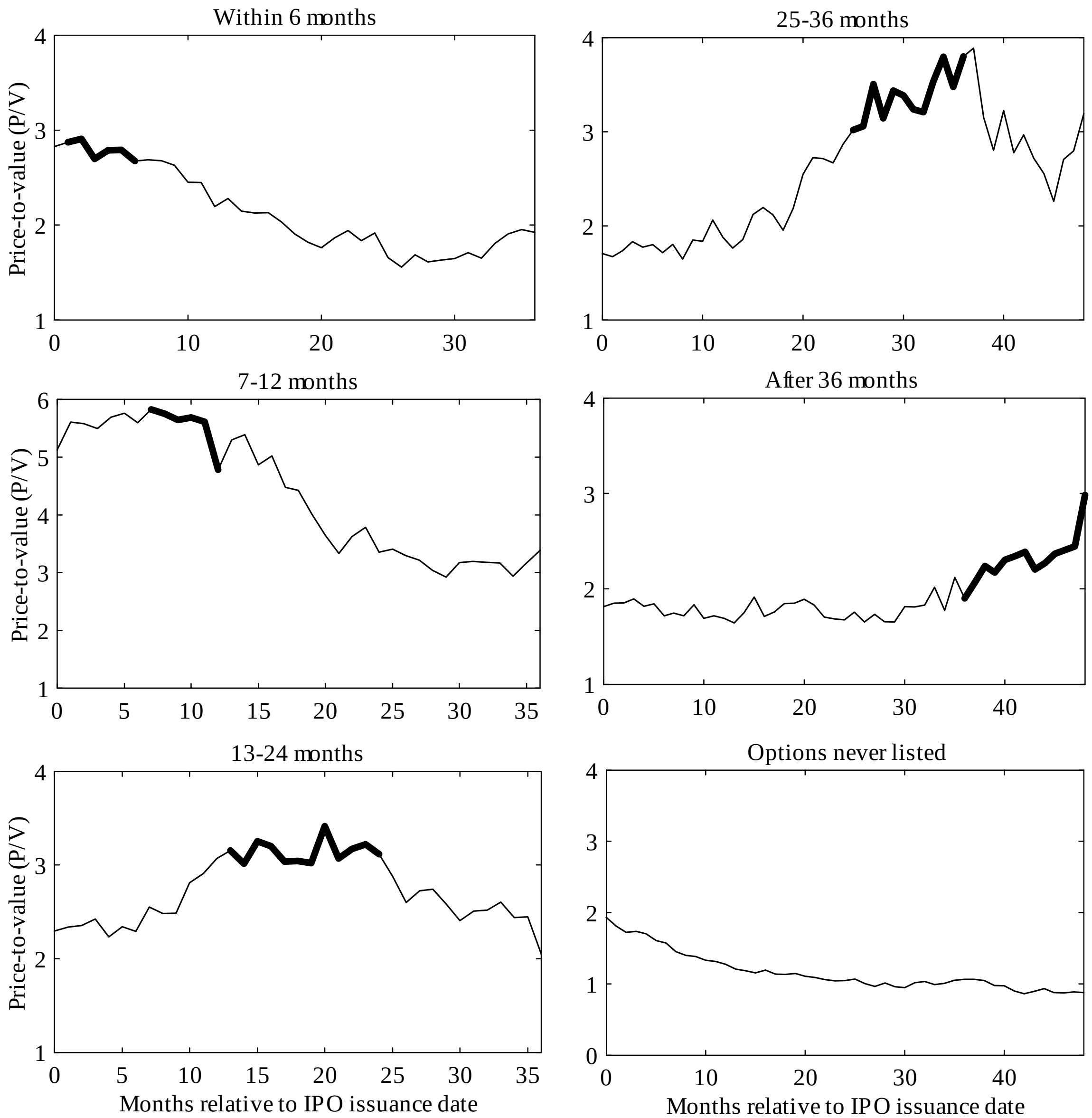

Figure 2. *Option listing and IPO long-run valuation.* The sample consists of firms that go public in 1996-2008. We plot the long-run IPO price-to-value (P/V) ratio starting from the event month 0 which represents to the offer date. The P/V ratio measures the valuation of the IPO equity price against the "fair/intrinsic value" of a comparable firm (see Purnanandam and Swaminathan (2004)). The P/V ratio is computed by dividing the IPO equity price muliple in term of sales, $(P/S)_{IPO}$, by the comparable firm's market multiple in term of sales, $(P/S)_{Match}$. The comparable firm to each IPO is found by matching based on industry, past sales and past EBITDA proffit margin. We plot the monthly median P/V ratios for six IPO groups categorized by when they have options listed after the IPO issuance date. Darkened solid line in each panel indicates when options are listed relative to the IPO issuance date. Panels in the left column plot median P/V ratios for IPO firms that have options listed within 6 months (top), in 7-12 months (middle), and in 13-24 months (bottom), of the IPO issuance date. Top and middle panels in the right column plot median P/V ratios for IPO firms that have options listed in 25-36 months (top), and after 36 months (middle), respectively. The bottom-right panel plots median P/V ratios for IPO firms that never have options listed.

