

What is the Role of the Options Market? Evidence from Newly Public Companies

December 2025

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Abstract

We examine how options markets affect price discovery for newly public companies. Stock returns decline by approximately 100 basis points following options listing for newly public firms but not seasoned firms. Security-lending data reject the short-sale constraint relaxation hypothesis: both share borrowing costs and quantities increase after listing. Using comprehensive open-close volume data, we show that proprietary traders accumulate negative equity exposure through put purchases that predict subsequent returns. This predictive power concentrates among firms whose IPO underwriters are affiliated with proprietary traders. Market maker hedging explains the increased short-sale constraints. Our findings support information-based trading as the primary channel.

Keywords: Informed trading; Options trading; Signed volume; Short-sale constraint; IPO

JEL classification: G14; G12; G24.

1 Introduction

How do options markets affect price discovery in the underlying equity? This question has motivated decades of research, yet remains contested. Two competing theories offer different answers. The first argues that options relax short-sale constraints, allowing pessimistic investors to express their views and correcting overpricing.¹ The second posits that options provide a venue for informed trading, with private information revealed through derivatives markets.² Understanding which channel dominates has renewed importance given dramatic shifts in options market participation. Recent evidence documents that retail investors now comprise a substantial share of options trading activity, particularly during the pandemic period.³ When retail participation is elevated, the return predictability of options trading volumes substantially diminishes, suggesting the informed trading channel weakens precisely when uninformed order flow dominates. This raises a puzzle: if options serve primarily to relax short-sale constraints, return predictability should persist regardless of who trades. The observed sensitivity to market composition instead suggests the information channel plays a critical role. We test these competing explanations using newly public companies, where information asymmetry and short-sale constraints are most acute.

We test these hypotheses using the IPO aftermarket, where newly public companies face substantial information asymmetry and significant short-sale constraints.⁴ These features make the IPO aftermarket well-suited for identifying which channel dominates.

We document three main findings. First, options listing produces a strong negative price im-

¹Danielsen and Sorescu (2001) and Jarrow (1980), among others, argue that the options market could relax short-sale constraints on the underlying equity, leading to a decrease in the company's share price in a setting with heterogeneous beliefs among investors. Bakshi, Madan, and Panayotov (2010) develop theory where short-sale constraints drive demand for options as protective instruments.

²Easley, O'Hara, and Srinivas (1998) finds that the options market provides an important avenue, in addition to the stock market, through which those with private information can trade.

³For evidence on how market composition affects options informativeness, see Cashman, Harrison, Scheick, and Sheng (2025) and Bryzgalova, Pavlova, and Sikorskaya (2023), who document that retail participation now comprises a substantial share of options trading activity and affects the information content of options prices.

⁴As explained by Geczy, Musto, and Reed (2002), short selling shares of newly public companies is difficult because there is a restricted supply of shares in the immediate aftermarket, lockup agreements, prohibitions on banks in the IPO syndicate from lending shares in the first 30 days, and the relatively small public float. Further, investors in newly listed companies face greater information asymmetry because little information about their performance and prospects is available before these companies go public. See also Boulton, Braga-Alves, and Shastri (2020) for cross-country evidence on how short-sale constraints impede price discovery and liquidity in IPO aftermarkets.

impact on newly public companies— approximately 100 basis points—but not on seasoned firms. Using a novel empirical insight that 15% of options listings are not accompanied by immediate trading, we show that listings without trading produce no price effect, helping to address concerns about exchanges’ timing ability. Second, we reject the short-sale constraint relaxation hypothesis. Security-lending data reveal that both borrowing costs and quantities increase, indicating that constraints intensify rather than relax. Third, we find strong support for the information-based trading channel. Using comprehensive open-close volume data from eight options exchanges covering 60–80% of U.S. equity options trading, we document that proprietary traders accumulate negative equity positions through put purchases that predict subsequent returns. This predictive power is concentrated among firms whose IPO underwriters are likely affiliated with proprietary traders. Market maker hedging explains the increased short-sale constraints as their hedging needs require shorting the underlying stock, driving up borrowing demand.

Our first contribution documents the differential impact of options listing across firm types. Figure 1 illustrates this finding for U.S. companies that went public from 1996 to 2007. Panel A shows that cumulative adjusted stock returns (CAR) decline by 100–150 basis points in the weeks following options listing for newly public companies. Panel B addresses the concern that exchanges time their listing when share prices have peaked (Mayhew and Mihov, 2004). We exploit the fact that options do not always begin trading immediately when listed. About 15% of listings experience trading delays exceeding one week, with median delays of 17 days.⁵ Using these listing-without-trading events as a counterfactual, Panel B shows that options listing without actual trading produces no price effect, suggesting that the negative returns stem from options trading itself rather than exchanges’ timing ability. For seasoned public companies, we find no robust evidence that options listing affects stock returns, consistent with theories predicting stronger effects when information asymmetry and short-sale constraints are severe.

Our second contribution directly tests and rejects the short-sale-constraint relaxation hypoth-

⁵Previous studies use options exchanges’ listing dates as the event of interest, with mixed findings depending on sample periods. Branch and Finnerty (1981) and Conrad (1989) find positive price effects for U.S. options listings in the 1970s, while Sorescu (2000) and Danielsen and Sorescu (2001) report negative abnormal returns after 1980. Our paper differs by examining newly public versus seasoned firms and distinguishing between listing dates and trading start dates.

esis. Following D’Avolio (2002), we examine security-lending market data where investors must borrow shares to establish short positions. If options relax short-sale constraints, borrowing costs should decrease or remain stable while quantities increase. Instead, we find that both the cost to borrow shares and the quantity borrowed increase significantly after options listing, by 15–20% and 35–43%, respectively. This simultaneous increase in price and quantity indicates a positive shock in short-selling demand that outstrips supply (Cohen et al., 2007), demonstrating that short-sale constraints become more binding after options are listed on newly public companies.

Our third contribution establishes the information-based trading channel. Easley, O’Hara, and Srinivas (1998) theoretically characterize conditions under which informed traders employ options alongside stock markets. This occurs when options’ leverage is valuable, equity liquidity is limited, and the fraction of informed traders is substantial. The IPO aftermarket embodies these conditions. Newly public companies face restricted share supply and high information asymmetry. We test this channel using daily open-close option volume data from eight exchanges, the most comprehensive multi-exchange sample in the literature. These data capture directional trading volume (buy or sell) at the contract level and classify transactions by investor type: Customers and Firm proprietary. Customers represent public accounts including retail and institutional investors, while Firm proprietary, or proprietary traders, refers to trades from the proprietary trading accounts of banks, trading firms, or broker-dealers. Market makers absorb the net positions of these two groups.

We calculate *synthetic net open interests* to quantify investors’ equity exposure through options. Specifically, it is the delta-weighted sum of net option positions across all contracts. Synthetic net open interest represents the investor’s net exposure on the underlying stock, in the equivalent number of underlying shares, through their options holdings. We find that proprietary traders accumulate negative equity exposure by purchasing puts and writing calls, while customers take positive exposure through call purchases and put writing. The difference is largest for out-of-the-money puts and persists for several months. Crucially, proprietary traders’ synthetic positions significantly predict subsequent negative stock returns, while customers’ positions show no predictive power. This pattern is consistent with informed trading by proprietary traders whose put purchases signal negative information to the market, paralleling recent evidence by Harren et al. (2025) that

proprietary traders' put positions predict FOMC impacts in S&P 500 options markets.

We provide suggestive evidence linking proprietary traders' information advantage to IPO underwriter affiliations. All U.S. exchange-traded options are cleared by Options Clearing Corporation (OCC) members, approximately 100 of the largest broker-dealers. Trades by OCC members are labeled Firm proprietary. We classify an IPO as OCC-affiliated if its lead underwriter was an OCC member at the time of the offering.⁶ For example, Wells Fargo became an OCC member in 2007, so a 2009 Wells Fargo-underwritten IPO is classified as OCC-affiliated, while a 2006 Wells Fargo IPO is not. We find that proprietary traders' predictive power exists only for OCC-affiliated firms, suggesting their information advantage stems from underwriter connections.

Our final analysis explains why short-sale constraints worsen after options listing. Market makers hold positions opposite to end users for market clearing. After options begin trading, market makers hold net long equity exposure, they are net sellers of puts and buyers of calls, because proprietary traders' net short exposure exceeds customers' net long exposure. Market makers delta-hedge their net long option exposure by shorting the underlying stock (Ni et al., 2020), driving up both cost and demand in the security-lending market. Our regression estimates show that the increase in share borrowing demand after options listing can be almost entirely attributed to market makers' hedging needs, reconciling the seemingly paradoxical finding that options listing worsens short-sale constraints.

The remainder of this paper is organized as follows. Section 2 discusses our contribution relative to existing literature. Section 3 describes the empirical setup and data. Section 4 documents options listing effects on stock returns. Section 5 tests the short-sale constraint and information-based trading channels. Section 6 concludes.

2 Relation to the existing literature and contribution

This paper contributes to three strands of literature on the economic effects of options listing, informed trading in options markets, and price discovery in the IPO aftermarket. We describe our contributions in relation to each strand below.

⁶For details on OCC clearing members, see [OCC member directory](#).

2.1 Effects of options listing on the underlying equity

The empirical literature on options listing effects reveals a striking historical evolution. Early studies documented modest positive abnormal returns of approximately 2% around options introduction (Branch and Finnerty (1981); Conrad (1989)). However, Sorescu (2000) demonstrates that positive returns characterized only the 1973–1980 period, with negative effects emerging thereafter. Mayhew and Mihov (2004) examine options listing determinants on seasoned firms and show that after controlling for exchanges’ timing ability, options listing has no material impact on equity volatility.

We advance this literature on several dimensions. First, we provide the first systematic comparison of options listing effects across newly public and seasoned firms, a distinction that extant theories predict should reveal substantial differences. Danielsen and Sorescu (2001) and Easley et al. (1998) predict that options introduction should exert considerably stronger price effects when information asymmetry and short-sale constraints are most binding, which characterize newly public companies. Our empirical evidence supports these predictions. We document substantial negative returns following options listing for newly public firms but find no robust effects for seasoned firms, consistent with theory.

Second, we exploit a novel empirical observation. Approximately 15% of options listings experience trading delays of one week or longer. Using these listing-without-trading events as a counterfactual for exchanges’ timing ability, we show that without actual options trading, listing produces no significant stock price effect. This addresses concerns raised by Mayhew and Mihov (2004).

Conventional wisdom suggests that options markets alleviate short-sale constraints.⁷ We reject this hypothesis with new evidence. Both borrowing costs and borrowing quantities increase significantly by 15–20% and 35–43%, respectively. This simultaneous increase indicates demand for locating shares to short outstrips supply (Cohen, Diether, and Malloy (2007)), demonstrating that

⁷For instance, Danielsen and Sorescu (2001) show that put option purchases establish synthetic short positions, thereby circumventing short-sale constraints. Bakshi, Madan, and Panayotov (2010) develop theory showing constraints on short-selling create demand for options as protective instruments for pessimistic investors. Johnson and So (2012) argue that options’ leverage explains why informed traders use them for expressing negative views on stocks.

short-sale constraints become more binding after options listing on newly public companies. We provide the first empirical evidence that this increased constraint stems from option market makers' hedging needs. When market makers absorb proprietary traders' net short positions, they must hedge by shorting the underlying stock, driving up borrowing demand. Our finding underscores the importance of considering intermediary behavior when examining options markets' impact on short-sale constraints.⁸

Our results sharpen the findings of [Figlewski and Webb \(1993\)](#), who document that options trading coincides with increased short interest but conclude short-sale constraints relax. Critically, short interest represents only quantity, not cost. We demonstrate that both increase, indicating constraints tighten. For seasoned firms, borrowing costs do not increase after options listing, suggesting our findings are specific to newly public companies.

