DOCUMENT DE TRAVAIL

N° 495

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Résumé

Nous examinons plusieurs facteurs inexplorés jusqu'ici qui peuvent affecter les taux d'adoption ex post d'options sur actions récemment introduites en bourse. Nous montrons que diverses mesures d'asymétries d'information concernant les actions sous-jacentes prédisent les taux d'adoption des options. Ce pouvoir prédictif est robuste à la prise en compte de facteurs de contrôle dont la littérature a démontré le pouvoir prédictif, comme la volatilité du cours de l'action et le volume échangé. Néanmoins, les introductions d'options induisent une atténuation des asymétries d'information affectant l'action sous-jacente. En outre, les spreads bid-ask des options augmentent dans le temps à partir d'un faible niveau initial, ce qui est cohérent avec l'hypothèse d'un niveau initial d'agressivité des investisseurs informé relativement modeste.

Mots-clés : options sur actions, introductions d'options, asymétyries d'information, taux d'adoption, volumes d'options, intérêt ouvert.

Classification JEL: D82; G10; G14; O31.

Abstract

We examine a number of unexplored factors that affect the *ex-post* adoption rates of newly listed stock options. We show that a variety of measures of information asymmetries concerning underlying stocks predict option adoption rates. These predictive relationships are robust to control factors that have been found to be significant in earlier literature, such as stock volatility and volume. Nevertheless, option listings induce a reduction in the strength of the information asymmetries in the underlying stock. Further, option bid-ask spreads start from low initial levels and increase over time, which is consistent with a modest initial aggressiveness of informed investors.

Keywords: Stock options; option listings; asymmetric information; adoption rates; option volume, open interest.

JEL Codes: D82; G10; G14; O31.

Executive Summary

Financial innovations are pervasive and deeply affect the life experience of investors in all financial markets worldwide. As a result, investors appear to be constantly adapting to new financial products, practices, and institutional arrangements; thus investors engage in a painstaking and costly cognitive process aimed at learning the implications and benefits of financial developments. In this context, probably one of the most significant financial innovations in the last decades is the introduction of derivatives.

Equity option listings are common financial innovations (in the sense that they can be observed multiple times per year) in which completely new option contracts are introduced onto the market for the first time by option exchanges. For instance, the number of stocks with option contracts has grown on average 20% per year between 1973 and 2009 in the United States alone. In 2009 over 3,366 million contracts were traded on more than 3,500 stocks, in contrast to the 911 contracts that were traded on 16 underlying stocks on the first day of trading on the Chicago Board Option Exchange in April 1973.

Regarding the ex-post adoption process, despite the enormous expansion observed in option markets, it is also true that new equity options listed by option exchanges have presented diverse adoption levels among investors. Moreover, some of the options introduced have disappeared in a de-listing process due to low demand for these securities. For instance, 20% of all the equity option listings between 1996 and 2009 in the United States were de-listed in the two years following their introduction dates.

In fact, option listings offer an excellent opportunity to study the adoption process of security innovations since the number of option contracts traded is endogenously determined by investors. In an equity option listing a set of standardized and brand new (i.e. never traded before) option contracts with the same underlying stock are introduced and allowed to be traded in an option exchange. However, for each option contract there is no initial 'established number' of contracts that should be traded in the exchange, which is contrary to other financial offerings where the number of assets is determined exogenously by an institution (e.g., a company decides the number of stocks or corporate bonds to be issued). Instead, in option markets investors themselves create the contracts in an endogenous process based on their demands and following the characteristics of the standardized contracts that offer the option exchange. For instance, an investor has to buy a call option contract while another one has to sell it in order to create that call option contract, which is coordinated by an option market maker.