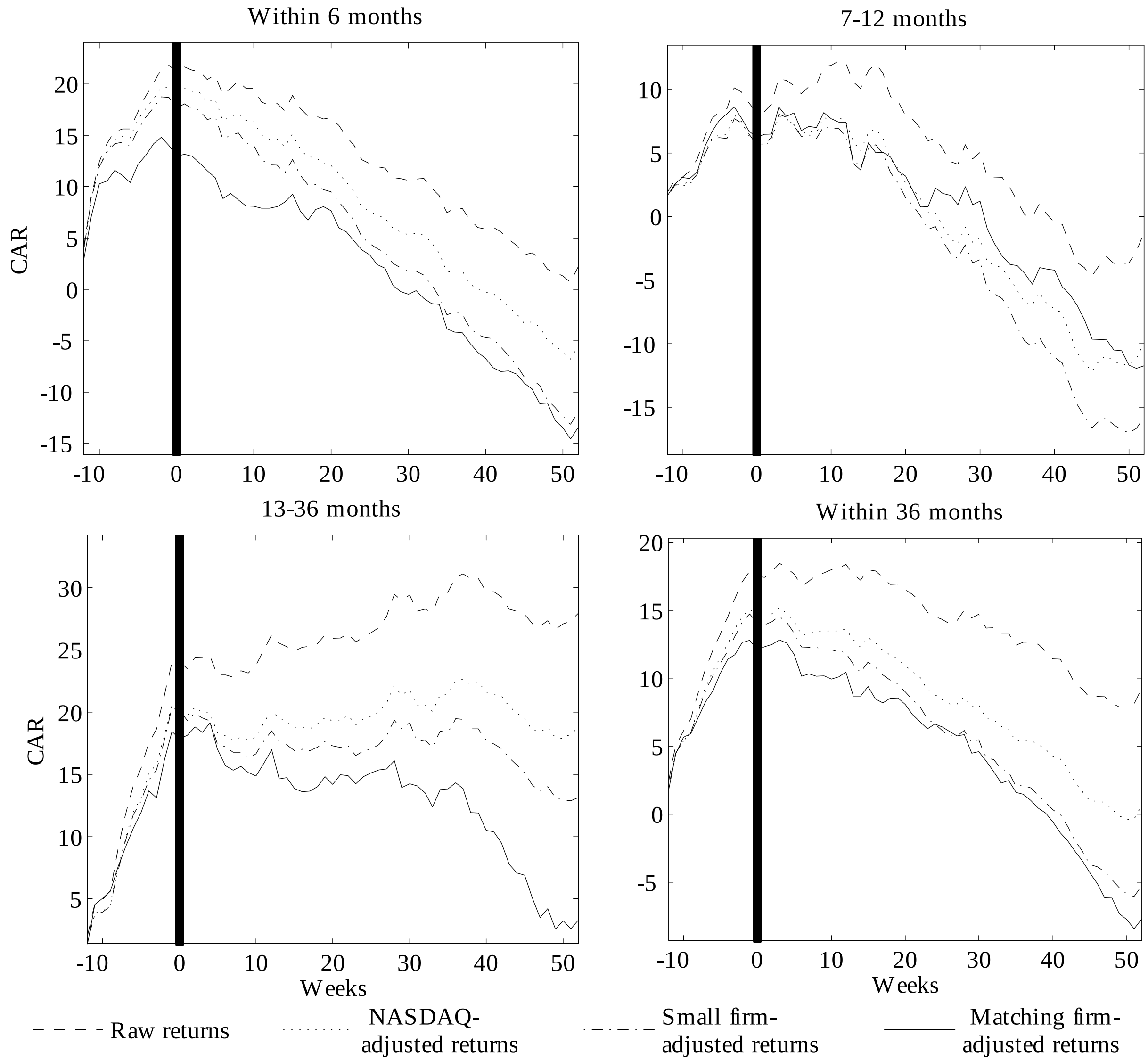

Figure 3. *Weekly cumulative adjusted returns to option listing.* The sample contains initial public offerings in 1996-2008 that have options listed in 0-6 months (Top-left), 7-12 months (Top-right), 13-36 months (Bottom-left), and 0-36 months (Bottom-right), after IPO issuance date. Cumulative average adjusted returns (CAR) are computed weekly centered on the option listing date. Week 0 refers to the option-listing window and is computed from trading day -2 to day +2 relative to the option listing date. In each panel, we plot four CAR series starting 12 weeks before option listing dates. Four CAR series are plotted: 1) raw returns, 2) CRSP-value-weighted NASDAQ index adjustment (NASDAQ-adjusted), lowest decile of NYSE market capitalization index adjustment (small firm-adjusted), and 4) matching firm adjustment (matching firm-adjusted).