2.2 Informed trading in the options market

The second dimension of our contribution addresses information-based trading in options markets. We demonstrate that options markets function as information-aggregation venues specifically for newly public firms, where information asymmetries and short-sale constraints are most acute. [Easley, O'Hara, and Srinivas \(1998\)](#) theoretically establish that options attract informed traders when equity liquidity is limited and private information substantial. Our analysis provides the first comprehensive characterization of how different investor types use options in the IPO aftermarket.

Several empirical studies document informed options trading ahead of corporate events.⁹ Our study pursues a complementary research question by examining how informed traders use options when a new trading venue is introduced. We focus on trading patterns following options listing rather than ahead of specific events.

Our analysis extends the foundational frameworks of [Pan and Poteshman \(2006\)](#) and [Ge, Lin, and Pearson \(2016\)](#). While those studies examine return predictability using one or two exchanges, we leverage comprehensive open-close volume data spanning eight options exchanges covering 60–

⁸[Murayev, Pearson, and Pollet \(2022\)](#) and [Ramachandran and Tayal \(2021\)](#) provide supporting evidence on option pricing reflecting short-sale constraint frictions. [Banerjee and Graveline \(2014\)](#) show that derivatives theoretically reduce price distortion when underlying securities are scarce, but acknowledges their model excludes market makers.

⁹See [Augustin et al. \(2019\)](#), [Augustin et al. \(2023\)](#), [Chan et al. \(2015\)](#), and [Cao et al. \(2005\)](#).

80% of total U.S. equity options trading from 2005–2021. Both [Pan and Poteshman \(2006\)](#) and [Ge et al. \(2016\)](#) find that return predictability comes from public customer positions while proprietary traders’ positions show no predictive power. Our conditional analysis of newly public firms, where information asymmetry is most severe, reveals a different pattern.

Proprietary traders, not public customers, drive return predictability for newly public companies. We document that proprietary traders’ options positions robustly predict subsequent stock returns, while customer positions show no predictive power.¹⁰ This conditional informativeness helps address the surprising result in [Pan and Poteshman \(2006\)](#) that proprietary traders’ option volumes were uninformative in their aggregate sample.¹¹ [Harren et al. \(2025\)](#) find that proprietary traders’ morning option trades on FOMC days predict monetary policy shocks, demonstrating predictive advantages in macro-event contexts. Our results suggest that proprietary trader informativeness is contingent on the information environment, with predictive power emerging when traders possess informational advantages through specific channels—underwriter connections in our setting, policy information in theirs.

We provide evidence linking proprietary traders’ information advantage to IPO underwriter affiliations. All U.S. exchange-traded options clear through the Options Clearing Corporation (OCC), with proprietary trades labeled accordingly. We classify an IPO as OCC-affiliated when its lead underwriter held OCC membership at the offering date. Proprietary traders’ return-predictive power concentrates exclusively among OCC-affiliated firms, suggesting their information advantage derives from underwriter connections. In other words, not all trades from proprietary accounts are informed; they predict returns only for newly public companies when they likely possess information advantages through IPO underwriter affiliation.

Recent advances in measuring informed trading validate this mechanism. [Bogousslavsky et al. \(2024\)](#) develop a machine-learning informed trading intensity measure capturing nonlinearities in trading behavior, demonstrating spikes before earnings and corporate events. Additionally, [Augustin et al. \(2023\)](#) document substantial informed options strategies before corporate announce-

¹⁰We infer future return direction using all option volume components, including new positions and position closures across all contract types.

¹¹[Pan and Poteshman \(2006\)](#) noted: “The most surprising result is that the open-buy volume from firm proprietary traders is not informative at all about future stock prices” (p. 902).

ments, with particular emphasis on put positions, precisely the instruments concentrated in proprietary trader positions in our sample.

2.3 Price discovery in the IPO aftermarket

The IPO aftermarket presents an information-opaque environment characterized by pronounced information asymmetries and binding short-sale constraints. The existing literature establishes two stylized facts: IPO shares trade substantially above offer prices in the immediate aftermarket (Ljungqvist, 2007; Ritter, 1991), and this positive performance often reverses over longer horizons. A substantial theoretical literature attributes these patterns to heterogeneous investor beliefs combined with frictions preventing pessimistic investors from shorting.¹²

The role of short-sale constraints in generating mispricing has been extensively documented. Ofek and Richardson (2003) provide vivid evidence from the dot-com era that binding short-sale frictions explain extreme overpricing of Internet stocks. Recent cross-country evidence strengthens this conclusion. Boulton, Braga-Alves, and Shastri (2020) demonstrate that short-sale constraints substantially impede both price discovery and liquidity in IPO aftermarkets, with underpricing significantly higher in countries where short selling is prohibited or restricted. Edwards and Hanley (2010) examine short-selling activity using Regulation SHO pilot data and find constraints, while material, reveal substantial heterogeneity across firms, with constraints tightest for newly public firms having limited public float and restricted lending availability.

Our results suggest that options listing both helps correct short-term equity mispricing in newly public firms and reveals an unexpected mechanism. Contrary to theories predicting short-sale constraint relaxation, it intensifies. The negative returns following options listing reflect not the relaxation of short-sale frictions, but rather the revelation of private information held by informed traders who are likely affiliated with IPO underwriters. This insight suggests that options markets serve as information-aggregation venues rather than substitutes for equity short-selling, consistent with the theoretical framework of Easley, O'Hara, and Srinivas (1998).

¹²See Miller (1977) for informal arguments and Morris (1996) for formal theory.

3 Data and Sample

3.1 Data sources and sample selection

Our analysis integrates multiple comprehensive data sources spanning 1996–2022. We obtain U.S. initial public offering data on ordinary common shares from the Securities Data Company (SDC) New Issues database for January 1996 through December 2017, excluding closed-end funds, real estate investment trusts (REIT), Dutch auction offerings, and issues flagged with “shelf registration” according to SEC Rule 415. Following corrections suggested by Jay Ritter,¹³ we eliminate observations with trading history in CRSP predating SDC records. Our initial sample consists of 3,723 companies across 29 industries.

Security prices and trading information come from the Center for Research in Security Prices (CRSP). Options listing dates and prices are obtained from OptionMetrics, which covers all U.S. exchange-listed equities from January 1996 onward. We identify the first listing date on any exchange as the official options listing date,¹⁴ excluding over-the-counter markets. For companies with multiple share classes, we focus on the first listing across all classes. We exclude new listings resulting from mergers, acquisitions, or temporary trading pauses that would create spurious new listings in OptionMetrics. We define the trading start date using the first date with nonzero open interest in end-of-day option price data.

Daily security-lending market data, which includes borrowing costs (indicative fees) and quantities on loan, are obtained from Markit Securities Finance for May 2002 through October 2015. The indicative fee represents the expected cost for a hedge fund to borrow shares, calculated by Markit from borrowing costs between agent lenders, prime brokers, and hedge funds. Quantity on loan represents total shares borrowed daily. This data captures equilibrium prices and quantities in the security-lending market, consistent with recent informed trading literature.¹⁵

Our novel contribution leverages open-close options volume data from eight options exchanges

¹³<http://bear.warrington.ufl.edu/ritter/SDC%20corrections.pdf>

¹⁴Listings with status code “L—The security has been listed” or “4—The security is new, and this is the first day priced.”

¹⁵Ramachandran and Tayal (2021) and Murayev et al. (2022) provide recent applications of this security-lending data to examine short-selling costs. We refer to their papers for further details on the security lending market.

covering 2005–2022. This data comprises approximately 60–80% of total U.S. equity options trading volume and provides directional end-of-day contract-level trading volume. The exchanges include four operated by the Chicago Board of Options Exchange (CBOE-C1, CBOE-C2, EDGX, BZX) and four by NASDAQ (PHOTO, NOTO, ISE, GEMX). Coverage expands over time as exchanges make data commercially available, beginning with CBOE’s C1 exchange in January 2005.¹⁶ Figure 2 displays each exchange’s percentage of total trading volume by quarter through December 2022.

3.2 Sample characteristics

Of the 3,723 initial public offerings from January 1996 to December 2017, we observe 2,332 options listings (62.6%). Approximately one-third of listings occur within 6 months of IPO, with approximately half within 12 months. The average days-to-list is 654 days (median 287), reflecting a left-skewed distribution with the shortest gap at 8 days.

A critical feature emerges. Approximately 15% of options listings do not experience immediate trading. The *trading-listing gap*—the number of days between listing and trading start—averages 6.37 days but exhibits substantial variation. We classify option listings with trading that starts within 7 calendar days as immediate trading starts (Trading = Listing), treating those longer gaps as delayed trading starts (Trading > Listing). Figure 3 illustrates this distinction, displaying the distribution of option listings across different time windows relative to the IPO date, separately for immediate and delayed trading starts.¹⁷ This variation provides a falsification test: if negative returns stem from exchange timing rather than trading, we should observe price declines even without trading. Section 4.2 shows we do not.

Table 1 summarizes IPO characteristics used as control variables. *Venture backed* indicates venture capital financing (48.3% of sample). *Underwriter rep*, obtained from Loughran and Ritter (2004)’s website, scales lead underwriter reputation from 0 (lowest) to 9 (highest), averaging

¹⁶CBOE refers to its primary options exchange as C1 starting in 2011. CBOE’s C1 data before 2011 correspond to the primary CBOE exchange’s open-close volume.

¹⁷This cutoff reflects that most option-listing dates fall on Thursday or Friday, affecting trading start timing due to weekend proximity. The 7-day window allows exchanges time to build limit-order books and option market makers to develop hedging capability on newly listed contracts. Appendix Table A2 provides descriptive statistics for immediate versus delayed trading starts. For the 15% listings with delayed trading, the average delayed days is 33 days with the mean of 17 days.

7.43. *First day ret*—the simple return from offer price to first trading day end—measures short-run underpricing, averaging 25.17%. *Proceeds amount* is the log of total offer proceeds including over-allotment options. *Total shares offered* measures shares issued in thousands. *Lock up agreement* indicates presence of lockup provisions (78% of sample), with average 198-day expiration. These characteristics align with established IPO literature patterns.

For security-lending analysis, we retain observations with non-missing Markit data, yielding 616–626 option-listing events over 2002–2015. For open-close volume analysis, expanding data availability across exchanges (shown in Figure 2) provides sufficient coverage by 2010 onward, with sample sizes of 384–523 firms across our 24-week event windows.

4 Options listing and the underlying equity

4.1 Empirical design

We examine the stock return impact of options listing using the following regression model:

$$CAR_{i,t} = \alpha_i + \theta_t + \beta \cdot Optionstat_{i,t} + \sum_{j=1} \delta_{ij} Weekly\ Controls_{ij,t-1} + \varepsilon_{i,t}. \quad (1)$$

The dependent variable is the cumulative adjusted return (CAR) for firm i in week t . Event week $t = 0$ refers to the week options are listed. $Optionstat_{i,t}$ is a dummy variable equal to 1 for $t \geq 0$ (post-listing) and 0 otherwise; β measures the return differential between the post- and pre-listing periods.

We include firm fixed effects α_i and time fixed effects θ_t . Time-varying controls variables represented by $Weekly\ Controls_{ij,t-1}$ are lagged by one week and calculated over a rolling 12-week window. They include stock return momentum (raw cumulative return), volatility (standard deviation of daily returns), average shares turnover (volume/shares outstanding), and market capitalization (log of daily average market cap).

Cumulative adjusted returns are calculated as $CAR_{i,t} = \sum_j (r_i(t, j) - r_m(t, j))$, where $r_i(t, j)$ is the daily return of company i on day j and $r_m(t, j)$ is the daily benchmark return. We employ three benchmarks: (1) CRSP value-weighted return, (2) NASDAQ index return, and (3) lowest-decile

NYSE portfolio by market capitalization.

We define $t = 0$ as the calendar week (Monday–Friday) in which options are listed. This approach mitigates day-of-the-week effects and facilitates replication. Our results remain robust when using the actual listing date instead of the calendar week.

Options listing decisions are not exogenous. Exchanges may time listings strategically. We address selection bias in two ways. First, we exploit listing events not accompanied by immediate option trading as a counterfactual. If observed negative returns reflect pure exchange timing rather than trading effects, they should persist absent trading. Second, we employ propensity-score matching on observable factors to mitigate selection bias. Specifically, we follow the method of Mayhew and Mihov 2004 with added IPO characteristics in the matching.

4.2 Options listing and the stock returns of newly public companies

Figure 1 displays average weekly CAR around options listing. Panel A (immediate trading start: Trading = Listing) reveals that newly public firms experience positive CAR before event week 0, followed by an immediate and persistent decline. This pattern is robust across benchmarks. Panel B (delayed trading start: Trading > Listing) shows no comparable decline. This suggests that that negative returns reflect trading activity, not exchange timing of when to list options.

Table 2 reports OLS regression results. We estimate the model in equation (1) using two observation windows. The first is a balanced window from event weeks -12 to $+12$. The second observation window is from event weeks -12 to $+24$. Panel A (Trading = Listing) shows that β is negative and significant. In the $[-12, +12]$ window, $\beta = -0.959\%$ (t-stat = -6.00), indicating weekly returns decline by 96 basis points post-listing. In the $[-12, +24]$ window, $\beta = -1.304\%$ (t-stat = -8.73), demonstrating substantial persistence. Panel B (Trading > Listing) reports much smaller and statistically insignificant coefficients, confirming that absent of options trading, listing has no material impact on returns.

To further address exchange timing concerns, we estimate a difference-in-difference (DID) regression on propensity-score matched samples. Appendix Section A1 explains how we construct the matched sample and shows that the control and treated groups are well matched on observable

characteristics. The DID regression is as follows:

$$CAR_{i,t} = \alpha_i + \theta_t + \beta_1 Treated_i + \beta_2 Optionstat_{i,t} + \beta_3 (Optionstat_{i,t} \times Treated_i) + \sum_{k=1} \gamma_{ik} IPO\ variable_{ik} + \sum_{j=1} \delta_{ij} Weekly\ Controls_{ij,t-1} + \varepsilon_{it}. \quad (2)$$

Here, $Treated_i$ equals 1 for listed firms (treatment group) and 0 otherwise. The coefficient on $Optionstat \times Treated$ is the DID estimator of interest. All matching variables are included as controls; they include IPO characteristics and weekly time-varying control variables. We include industry fixed effects, represented by α_i based on the 3-digit SIC code.

Table 3, Panel A reports DID results (Trading = Listing). The interaction coefficient is negative and significant, though slightly smaller in magnitude than OLS estimates, yet economically substantial. The $Treated$ coefficient is insignificant, indicating treated and control firms had similar pre-listing CAR performance, which is a desirable outcome. Panel B (Trading > Listing) shows positive, insignificant interactions, further confirming results.

In summary, our findings suggest that options listing on newly public companies generates a negative, and persistent price impact driven by actual trading rather than exchange timing decisions.

4.3 Options listing and seasoned public companies

We next examine whether the negative return effect of options listing extends beyond newly public companies to seasoned firms. We define seasoned companies as those that went public in 1996–2017 but lacked options for at least 36 months following their IPO. This definition ensures that seasoned firms have established public trading histories and less acute information frictions compared to their newly public counterparts. Our seasoned-firm sample comprises 406 options listing events.

Appendix Table A4 reports OLS regression results from equation (1). For immediate-trading listings (Trading = Listing), the $Optionstat$ coefficient is not significantly negative in the $[-12, +12]$ window. For delayed-trading events (Trading > Listing), coefficients are positive and statistically

insignificant. This pattern contrasts sharply with the robust negative returns documented for newly public companies.

Appendix Table A5 presents corresponding difference-in-differences results using the propensity-score matched sample. The interaction term $Optionstat \times Treated$ never achieves negative statistical significance, and for delayed-trading listings, it turns positive. These findings provide additional confirmation that options listing lacks a robust negative price effect on seasoned firms.

This differential impact across newly public versus seasoned companies aligns squarely with theoretical predictions. Danielsen and Sorescu (2001) and Easley, O'Hara, and Srinivas (1998) both predict substantially stronger options listing effects when information asymmetry and short-sale constraints are most binding—precisely the conditions that characterize newly public, rather than seasoned, companies. Appendix Table A7 examines robustness across IPO characteristics.

For all subsequent analyses, we focus exclusively on immediate-trading-start listings to isolate the economic impact of options market introduction itself, reserving seasoned-firm evidence primarily for robustness and specification checks.

4.4 Options listing and the short-sale constraint

The introduction of the options market is commonly believed to mitigate short-sale constraints on the underlying stock. Therefore, the relaxation of the short-sale constraint associated with options listing could explain the negative stock returns on newly public companies. This section tests this conjecture. Danielsen and Sorescu (2001) and Banerjee and Graveline (2014) provide two different, although related, explanations for the negative stock returns when the options market is introduced. Danielsen and Sorescu (2001) argue that when the options market is introduced, investors can establish synthetic short positions by purchasing put options, and this de-facto removal of short-sale constraints drives down the price of the underlying stock. In this case, market makers who sell put options would carry the responsibility of shorting the underlying share in order to hedge their exposure. But they likely do so at a much lower cost than regular investors.¹⁸ In

¹⁸Prior to 2008, option market makers were exempted from the requirement to borrow shares before shorting. Evans, Geczy, Musto, and Reed (2009) find that half of the time, option market makers choose not to borrow “special” stocks before short selling and instead fail to deliver them to their buyers.

the model of Banerjee and Graveline (2014), the introduction of redundant derivatives can reduce price distortion on an underlying security that is scarce by serving as substitutes for its long and short positions. In this case, options listing would lower the price associated with short-sale constraints by relaxing the scarcity of its underlying.

Shorting a stock requires that the short seller first borrow the share. As described in D'Avolio (2002), for an investor to short a stock, she must find an existing owner of the stock who is willing and able to lend her shares. Having negotiated a loan, she may short sell the borrowed share to any willing buyers. The cost to borrow shares (or loan fee) is the common constraint associated with short selling, and this fee is a function of supply and demand in the security-lending market.¹⁹

The explanations of Danielsen and Sorescu (2001) and Banerjee and Graveline (2014) share a key testable implication: the introduction of the options market should lower frictions associated with the cost to borrow (loan fee) and the difficulty of locating shares to borrow (scarcity). We test this hypothesis directly using data from Markit Securities Finance. We employ two variables that characterize the shorting demand and shorting supply of shares before and after options listing. The first variable is *Indicative fee*, representing the expected cost to borrow a share that is required before shorting. This value is calculated daily by the data provider and indicates the standard market cost to borrow. The second variable is *Quantity on loan*, representing the total number of shares being borrowed each day. *Indicative fee* and *Quantity on loan* represent the equilibrium price and quantity of the underlying share in the lending market. We winsorize their values at the 1st and 99th percentile to mitigate outliers.

It is important to examine how the loan fee and the quantity of shares on loan move together. When both the fee and the quantity on loan increase, it signals that the short-sale constraint worsens and that shorting demand is the dominant force. This price-and-quantity increase indicates an outward shift in shorting demand that outpaces the supply of shares available for shorting, and thus, the stock becomes more scarce. Conversely, when the loan quantity increases but without an increase in fee, it suggests that the supply of shares sufficiently shifts outward to keep up with the demand for shorting. In this latter case, the short-sale constraint is unaffected.

¹⁹Other short-sale constraints include regulations and the risk of borrowed shares being recalled by their lenders.

We test the effect of options listing on short-sale constraints using a regression model similar to equation (1). We replace the dependent variable CAR by weekly average *Indicative fee* and *Quantity on loan*, which is calculated by simply averaging their daily values over each event week. Table 4 reports the results. As discussed earlier, we consider only options listings that are accompanied by immediate option-trading start for the remainder of this paper. Firm and year fixed effects are included in all regression specifications.

We find that *Indicative fee* and *Quantity on loan* increase significantly after options listing. Regarding economic magnitude, Columns (1)-(2) of Table 4 show that the coefficient estimate on *Optionstat* suggests the indicative fee increases by 25–36 bps per annum. The average indicative fee in the 12 weeks before options listing is 1.74 %, indicating an increase of about 14–21% in relative magnitude after options listing. The increase in the quantity of shares on loan is between 329,000 to 428,000 units. We can compare this amount to the average quantity on loan of 0.93 million units in the 12 weeks before options listing, indicating an increase of 35% or greater. Our findings in Table 4 demonstrate that short-sale constraints on the equity of newly public companies are more binding after options listing and that this reflects an outward shift in shorting demand that outstrips the supply of shares available for borrowing.

Appendix Table A6 reports our findings for seasoned public companies. Although we find that the quantity of shares on loan significantly increases after options listing, the cost to borrow shares does not. The increase in the number of shares on loan is consistent with Figlewski and Webb (1993), who find that the short interest on a company increases significantly after the introduction of the options market. However, we emphasize that the level of short interest represents only the quantity and not the cost associated with short selling.

5 Informed options trading on newly public companies

This section explains how information-based trading reconciles the documented negative stock price impact and increased short-sale constraints following options listing. Our empirical analysis employs the most comprehensive open-close volume data of options trading activity known to date, covering 60–80% of U.S. equity options trading from 2005–2022. This comprehensive data cov-

erage enables examination of how different investor types trade options on the underlying equity and allows us to reconcile the observed increase in short-sale constraints after options listing.

5.1 Open-close options volume

Open-close volume data provides daily opening and closing positions on each option contract, including trade direction (buy or sell) and investor type. The Options Clearing Corporation (OCC) classifies all transactions by origin code: F for firm proprietary traders, C for public customers, and M for market makers.

Firm proprietary, or proprietary traders, refer to proprietary accounts of banks, broker-dealers, or trading firms—all clearing members of the OCC. These include, for example, employees at Barclays Capital LLC who trade on behalf of the bank's own account. The names of clearing members are published in the OCC annual report, with reports from 2014 onward available from the OCC website and earlier reports obtained from the OCC archives. Public customers range from retail and professional traders to institutional investors, with exchanges further identifying trades from public customer accounts according to their sizes. Market makers trade in opposite directions to absorb net positions of both proprietary traders and public customers, satisfying the zero net supply constraint inherent in options markets.

We merge contract-level open-close volume data from eight options exchanges into the OptionMetrics database. These eight exchanges comprise CBOE-C1, CBOE-C2, EDGX, and BZX (operated by the Chicago Board of Options Exchange) and PHOTO, NOTO, ISE, and GEMX (operated by NASDAQ). Our data covers 60–80% of total U.S. equity options trading volume from 2005–2022, with coverage expanding over time as exchanges made data commercially available beginning with CBOE's C1 exchange in January 2005. This represents the most comprehensive multi-exchange open-close volume dataset used to date in the options literature.

The OptionMetrics database provides complete end-of-day information on all U.S. exchange-traded options across all options exchanges. In contrast, open-close volume data from individual exchanges exist only when contracts have non-zero trading volume on that platform. We merge these datasets at the contract level, excluding contracts that cannot be matched into OptionMetrics.

We examine trading activity across three investor groups: Firm proprietary, Customer, and Market maker.

Open-close volume data reports four trade categories: (1) *Open-buy* represents contracts purchased to open new positions; (2) *Open-sell* represents contracts sold to open positions; (3) *Close-buy* represents contracts purchased to close existing positions; (4) *Close-sell* represents contracts sold to close existing positions.

Daily net open interest for option series j on day t is calculated as:

$$Net\ open\ int(j, t) = \sum_q^t (Open-buy_{j,q} + Close-buy_{j,q} - Open-sell_{j,q} - Close-sell_{j,q}). \quad (3)$$

This calculation accumulates daily order imbalance (buy minus sell) from the inception of each option series through date t . Since our data covers 60–80% of equity options volume, this approach introduces approximation error. However, [Ni, Pearson, Potesman, and White \(2020\)](#) provide evidence that errors from this approximation procedure are unlikely to bias results. Notably, their evidence is based on data from only two exchanges (CBOE-C1 and ISE) covering 20–40% of all equity options volume. Therefore, any errors induced by this approximation are likely to be considerably smaller given our more comprehensive data coverage of 60–80% from eight exchanges. Each option contract exists in zero-net supply, so we estimate market maker net open interests as the negative value of the combined net positions of proprietary traders and public customers.