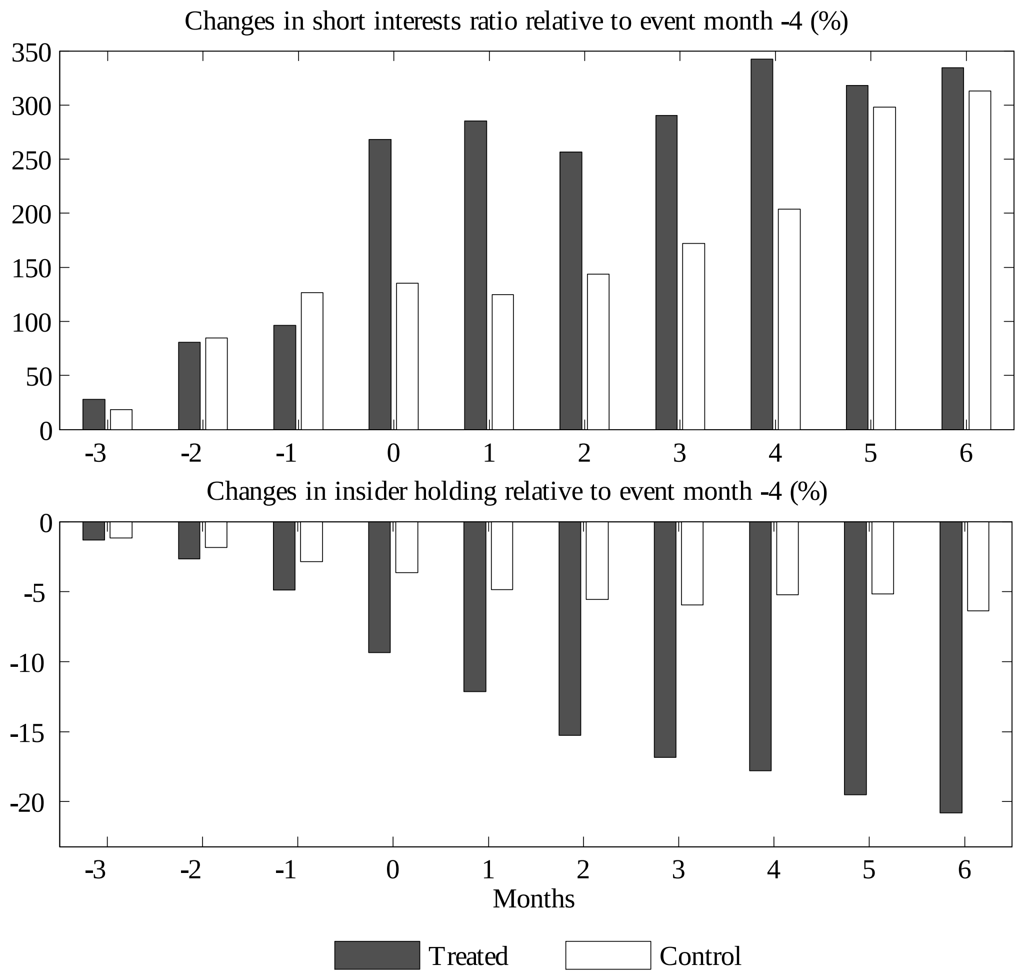

Figure 4. *Relative changes in short interest ratio and insider holding.* The top (bottom) panel plots average monthly changes in short interest ratios (insider holding) relative to event month -4 around option listing date. We plot average values of the winsorized mean (at 1 and 99%) for each event month for the treated and control groups. We obtain short interest levels from COMPUSTAT. Short interest ratio is calculated as the ratio of short interest to the number of shares outstanding. We obtain insider filing data from Thomson Reuter's Insider Filing Data Feed (IFDF). Insider holding is the total direct and indirect holding of conventional stock by the company's insiders with relationship hierarchy level between 1 to 3. The treated group consists of IPO firms in 1996-2008 that have options listed within 36 months after the issuance date. Firms in the treated group are matched to IPO firms that do not have options listed (Control group) based on their propensity score of having options listing. We require that firms in the control group must not have been public for more than four years and will not experience option listing for at least one year after they have been matched. The matching is done at the firm-month level based on IPO characteristics, size, turnover, volatility, and momentum (see Table A1).