5.2 Synthetic net open interest

To quantify each investor’s net equity exposure through multiple option series, we calculate the *synthetic net open interest* (SNOI):

$$Synthetic\ net\ open\ int(t) = 100 \cdot \sum_{j=0} \Delta(j, t) \cdot Net\ open\ int(j, t), \quad (4)$$

where the summation applies across different option series and $\Delta(j, t)$ is the delta of option series j on day t . Since each option contract represents 100 underlying shares, multiplying by 100 expresses synthetic net open interest in equivalent share units.

Using option delta as weights in equation (4) serves two critical purposes. First, delta combines an investor’s positions in puts and calls into a single net equity exposure measure. Option delta ranges from -1 to 1 , with sign indicating the direction of positive or negative exposure on the underlying share. Delta is always positive for calls and negative for puts. Second, using option delta as weights allows us to combine holdings across moneyness and maturity. Under the Black-Scholes framework, option delta represents the fraction of shares required to hedge an option price against changes in the underlying price. Equivalently, delta magnitude indicates the probability the option expires in-the-money (ITM). Out-of-the-money (OTM) options have much smaller deltas than at-the-money (ATM) options because ATM options are more likely to be exercised and thus require a larger fraction of underlying shares to hedge. Under this interpretation, option delta represents the level of bullishness (or bearishness) that the holder places on the underlying stock.

OptionMetrics frequently omits delta values for deep ITM and deep OTM contracts. We therefore calculate option delta using the interest rate and dividend-yield information from OptionMetrics. We exclude options with less than one day to expiration since their delta cannot be precisely inferred using end-of-day information. We confirm in robustness checks that our conclusions are qualitatively unaffected when using the delta values reported in OptionMetrics.

SNOI quantifies the net exposure (or view) that each investor holds on the underlying stock through option contracts. A positive value indicates a bullish position, while a negative value indicates a bearish position. Table 5 provides descriptions of weekly SNOI over the 24 weeks after options are listed on newly public companies. Weekly SNOI is the average of daily values over each week. We report results for the three investor groups: *Firm proprietary*, *Customer*, and *Market maker*. The SNOI of *Market maker* equals the negative sum of *Customer* and *Firm proprietary* SNOI. Values are measured in equivalent 100,000 shares. Columns (1), (2), and (3) report results based on put holdings, call holdings, and combined put and call holdings, respectively.

We find striking differences in how proprietary traders and public customers trade options. Proprietary traders, on average, hold a negative synthetic position (short) on the underlying equity by purchasing puts and writing calls. Customers, on the other hand, hold a positive synthetic position (long) by purchasing calls and writing puts. The difference between their SNOI is largest for put

options, suggesting that customers hold a bullish view on the underlying share while proprietary traders hold a bearish view. On average, after options begin trading, proprietary traders hold a daily synthetic short position in put and call options equivalent to approximately 19,900 shares (Column 3), while public customers hold a synthetic long position equivalent to approximately 16,500 shares. The range of these SNOI values are large but clearly left-skewed for proprietary traders.

We test whether SNOI differences between customers and proprietary traders are statistically significant in Table 6. We estimate an OLS regression on a weekly panel over the 24 weeks after options are listed. The independent variable of interest is *Firm proprietary*=1, which equals 1 if the SNOI corresponds to proprietary traders and 0 for customers. We include firm fixed effects and event-week fixed effects, allowing us to estimate the difference in SNOI between customers and proprietary traders within the same company during the same week. The coefficient estimates on *Firm proprietary*=1 are negative, though they are statistically significant only in Columns (1) and (3), indicating that the difference in SNOI originates from holdings of put options. Examining economic magnitudes, Column (3) shows that customers hold a daily synthetic long position equivalent to 16,500 shares, while proprietary traders hold a synthetic short position equivalent to $16,500 - 36,400 = 19,900$ shares. These estimates are consistent with the summary statistics in Table 5.

Column (3) shows that market makers, on average, hold a net synthetic long position in puts and calls. This occurs because the magnitude of proprietary traders' synthetic short position exceeds the magnitude of customers' synthetic long position. Therefore, to clear the market, market makers must hold a net synthetic long-option position.

5.3 Synthetic net open interests and future stock returns

A key implication of the model in Easley, O'Hara, and Srinivas (1998) is that if options serve as a venue for information-based trading, the private information component in option trades should predict future stock returns. We test this hypothesis by estimating a panel regression at the firm-week level over the weeks following options listing on newly public companies. Table 7 reports

the results, where the dependent variable is cumulative adjusted returns (CAR) for the subsequent week and the independent variable of interest is weekly SNOI for different investor types. We include firm and year fixed effects as well as weekly control variables. Columns (1)-(2) report results for SNOI calculated from put options only, while Columns (3)-(4) for results for call options only. Columns (5)-(6) report results for SNOI calculated using put and call options.

Our results reveal three distinct patterns. First, proprietary traders' put option positions strongly predict subsequent returns. Second, customers' option trades show no return predictability. Third, market makers' positions reflect passive liquidity provision rather than informed trading. We detail each finding below.

Proprietary traders' put positions exhibit strong predictive power for future returns. The coefficient on *Firm proprietary* in Panel A is positive and highly significant (t-statistic = 3.42), indicating that proprietary traders' put option positions strongly predict subsequent stock returns in the expected direction. A more negative SNOI—reflecting greater put purchases—predicts lower stock returns, while a more positive SNOI—reflecting put sales—predicts higher returns. Economically, when proprietary traders buy put options equivalent to 100,000 underlying shares, the company's stock price decreases by approximately 15 basis points over the next week. The coefficient on *Firm proprietary* is insignificant for call-only SNOI (Panel B), but becomes positive and significant when combining puts and calls (Panel C). This pattern indicates that proprietary traders' return predictability originates exclusively from their put option positions.

In contrast, customers' option trades contain no predictive information. We find that customers' put-only SNOI (Column 1) is negative and weakly significant, with a direction opposite to proprietary traders. This counterintuitive sign is inconsistent with informed trading. When examining combined put and call positions, customers' SNOI shows no significant relationship with future returns. Collectively, these results provide no evidence that customers' option trades are informed.

Market makers' positions reflect their role as liquidity providers rather than informed traders. Panels showing market maker SNOI (Columns 2, 4, and 6) reveal negative and insignificant coefficient estimates, indicating market makers' positions do not predict returns. This finding is unsurprising since market makers function primarily as liquidity providers who absorb net positions of

end users rather than trading on information.

The negative but insignificant coefficients on market maker SNOI warrant explanation. Market makers hold positions opposite to the aggregate of proprietary traders and customers; thus, their SNOI reflects the combined influence of both informed proprietary trading and uninformed customer trading. In an untabled result, we find that customer trading volume exceeds proprietary trader volume by approximately threefold. Consequently, uninformed customer order flow dominates market maker positions in aggregate, which could crowd out or obscure any private signals embedded in proprietary traders' option trades.

5.4 *Why do options trades of proprietary traders contain new information?*

Our results in Table 7 show that proprietary traders' put option purchases predict negative stock returns, revealing private information about newly public companies. We investigate whether this information derives from underwriter affiliations.

The IPO literature has consistently documented the critical roles of the investment bank that helps the company prepare for its IPO process, i.e., the company's lead underwriter. This includes conducting due diligence on the company, setting the offering price, advising the company on market conditions, marketing and distributing the company's shares, and facilitating trading in the company's common stock after the IPO. The lead underwriter therefore possesses substantially more information about the newly public company's performance and financial health than other investors. This informational advantage can reach the proprietary trading desk within the same investment bank. Affiliated proprietary traders may use this access in their options trading, with put purchases signaling private information about the company. When other investors observe and decode these trading patterns, stock prices adjust downward.²⁰

We test the hypothesis that proprietary traders' informational advantage stems from their connection with the company's lead IPO underwriter. We identify IPO underwriters that were OCC clearing members at the time of the offering. The OCC annually publishes its clearing members; companies with underwriters in this registry are classified as OCC-affiliated. For example,

²⁰Several options exchanges make their open-close volume data commercially available at the end of each business day, allowing the public to infer the direction of trades originating from different investor groups.

Wells Fargo became an OCC member in 2007, so 2009 IPOs underwritten by Wells Fargo are OCC-affiliated, whereas 2006 offerings are not. We find that proprietary traders' option trades predict returns exclusively for OCC-affiliated firms. Table 8 reports this finding.

Panel A reports results for the OCC-affiliated companies. Column (1) shows that the coefficient on *Firm proprietary* is positive and significant, with magnitude and statistical significance substantially exceeding the baseline estimate in Table 7. The predictability remains specific to put options. Panel B reports results for non-OCC-affiliated companies, where proprietary traders' positions show no predictive power; Column (4) shows a negative coefficient, inconsistent with information-based trading. This sharp bifurcation suggests that underwriter affiliation confers the information advantage driving return predictability.

Our OCC-affiliation classification leverages aggregate proprietary trader volumes at the exchange level, a measurement approach inherent to our data structure. This is because open-close option volume data labels all option trades that originated from any proprietary traders account as one entity. However, the empirical evidence suggests this measurement approach would work against our finding. Specifically, if substantial return predictability arose from non-affiliated proprietary traders, we would observe significant predictive power in Panel B (non-OCC-affiliated firms). Instead, the predictive signal entirely disappears for these firms, suggesting that informed trading activity concentrates among OCC-affiliated proprietary traders.

A related consideration concerns the composition of firms in our OCC-affiliated sample. Most OCC clearing members are affiliated with major investment banks that serve as lead underwriters for large-capitalization companies. Thus, our analysis primarily captures informed trading dynamics among larger, higher-profile firms. However, if this argument is true, our results in Table 8 are likely to underestimate the importance of informed option trading by affiliated proprietary traders. Large-cap companies generally have better analyst coverage, better media following, and more liquid shares traded shares. As a result, their equity is more efficiently priced, and any informed option trades in these large-cap companies are likely to produce a smaller impact on their equity value relative to a smaller and less well-recognized company.

Our results are consistent with two mechanisms through which underwriter information reaches

equity markets. First, information may flow from underwriters to affiliated proprietary traders within the same investment bank. In this case, put purchases by informed traders would signal negative information. Second, underwriters may hedge substantial post-IPO equity stakes through proprietary trading desk, and such purchases would reflect rational hedging of downside risks. Both mechanisms operate through affiliated proprietary traders and produce the observed pattern. The data cannot distinguish between these channels, but both imply that options markets effectively aggregate underwriter information about IPO firm prospects.

5.5 Synthetic net open interests and short-sale constraints

An important finding in Section 4.4 is that short-sale constraints worsen after options begin trading on newly public companies. This section explains this seemingly puzzling result through options market trading activity.

As shown in Table 4, both the cost to borrow shares and the quantity of shares on loan increase significantly after options listing, indicating that increased shorting demand outstrips available supply. We provide evidence that this increased demand stems from market maker hedging requirements. Market makers must maintain positions opposite to end users to clear the market. As documented in Table 5, market makers hold net synthetic long positions—selling puts or buying calls on average—to offset end users’ aggregate positions. To hedge these long option positions against underlying price changes, market makers short the underlying stock, thereby driving up security-lending demand.

To illustrate this mechanism, consider a market maker holding w_j in option contract $O_j(S_t)$, where j denotes the contract and S_t denotes the current underlying stock price. To hedge this position, the market maker holds d_j shares of the underlying. The portfolio value is $w_j \cdot O_j(S_t) + d_j \cdot S_t$. Taking the first derivative with respect to S_t and setting it to zero, we solve for the optimal delta-hedge ratio $d_j = -w_j \cdot \Delta_j$, where $\Delta_j \equiv \frac{\partial O_j(S_t)}{\partial S_t}$ is the option delta.²¹ Call option deltas are always positive, while put option deltas are always negative. When a market maker sells put

²¹Under the Black-Scholes framework, option delta represents the fraction of shares required to hedge an option position against price changes. Equivalently, delta magnitude indicates the probability the option expires in-the-money. Call option deltas range from 0 to 1, while put option deltas range from -1 to 0.

options ($w < 0$, $\Delta < 0$), the optimal stock holding is $d < 0$, implying a short position in the underlying. Similarly, when buying call options ($w > 0$, $\Delta > 0$), the market maker's optimal stock holding is again to short the stock $d < 0$.

For multiple option contracts j 's, the number of shares that the market maker would hold is $-100 \cdot \sum_j \Delta_j \cdot w_j$, where the multiplier 100 is because each option contract provides exposure to 100 underlying shares. This expression is essentially the negative of the synthetic net open interest (SNOI) shown in Equation (4), with w_j as net open interest on contract j . Thus, the market makers' SNOI directly determines their required short selling quantity. The more synthetic long positions market makers hold, the greater their short-selling demand. This is our prediction. Market maker SNOI should be positively related to short-selling demand and cost.

Table 9 presents regression results examining the relationship between SNOI and short-selling costs and quantities over weeks $[0, +24]$ relative to options listing. Panel A reports results where the dependent variable is indicative fees, while Panel B reports results where the dependent variable is quantity on loan. The columns in both panels are organized identically: Columns (1), (3), and (5) report results for the SNOI calculated using option positions of proprietary traders and customers, while Columns (2), (4), and (6) report results for market makers.

We first examine market makers' hedging effects, which is our prediction. Column (6) in both Panels A and B shows that market makers' combined SNOI from puts and calls has positive and highly significant effects on both indicative fees and quantity on loan. A larger market maker SNOI, reflecting greater hedging needs, leads to substantially higher share borrowing costs and quantities. Examining puts and calls separately, Columns (2) and (4) show that market makers' hedging demand from put options drives the increased short-selling demand and cost. Notably, the coefficient on quantity on loan is approximately one, suggesting that the increase in short-selling demand can be mostly attributed to market makers' hedging requirements.

We next examine proprietary traders' and customers' effects, whose net positions are absorbed by market makers. Column (5) reports results for combined puts and calls positions. We find negative and statistically significant effects for proprietary traders but not for customers. Column (1) shows this relationship is specific to put positions, not call positions (Column 3). These results sug-

gest that increased market makers' hedging needs, which drive up short-selling demand, respond to put purchases by proprietary traders.

Increased shorting demand may also stem from proprietary traders' put purchases, which signal negative private information. When proprietary traders purchase more put options, corresponding to increasingly negative SNOI, this sends a bearish signal to the market. Observing these informed trades, other investors increase their shorting activity, raising both borrowing costs and quantities. Thus, the negative relationship between proprietary trader put-option SNOI and short-selling demand may also operate through the information channel.

In aggregate, results from Table 9 demonstrate that increased short-selling demand following options listing aligns with observed options trading patterns. Market makers' hedging activities (Column 6) drive borrowing demand proportionally to their net long option exposure. Additionally, proprietary traders' put purchases (Column 1), which could signal negative information, induce increased short-selling demand by market makers who supply these puts and hedge by shorting stock. These mechanisms are complementary rather than mutually exclusive. Order flow absorbed by market makers originates partly from informed traders, so observed market maker hedging demand inherently incorporates information-driven components. The identity of market participants ultimately responsible for increased borrowing costs and quantities remains ambiguous. Nevertheless, increased short-sale constraints after options listing prove consistent with typical options market trading patterns in newly public firms and reflect rational hedging and price discovery mechanisms rather than a puzzling market outcome.

6 Conclusion

We provide extensive evidence that options markets affect newly public firms through the information channel rather than the constraint relaxation channel, consistent with the theory of Easley, O'Hara, and Srinivas (1998). The negative price impact we document reflects not a relaxation of short-sale frictions but rather the revelation of private information held by informed traders likely affiliated with IPO underwriters.

This finding differs from the conventional explanation for options listing effects. Rather than

alleviating short-sale constraints, options introduction appears to intensify them. Both borrowing costs and quantities increase significantly following options listing, indicating that increased demand for short positions outstrips available supply. Market makers' delta-hedging activities drive this increased borrowing demand as they absorb proprietary traders' net short positions.

The information-based trading mechanism explains both outcomes. Proprietary traders accumulate negative synthetic positions through put purchases that strongly predict subsequent stock returns. This predictive power concentrates exclusively among firms whose lead IPO underwriters are OCC clearing members, suggesting that informed traders' information advantages stem from underwriter connections. When market makers absorb these positions, they hedge by shorting stock, simultaneously intensifying constraints while revealing information through price discovery.

These findings contribute to multiple research domains. We provide the first systematic comparison of options listing effects across firm types, testing theoretical predictions of differential impacts across information environments. We advance understanding of informed trading in derivatives by documenting that return predictability from proprietary trader positions is conditional on information advantages, which addresses the surprising finding in Pan and Poteshman (2006) that proprietary traders' aggregate positions lacked predictive power. For the IPO literature, we document that options market introduction can reveal private information held by IPO underwriters and their affiliates, potentially contributing to more efficient pricing of newly public firms.

Our findings present an interesting dynamic: options listing simultaneously worsens short-sale constraints while enabling price discovery through informed trading. This apparent paradox occurs because the mechanisms driving information revelation also drive increased borrowing demand. When informed traders accumulate large positions in options markets, market makers must absorb these positions to clear the market. To hedge their resulting long option exposure, market makers short the underlying stock. This hedging activity simultaneously accomplishes two outcomes: it drives up borrowing demand and costs in the security-lending market, and it communicates market makers' information about informed traders' positioning to other market participants through their trading activity. In this sense, the increased short-sale constraints are not a market friction that impedes efficiency but rather a signal of active information revelation. The short-term increase

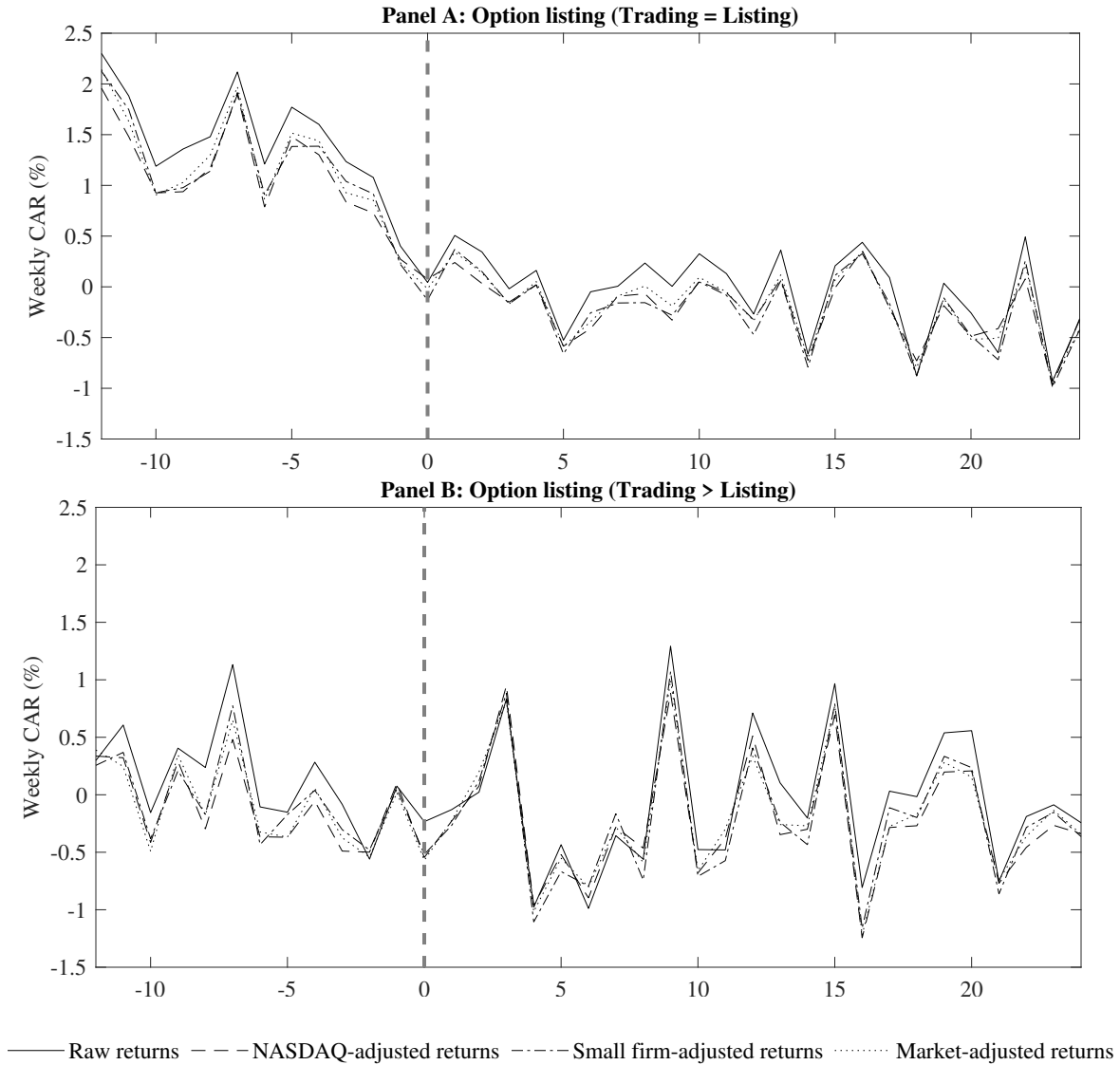
in borrowing costs and quantities represents the market's rational response to the presence of informed trading, where the costs of information discovery are borne through temporarily tighter short-sale constraints.

Our analysis suggests that policymakers and exchanges should carefully evaluate how derivative introductions affect market quality, particularly for securities already subject to trading frictions. The evidence that informed trading through derivatives generates value through information revelation supports the view that derivatives markets contribute to financial stability by improving price discovery, even when they simultaneously increase certain transaction costs.

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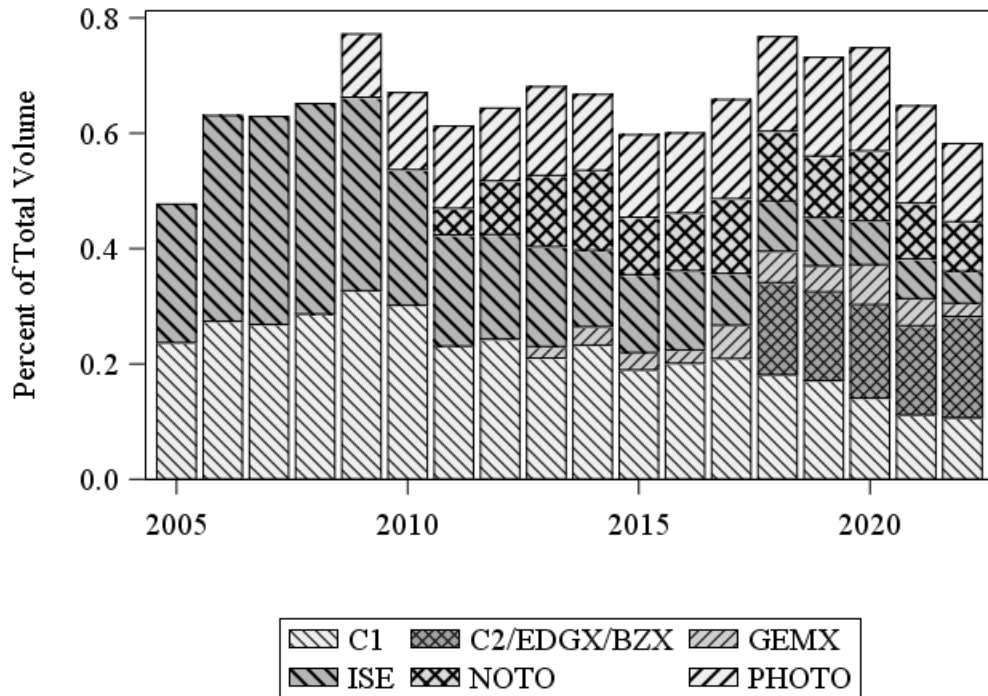
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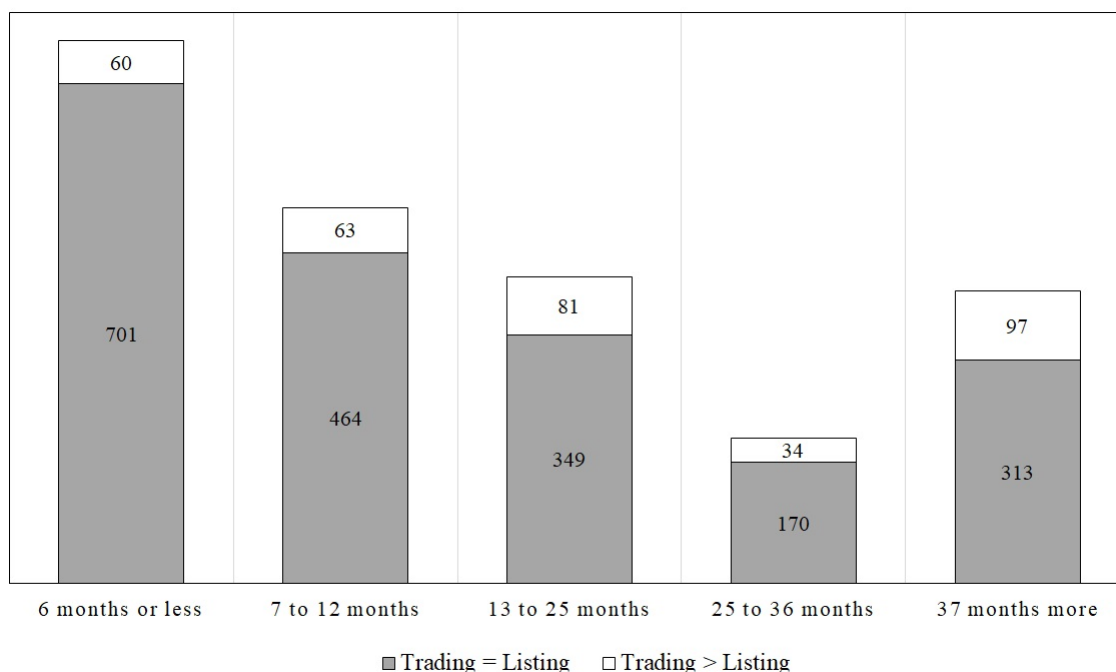
We plot average weekly cumulative adjusted returns (CAR) centered on two events: Option listing (Panels A and B). The sample consists of U.S. companies that went public in 1996–2017 and have option listed on their equity within 3 years after their IPO issuance date. Week 0 refers to the event week. Panel A plots average weekly CARs around 1,568 option listings that are accompanied by immediate option-trading starts (Trading=Listing). Panel B plots average weekly CARs around 232 option listings that have delayed option-trading starts (Trading>Listing). In each panel, we plot cumulative raw returns (Raw return) and CAR series calculated against three benchmarks: (1) CRSP-value-weighted NASDAQ index (NASDAQ-adjusted), (2) lowest decile portfolio based on market capitalization of NYSE index (Small company-adjusted), and (3) CRSP-value-weighted market index adjustment (Market-adjusted).

Figure 1: *Option listing and IPO shares lockup expiration*



We plot the coverage of our equity option open-close data from January 2005 through December 2022. The figure plots the percentage of option volume that originated from the following 8 exchanges: C1, C2, EDGX, BZX, ISE, NOTO, PHOTO, and GEMX. Open-close option volumes contain directional trading activity in the options market for different investor types. The percentage of the total volume is calculated relative to the total equity options volume reported in Optionmetrics. The figure plots the average value calculated over each quarter. Data for the C1, C2, EDGX, and BZX exchanges are obtained from the CBOE. Data for the NOTO, PHOTO, and GEMX are obtained from NASDAQ. Open-close option volume data from exchanges became available at different times. Data from different exchanges were added to the sample as they became available as illustrated in this figure.

Figure 2: Coverage of option trading volume from different exchanges



We plot the distribution of U.S. companies in our sample that went public in 1996–2017 and have options listed on their common equity shares. We group companies based on when option is listed, in months, after their initial public offering (IPO). See Table A1 for the distribution of companies in our sample grouped by their IPO issuance year. Options do not always begin trading immediately after they are listed. Each vertical bar plots the number of option-listing events for two distinct sub-samples: Listing = Trading, and Trading > Listing. We consider the option-listing event to have an immediate option-trading start if options begin trading within 7 days (i.e., Listing = Trading), otherwise, we consider the listing event to have a delayed option-trading start (i.e., Trading > Listing).

Figure 3: *Do options trade immediately after they are listed?*

Table 1: *Descriptive statistics of companies that went public in 1996–2017*

The sample consists of 3,723 US initial public offerings of ordinary common shares between January 1996 and December 2017. We exclude closed-end funds, REIT, Dutch auction offerings, and issuance that are flagged with “shelf registration”. *Offer price* is the share price for the IPO company set by the underwriter. *Venture backed* is an dummy variable equal to 1 if the IPO company is backed by a venture capitalist. *High underwriter rep* is a dummy variable equal to 1 if the company’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 return (%) on the offered price to the end of the first trading day. *Proceeds amount* is the total proceeds amount from the offering. *Total shares offered* is the number of shares issued at the offering in thousands. *Lockup 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, if the company has shares lockup expiration agreement. *Option listed* is a dummy variable equal to 1 if the company has options listed on their equity during the 1996–2021 period and 0 otherwise. Conditional on having options listed, *Days-to-list* is the number of days between IPO issuance and when the option is listed on an exchange. Conditional on having options listed, *Trading-listing gap* is the number of days between the option listing date to when the first option contract is traded.

	Nobs	Mean	Median	Std Dev	Minimum	Maximum
Offer price	3,723	13.43	13	5.80	5	97
Venture backed	3,723	0.46	0	0.50	0	1
First day ret (%)	3,723	25.17	10.42	51.60	-43.27	697.50
Total shares offered (*1000)	3,723	7,709.96	5,000	14,874.40	174.29	478,000
High underwriter rep	3,723	0.43	0	0.49	0	1
Proceeds amount	3,723	17.90	17.91	1.11	14.56	23.48
Underwriter rep	3,670	7.43	8	2.23	0	9
Lockup agreement	3,723	0.78	1	0.41	0	1
Days to lockup expiration	3,019	198.45	180	86.13	15	760
Option listed	3,723	0.63	1	0.23	0	1
Days-to-list	2,332	654.38	287	962.06	8	7,785
Trading-listing gap	2,332	6.37	1	23.91	0	771

Table 2: *CAR around option listing on newly public companies: OLS*

The dependent variable is weekly cumulative adjusted return, $CAR_{i,t}$, for company i in event week t . It is calculated as the weekly raw return on stock i minus the corresponding period market return. The sample consists of U.S. companies that went public in 1996–2017 and have options listed within 156 weeks (3 years) after the IPO issuance date. *Optionstat* is a dummy variable equal to 1 when event week $t \geq 0$ indicates that options have been listed and zero otherwise. We report OLS regression results for two observation windows: $[-12, +12]$ and $[-12, +24]$ weeks relative to option listing. Panel A reports results for option listings with immediate option-trading start (i.e., Listing = Trading). Panel B reports results for option listings with delayed option-trading start (i.e., Trading > Listing). Firm and year fixed effects are included. We include weekly time-varying control variables: stock return momentum, shares turnover, stock return volatility, and market capitalization. These variables are calculated using a one-quarter rolling window and lagged by one week. See text for complete details. Robust t-statistic, clustered at the company and year level, are reported in brackets below each estimate. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

	Weekly CAR			
	Panel A: Trading = Listing		Panel B: Trading > Listing	
	$[-12, +12]$	$[-12, +24]$	$[-12, +12]$	$[-12, +24]$
	(1)	(2)	(3)	(4)
Optionstat	-0.959*** (-6.00)	-1.304*** (-8.73)	-0.348 (-1.36)	-0.117 (-0.50)
Momentum	-6.357*** (-21.44)	-4.167*** (-19.99)	-7.544*** (-8.95)	-4.189*** (-7.30)
Volatility	8.659 (1.00)	10.498** (2.26)	5.451 (0.32)	23.770** (2.06)
Shares turnover	5.165 (0.77)	-4.677 (-0.99)	2.597 (0.16)	-5.300 (-0.32)
Market cap	-11.517*** (-24.40)	-7.434*** (-20.65)	-8.727*** (-7.36)	-6.580*** (-8.03)
Year and Firm FE	Yes	Yes	Yes	Yes
Number of observations	38182	58241	5632	8453
Number of events	1686	1686	239	239
Adjusted R^2	0.0580	0.0385	0.0499	0.0346

Table 3: CAR around option listing on newly public companies: Diff-in-diff

We report difference-in-difference (DID) regression results on weekly adjusted returns around option listing. The results are based on the propensity-score matched sample, i.e., treated and control groups. We report results for two observation windows: $[-12, +12]$ and $[-12, +24]$ weeks relative to option listing. Panel A reports results for option listing with immediate option-trading start (i.e., Listing = Trading). Panel B reports results for option listings with delayed option-trading start (i.e., Trading > Listing). The treated group consists of companies that went public in 1996–2017 and have options listed within 156 weeks after their IPO issuance date. We match companies in the treated group to newly public companies that do not have options listed (Control group) based on their propensity score of having options listed; see Appendix Table A3. *Optionstat* is a dummy variable equal to 1 in event weeks $t \geq 0$ indicating that options have been listed, and zero otherwise. *Treated* is a dummy variable equal to 1 if the company is in the treated group, and zero for the control group. IPO characteristics that are used in the propensity-score matching are included as control variables. Year and industry fixed effects, as well as weekly control variables (unreported) are included in the regressions. Weekly controls consist of one-week lagged variables: return momentum, shares turnover, return volatility, and market capitalization. Robust t-statistic, clustered at the company and year level, are reported in bracket below each estimate. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

	Weekly CAR			
	Panel A: Trading = Listing		Panel B: Trading > Listing	
	$[-12, +12]$	$[-12, +24]$	$[-12, +12]$	$[-12, +24]$
	(1)	(2)	(3)	(4)
Optionstat $\times$ Treated	-0.559** (-2.28)	-0.851*** (-3.68)	0.227 (0.54)	0.105 (0.26)
Optionstat	-0.446** (-2.50)	-0.384** (-2.30)	-0.526 (-1.48)	-0.340 (-0.99)
Treated	0.250 (1.19)	0.253 (1.21)	-0.556 (-1.56)	-0.542 (-1.55)
Year and Industry FE	Yes	Yes	Yes	Yes
Weekly controls	Yes	Yes	Yes	Yes
IPO characteristics	Yes	Yes	Yes	Yes
Number of observations	37782	57337	8027	12132
Number of events	819	819	172	172
Adjusted R^2	0.0074	0.0068	0.0069	0.0048

Table 4: *Short-selling demand and cost around option listing on newly public companies*

We examine the impact of option listing on short-sale constraints. The dependent variables are the weekly average indicative fee and the weekly average quantity-on-loan of shares in the security lending market. *Indicative Fee* is the expected shares borrowing cost for a hedge fund on a given day in the security lending market, measured in annualized percentage rate (%). *Quantity On Loan* is the total quantity of stock on loan on any given day, measured in 100,000 shares. We consider only option-listing events with immediate option-trading starts (Listing=Trading). *Optionstat* is a dummy variable equal to 1 for event weeks $t \geq 0$ indicating that option has been listed, and zero otherwise. Each panel reports results for two observation windows: [-12,+12] and [-12,+24] weeks relative to option listing. Firm and year-fixed effects are included. Weekly controls consist of one-week lagged variables: return momentum, shares turnover, return volatility, and market capitalization; they are calculated using one-quarter rolling windows. *Indicative Fees* is winsorized at the 1% and 99% levels. Robust t-statistics, clustered at the company and year level, are reported in brackets below each estimate. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

	Indicative Fee		Quantity On Loan	
	[-12,+12]	[-12,+24]	[-12,+12]	[-12,+24]
	(1)	(2)	(3)	(4)
Optionstat	0.248** (2.39)	0.358*** (3.04)	3.288*** (8.93)	4.280*** (9.03)
Year and Firm FE	Yes	Yes	Yes	Yes
Weekly controls	Yes	Yes	Yes	Yes
Number of observations	11556	18225	11561	18235
Number of events	616	626	616	626
Adjusted R^2	0.8235	0.7992	0.8053	0.7640

Table 5: *Synthetic net open interests: Descriptive statistics*

This table summarizes weekly synthetic net open interests (SNOI), measured in 100,000 unit shares, over the 24 weeks after option listing. We report summary statistics for three investor groups: *Firm proprietary*, *Customers*, and *Market maker*.

	Puts	Calls	Puts and Calls
	(1)	(2)	(3)
<i>Firm proprietary</i>			
Mean	-0.172	-0.027	-0.199
Median	-0.002	0.000	-0.001
Standard deviation	1.281	0.875	1.707
Min	-39.375	-24.958	-40.336
Max	4.519	10.458	10.332
<i>Customer</i>			
Mean	0.131	0.034	0.165
Median	0.008	0.000	0.012
Standard deviation	1.456	1.839	2.721
Min	-36.464	-23.101	-35.996
Max	25.765	53.418	63.409
<i>Market maker</i>			
Mean	0.041	-0.008	0.033
Median	-0.002	-0.002	-0.007
Standard deviation	1.815	1.455	2.582
Min	-75.840	-18.469	-76.332
Max	20.773	41.241	43.597

Table 6: *Synthetic net open interests: customer versus proprietary traders*

We estimate a weekly panel regression over the 24 weeks after option has been listed. The dependent variable is the weekly SNOI of proprietary traders and public customers. *Firm proprietary*= 1 is a dummy equal to 1 for the SNOI of proprietary traders and 0 otherwise. Models (1), (2), and (3) report OLS regression results trading activities in put options, call options, and both put and call options, respectively. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

	Puts	Calls	Puts and Calls
	(1)	(2)	(3)
Firm proprietary= 1	-0.302*** (-4.93)	-0.061 (-0.69)	-0.364*** (-2.76)
Constant	0.131*** (4.18)	0.034 (0.74)	0.165** (2.44)
Firm and Week FE	Yes	Yes	Yes
Number of observations	26120	26122	26122
Number of events	523	523	523
Adjusted R^2	0.1553	0.1535	0.1685

Table 7: *Synthetic net open interests and future stock returns*

The dependent variable is weekly cumulative adjusted return, $CAR_{i,t+1}$, for company i in week $t + 1$. Independent variables of interest are the synthetic net open interests (SNOI) of different investor types. SNOI represents the net exposure on the underlying stock that is held in option contracts and expressed in the equivalent 100,000 unit shares. *Market makers* is the weekly synthetic net open interest of market makers. *Firm proprietary* is the weekly synthetic net open interest of proprietary traders. *Customer* is the weekly synthetic net open interest of public customers. See text for complete details. Columns (1) and (2) report results for the SNOI measures calculated using put options only, and Columns (3) and (4) report results for call options only. Columns (5) and (6) report results for the SNOI measures calculated using both put and call options. Firm and year fixed effects are included. Weekly controls consist of lagged variables: momentum, average turnover, average volatility, and average market capitalization. Robust t-statistic, clustered at the company and year level, are reported in bracket below each estimate. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

	Weekly CAR [t+1]					
	Puts		Calls		Both	
<i>Synthetic net open interest</i>	(1)	(2)	(3)	(4)	(5)	(6)
Firm proprietary	0.148*** (3.42)		0.130 (1.18)		0.128*** (3.09)	
Customer	-0.105** (-2.18)		0.090 (1.16)		-0.014 (-0.38)	
Market makers		-0.020 (-0.74)		-0.096 (-1.29)		-0.037 (-1.50)
Year and Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Weekly controls	Yes	Yes	Yes	Yes	Yes	Yes
Number of observations	13016	13016	13016	13016	13016	13016
Number of events	523	523	523	523	523	523
Adjusted R^2	0.0381	0.0377	0.0378	0.0378	0.0380	0.0378

Table 8: *Affiliation of proprietary traders with the company's lead IPO underwriter*

The dependent variable is weekly cumulative adjusted return, $CAR_{i,t+1}$, for company i in week $t + 1$. It is calculated as the weekly raw return on stock i minus the corresponding period market return. Independent variables of interest are the synthetic net open interests (SNOI) of different investor types. SNOI represents the net exposure on the underlying stock that is held in option contracts and expressed in the equivalent 100,000 unit shares. Panels A and B report results for the OCC-affiliated IPO and non-OCC-affiliated IPO. We consider an IPO to be OCC-affiliated if the company's lead underwriter was an OCC member during the year of its IPO. The OCC annually publishes the names of its current clearing members. Options trades of OCC clearing members are labeled by exchanges as originating from a proprietary account. *Firm proprietary* is the weekly synthetic net open interest of proprietary traders. *Customer* is the weekly synthetic net open interest of public customers. See text for complete details. In each panel, we report results for the SNOI measures that are calculated using put options only, call options only, and both put and call options. Firm and year fixed effects are included. Weekly controls consist of lagged variables: momentum, average turnover, average volatility, and average market capitalization. Robust t-statistic, clustered at the company and year level, are reported in bracket below each estimate. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

	Weekly CAR [t+1]					
	Panel A: OCC-affiliated IPO			Panel B: Non-OCC affiliated IPO		
	Puts	Calls	Both	Puts	Calls	Both
<i>Synthetic net open interest</i>	(1)	(2)	(3)	(4)	(5)	(6)
Firm proprietary	0.165*** (3.74)	0.123 (1.16)	0.136*** (3.21)	-0.924 (-0.58)	0.911 (0.39)	0.111 (0.07)
Customer	-0.116** (-2.35)	0.111 (1.36)	-0.010 (-0.28)	-0.161 (-0.21)	0.221 (0.32)	0.090 (0.18)
Year and Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Weekly controls	Yes	Yes	Yes	Yes	Yes	Yes
Number of observations	10829	10829	10829	2187	2187	2187
Number of events	435	435	435	88	88	88
Adjusted R^2	0.0346	0.0341	0.0344	0.0548	0.0548	0.0545

Table 9: *Synthetic net open interests and short-selling demand*

We report weekly panel regression results over the event weeks $[0, +24]$ relative to when option is listed on the equity of newly public company. The dependent variables in Panels A and B, respectively, are the average indicative fees measured in annualized percentage rate and the average quantity on loan measured in 100,000 shares. Independent variables of interest are the SNOI of different investor types measured in 100,000 unit shares. *Market makers* is the weekly SNOI of market makers. *Firm proprietary* is the weekly SNOI of proprietary traders. *Customer* is the weekly SNOI of public customers. See text for complete details. Firm and year fixed effects are included. Weekly controls consist of one-week lagged variables: momentum, average turnover, average volatility, and average market capitalization. Robust t-statistic, clustered at the company and year level, are reported in bracket below each estimate. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

<i>Synthetic net open interest</i>	Panel A: Indicative Fees					
	Puts		Calls		Both	
	(1)	(2)	(3)	(4)	(5)	(6)
Firm proprietary	-0.245** (-2.41)		-0.123 (-0.36)		-0.228*** (-2.80)	
Customer	-0.107 (-1.26)		-0.005 (-0.02)		-0.070 (-0.83)	
Market makers		0.176** (2.42)		0.015 (0.06)		0.132* (1.89)
Year and Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Weekly controls	Yes	Yes	Yes	Yes	Yes	Yes
Number of observations	9211	9211	9211	9211	9211	9211
Number of events	384	384	384	384	384	384
Adjusted R^2	0.8148	0.8147	0.8133	0.8133	0.8148	0.8143

<i>Synthetic net open interest</i>	Panel B: Quantity on Loan					
	Puts		Calls		Both	
	(1)	(2)	(3)	(4)	(5)	(6)
Firm proprietary	-2.263*** (-3.89)		-0.494 (-0.31)		-1.581*** (-2.74)	
Customer	0.009 (0.02)		-1.630* (-1.73)		-0.842 (-1.56)	
Market makers		1.132*** (3.50)		1.534 (1.41)		1.130*** (3.33)
Year and Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Weekly controls	Yes	Yes	Yes	Yes	Yes	Yes
Number of observations	9211	9211	9211	9211	9211	9211
Number of events	384	384	384	384	384	384
Adjusted R^2	0.8095	0.8037	0.8011	0.8001	0.8076	0.8064

Internet Appendix

A1 Propensity-score matching

We construct a sample of control firms using recently public firms that are eligible for option listing but *do not yet* have options listed. We use firms that went public in 1996-2017. ADRs, country funds, REIT, and closed-end funds are excluded. We also require that eligible control firms must have options listed within 156 weeks (i.e., 3 years) after their IPO issuance date. This requirement ensures that these control-group firms are, at some point, selected for option listing. In the next step, we estimate the following logit model on all eligible firms each year from 1996 to 2020:

$$L(List_{i,t}) = \alpha_i + \sum_{i=1} \gamma_i IPO\ variable_i + \sum_{j=1} \delta_j Weekly\ Controls_{j,t-1}, \quad (4)$$

where $L(List_{i,t})$ is the log-odds ratio, or the propensity score, that firm i will be selected for option listing during week t . The independent variables that we use are IPO characteristics, denoted by $IPO\ variable_i$, and time-varying covariates, $Weekly\ Controls_{j,t-1}$. IPO characteristics that we use include *Venture backed*, *High underwriter rep*, *Time to lockup expiration*, *First day ret*, and *Proceeds amount*. These variables are defined in Section 3.2. The choice of IPO characteristics is motivated by our preliminary analysis that examines the determinants of when options are listed after their IPO issuance. We include a set of one-week lagged time-varying covariates denoted by $Weekly\ Controls_{j,t-1}$. It includes stock return momentum, stock return volatility, average shares turnover, and market capitalization as described in Section 4.1. Finally, α_i represents the industry fixed effects.

After estimating the logit model in equation (A1), we use the nearest-neighborhood caliper matching approach of Cochran and Rubin (1973) to construct the matched sample on a weekly basis. For each week t , firms in the treated group are newly public firms that have option listed in that week. On the other hand, firms in the control group are newly public firms that *will not* have options listed on their equity for at least 26 weeks (i.e., 6 months) from the current week t . This last requirement is critical. It ensures that control-group firms are free from the effects associated with option listings. We require that the propensity score of observations from the control-group and the treated-group firms are within $\pm 5\%$ of each other. We allow each control firm to be matched with up to two treated-group firms. The final matching yields 991 matched pairs, where 991 unique treated-group firms are matched to 591 unique control-group firms.

Appendix Table A3 reports logit regression results for the samples before and after matching. We report results from estimating the model in Equation (A1) on the full sample (“Before matching”), and on the propensity score-matched pairs (“After matching”). Most of the IPO characteristics and time-varying covariates lose their statistical significance in the propensity score-matched sample. *First day ret* and *Shares turnover* are two only two variables that remain weakly significant. Overall, we find that the treated- and control-group firms are similar in their observable characteristics.

Table A1: *Distribution of when option is listed after the initial public offering (1996–2017)*

The sample consists of 3,723 US initial public offerings on ordinary common shares between January 1996 and December 2017, excluding closed-end funds, REIT, Dutch auction offerings, and issuances flagged with “shelf registration”. This table 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. IPO data is from the Securities Data Company (SDC) New Issues database. Option listing dates are obtained from the Exchange file in Optionmetrics.

Year IPO issuance	Months relative to the IPO issuance date					Never optioned	Total
	6 or less	7–12	13–24	25–36	37 or more		
1996	33	52	87	42	71	355	640
1997	39	33	42	15	49	250	428
1998	34	32	16	11	32	145	270
1999	101	102	30	10	28	158	429
2000	55	59	27	7	42	134	324
2001	30	12	3	5	13	12	75
2002	23	3	10	2	8	17	63
2003	10	5	7	6	16	16	60
2004	34	16	24	26	18	45	163
2005	35	10	20	9	26	37	137
2006	34	16	19	6	22	39	136
2007	26	26	12	12	27	35	138
2008	10	1	1	1	3	2	18
2009	23	1	4	0	2	4	34
2010	33	13	15	9	6	13	89
2011	37	12	10	3	1	9	72
2012	41	17	9	2	6	12	87
2013	56	23	13	6	4	19	121
2014	66	26	42	10	14	22	180
2015	36	10	16	8	11	22	103
2016	17	14	19	2	5	13	70
2017	28	12	7	5	2	32	86
Total	801	495	433	197	406	1391	3723

Table A2: *Summary statistics: Option-listed firms that went public in 1996–2017*

We report descriptive statistics for US companies that went public between January 1996 and December 2017 and have options listed on their ordinary common shares. We separately classify option-listed firms into two sub-samples. Panel A reports descriptive statistics of firms that have immediate option-trading start when option is listed on their equity, i.e., within 7 calendar days. Panel B reports descriptive statistics of firms that have delayed option-trading start after option is listed on their equity, i.e., after 7 calendar days. *Days-to-list* is the number of days between the firm's IPO issuance date and when option is listed on its equity. *Trading-listing gap* is the number of days between the first option trading date and the option listing date. All other variables are defined in Table 1.

Panel A: Trading = Listing

	Nobs	Mean	Median	Std Dev	Minimum	Maximum
Days-to-list	1,997	574.21	247	826.03	8	7,785
Trading-listing gap	1,997	1.95	1	2.05	0	7
Offer price	1,997	14.89	14.50	5.57	5	65
Venture backed	1,997	0.52	1	0.50	0	1
First day ret (%)	1,997	33.09	14.06	61.59	-34.38	697.50
Total shares offered ($\times 1000$)	1,997	10,234.42	6,000	17,904.59	174.29	478,000
High underwriter rep	1,997	0.57	1	0.50	0	1
Underwriter rep	1,974	8.08	9	1.78	0	9
Proceeds amount	1,997	18.31	18.22	1.04	14.56	23.48
LockupFlag	1,997	0.76	1	0.42	0	1
Days to lockup expiration	1,544	182.46	180	38.04	15	730

Panel B: Trading > Listing

	Nobs	Mean	Median	Std Dev	Minimum	Maximum
Days-to-list	335	1,132.29	552	1,454.10	11	7,133
Trading-listing gap	331	33.06	17.00	56.33	8.00	771.00
Offer price	335	14.06	14	7.13	5	91
Venture backed	335	0.42	0	0.49	0	1
First day ret (%)	335	12.46	6.20	23.51	-23.64	206.67
Total shares offered ($\times 1000$)	335	7,777.44	6,150	5,560.78	625	53,191.49
High underwriter rep	335	0.47	0	0.50	0	1
Underwriter rep	330	7.80	8.50	1.94	0	9
Proceeds amount	335	18.21	18.23	0.81	15.32	20.77
LockupFlag	335	0.91	1	0.28	0	1
Days to lockup expiration	308	183.80	180	38.91	90	540

Table A3: *Propensity score matching for newly public firms*

We estimate the firm's propensity of having options listed on any given week following the logistic regression model in Equation (A1). 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 U.S. firms that have gone public in 1996-2017 and have options listed within 156 weeks after their IPO issuance. When firms become seasoned for more than 156 weeks, they are removed from the sample. The regression model (2), titled "After matching," reports results for the propensity-score-matched sample: treated and control groups. The matching is done at the firm-week level using relevant IPO characteristics as well as lagged time-varying covariates, i.e., size, trading volume, volatility, and momentum; see texts for descriptions. Robust t-statistic, clustered at the firm and year level, are reported in bracket below each estimate. Number of observations refers to the number of firm-week used in the estimation. 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)
Venture backed	0.047 (0.86)	0.019 (0.17)
Time to lockup expiration	0.008*** (2.93)	-0.000 (-0.06)
High underwriter rep	0.050 (0.91)	0.053 (0.47)
First day ret	-0.344*** (-7.12)	0.269** (2.18)
Proceeds amount	0.100** (2.23)	-0.096 (-1.02)
Momentum	-0.107 (-1.32)	0.198 (1.37)
Volatility	6.261*** (5.68)	-0.332 (-0.14)
Shares turnover	7.009*** (9.06)	-5.391** (-2.55)
Market cap	0.866*** (20.19)	0.126 (1.45)
Constant	-25.802*** (-19.50)	-1.252 (-0.64)
Year and Industry FE	Yes	Yes
Number of observations	81560	1971
Number of events	1914	991
Pseudo R^2	0.0847	0.0135

Table A4: *CAR around option listing on seasoned firms: OLS*

The dependent variable is weekly $CAR_{i,t}$ for firm i in event week t . The sample consists of seasoned firms going public in 1996-2017 that have options listed 36 months after the issuance date. *Optionstat* is a dummy variable equal to 1 when event week $t \geq 0$ indicates that options have been listed and zero otherwise. Panel A reports results for option listing with immediate option-trading start (i.e., Listing = Trading). Panel B reports results for option listings with delayed option-trading start (i.e., Trading > Listing). Weekly controls consist of one-quarter lagged variables: momentum, average turnover, average volatility, and average market capitalization. We report robust t-statistic is clustered at the firm and year level. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

Dependent Variable	Weekly CAR			
	Panel A: Trading = Listing		Panel B: Trading > Listing	
	[-12, +12]	[-12, +24]	[-12, +12]	[-12, +24]
	(1)	(2)	(3)	(4)
Optionstat	-0.385 (-1.41) (-9.61)	-0.614** (-2.27) (-10.22)	0.033 (0.09) (-8.94)	-0.214 (-0.64) (-10.44)
Year and Firm FE	Yes	Yes	Yes	Yes
Weekly controls	Yes	Yes	Yes	Yes
Number of observations	7703	11312	2380	3488
Number of events	310	310	96	96
Adjusted R^2	0.0706	0.0534	0.0702	0.0555

Table A5: *CAR around option listing on seasoned firms: Diff-in-diff*

We report the difference-in-difference (DID) regression results for weekly adjusted returns around option listing on seasoned firms using the propensity-score matched sample, i.e., treated and control groups. Firms in the treated group are matched to seasoned firms that do not have options listed (control group). The matching is done at the firm-month level following the method of [Mayhew and Mihov \(2004\)](#). We report robust t-statistic is clustered at the firm and year level. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

Dependent Variable	Weekly CAR			
	Panel A: Trading = Listing		Panel B: Trading > Listing	
	[-12, +12]	[-12, +24]	[-12, +12]	[-12, +24]
	(1)	(2)	(3)	(4)
Optionstat $\times$ Treated	-0.337 (-0.58)	-0.262 (-0.47)	1.227* (1.74)	0.752 (1.09)
Optionstat	-1.134** (-2.12)	-1.233** (-2.38)	-1.367** (-2.10)	-0.915 (-1.52)
Treated	0.721 (1.59)	0.622 (1.36)	-0.392 (-0.64)	-0.400 (-0.68)
Year and Industry FE	Yes	Yes	Yes	Yes
Weekly controls	Yes	Yes	Yes	Yes
Number of observations	8175	12120	2848	4184
Number of events	266	266	94	94
Adjusted R^2	0.0170	0.0129	0.0373	0.0247

Table A6: *Short-selling demand around option listing: Seasoned firms*

We examine the impact of option listing on short-selling demand and cost on seasoned firms. The dependent variables are the weekly average indicative fee and the weekly average quantity of loans in the security lending market. *Indicative Fee* is the expected shares borrowing cost for a hedge fund on a given day in the security lending market, measured in annualized percentage rate (%). *Quantity On Loan* is the total quantity of stock on loan on any given day, measured in 100,000 shares. We define seasoned firms as those that have been publicly traded for more than 156 weeks, i.e., 3 years. We consider only option-listing events with immediate option-trading starts (Listing=Trading). *Optionstat* is a dummy variable equal to 1 for event weeks $t \geq 0$ indicating that option has been listed, and zero otherwise. Weekly controls consist of one-week lagged variables: return momentum, shares turnover, return volatility, and market capitalization; they are calculated using one-quarter rolling windows. *Indicative Fees* is winsorized at the 1% and 99% levels. Robust t-statistics, clustered at the firm and year level, are reported in brackets below each estimate. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

	Indicative Fees		Quantity On Loan	
	[-12,+12]	[-12,+24]	[-12,+12]	[-12,+24]
	(1)	(2)	(3)	(4)
Optionstat	0.188 (1.07)	0.268 (1.28)	2.761*** (5.00)	3.294*** (5.45)
Year and Firm FE	Yes	Yes	Yes	Yes
Weekly controls	Yes	Yes	Yes	Yes
Number of observations	6076	9122	6092	9139
Number of events	284	285	284	285
Adjusted R^2	0.9097	0.8780	0.9120	0.8741

Table A7: *IPO characteristics and CAR around options listing on newly public firms*

The dependent variable is weekly cumulative adjusted return, $CAR_{i,t}$, for firm i in event week t . It is calculated as the weekly raw return on stock i minus the corresponding period market return. Model (1) reports results for the balanced window $[-12, +12]$, and Model (2) reports results for the long-run window $[-12, +24]$. The sample consists of firms that went public in 1996–2017 and have options listed within 156 weeks (3 years) after the issuance date. Only option listing events with immediate option-trading start are considered. *Optionstat* is a dummy variable equal to 1 when event week $t \geq 0$ indicates that options have been listed, and zero otherwise. Our variable of interests are the interactions between *Optionstat* and various IPO characteristics. See Table 1 for their descriptions. IPO characteristics, firm and year fixed effects are included. See text for complete details. Weekly controls consist of one-quarter lagged variables: momentum, average turnover, average volatility, and average market capitalization. Number of observations refers to the number of firm-week used in the estimation. Number of events refers to the number of option listings available for the estimation. Robust t-statistic, clustered at the firm and year level, are reported in bracket below each estimate. ***, **, and * indicate significance at the 1, 5, and 10 percent levels.

	Weekly CAR (Trading = Listing)	
	[-12,+12]	[-12,+24]
	(1)	(2)
Optionstat	-0.417 (-0.95)	-1.174*** (-2.84)
Optionstat $\times$ Venture backed	-0.846*** (-2.81)	-0.935*** (-3.34)
Optionstat $\times$ First day ret	-0.700** (-2.15)	-0.444 (-1.49)
Optionstat $\times$ High underwriter rep	0.155 (0.49)	0.701** (2.39)
Optionstat $\times$ Time to lockup expiration	0.008 (0.56)	0.011 (0.84)
Momentum	-6.383*** (-21.43)	-4.201*** (-19.26)
Volatility	6.778 (1.12)	9.151** (2.27)
Shares turnover	7.050 (0.96)	-2.457 (-0.44)
Market cap	-11.742*** (-25.51)	-7.527*** (-22.06)
Year and Firm FE	Yes	Yes
Number of observations	38182	58241
Number of events	1686	1686
Adjusted R^2	0.0586	0.0390