In this paper, we show that high levels of asymmetric information also predict the ex-post option adoption. The intuition behind the relationship between option adoption and asymmetric information is that heterogeneous levels of information also generate differences of opinions among investors. On the one hand, the general public of investors may wish to hedge the adverse effects of informed trading on their equity positions by trading options written on the stock. In this case, options markets will be perceived as venues in

which uninformed investors try to shield themselves from the existence of informed investors. On the other hand, informed investors may be eager for options markets to be created on the stocks for which they have access to superior information: options offer cheap ways in which private information may effectively be turned into profits. In fact, the literature (see e.g., Anand and Chakravarty, 2007; De Jong et al., 2006, and references therein) tells us that there is strong empirical evidence of informed investors adopting fragmented trading strategies within option markets to try and maximize the trading profits from their private information. Thus, option listings enjoy higher chances of ex-post realized success when the listings concern underlying stocks that are characterized by pervasive information asymmetries. This is also consistent with the theoretical literature (see e.g., Brennan and Cao, 1996; Vanden, 2008) that has emphasized how information asymmetries will normally increase both the demand and the traded volume of option-like derivatives.

We also show that equity options also induce a reduction in the asymmetric information through a learning process based on the private information revealed in the new option market. The learning explanations for the asymmetric information reductions involve two main informational origins. First, option trades (which provides an additional source of private trading information since two markets are now available) should accelerate the rate of disclosure of information from informed investors as result of the new observable market activity (e.g., Jennings and Starks, 1986, and Diamond and Verrecchia, 1987). Secondly, the increasing tendency in the number of analysts (as shown in Figure 1) is also fundamental for changes in the levels of asymmetric information, since more skilled and specialized people facilitate the detection of private information disclosed in trades of informed agents. In fact, we show that equity options reduce the levels of heterogeneous levels of information; in this study they use microstructure-based measures of information asymmetries. Consequently, option trading is expected to improve the informational efficiency of the security market as whole, in the sense that option trades contribute to reveal private information and improve flows of information (e.g., Chern *et al.*, 2008; De Jong *et al.*, 2006; Kumar *et al.*, 1998; Senchack and Starks, 1993).

1. Introduction

Since the first day of negotiations on the Chicago Board Option Exchange (CBOE) on April 26, 1973, when 911 plain vanilla contracts were opened for trading on 16 stocks, the U.S. equity option market has experienced an explosive growth. Between 1973 and 2011, in the United States alone, the equity option volume and the number of optioned stocks have grown on average by 34% and 19% per year, respectively. In 2011 over 1,534 million contracts have been traded on more than 3,684 stocks, for a total cleared value in excess of 426 billion dollars.[1] In spite of the enormous expansion of equity option markets, the newly listed options series have been characterized by rather heterogeneous adoption rates—as proxied by traded option volume and open interest—and hence, rather different success with investors. In fact, some of the options introduced have subsequently disappeared over time (i.e., the underlying stock has stopped being optioned) as a result of de-listings that may be imputed by low demands for the options themselves and not to the liquidation or merger of the underlying stock-issuing company.[2] In this context, our objective is to examine a number of unexplored factors that may affect the *ex-post* adoption process of newly listed stock options; in particular factors related to the levels of asymmetric information in the underlying asset.

Two important features of option markets are the presence of sophisticated and the influential trading activity of informed agents (e.g., see Easley *et al.*, 1998b; Chakravarty *et al.*, 2004; Pan and Poteshman, 2006). In addition, information asymmetries clearly cause—at least before a sequence of trades revealing such superior information—the existence of differences of opinions among investors. These differences of opinion between informed and uninformed agents could induce (or trigger) high levels of trading in new options series introduced into the market. On the one side, informed investors may be eager for

[1] Information obtained from the Option Clearing Corporation (the common clearinghouse shared by all the option exchanges) web page, http:// www.optionsclearing.com.
[2] See "Exchanges Agree To Delist Options; More To Come", *Inside Market Data*, 26 Feb. 2006, Vol. 11 No 11.

stocks on which they have access to superior information to be optioned: options offer cheap ways in which private information may effectively be turned into profits. On the other side, uninformed agents may wish to hedge the adverse effects of informed trading on their equity positions by trading options written on the stock. We therefore empirically investigate whether and how popular microstructure-based measures of information asymmetries (see e.g., Easley *et al.*, 2002), characterizing a stock prior to option introduction, forecast a higher degree of ex-post success of options written on the same underlying asset.

For instance, consider the option series written on Seagate Technology (ticker STX, an Irish-based technology company listed on the NADSAQ), that was introduced on the CBOE on Jan. 21, 2003. This option series (i.e., including all puts and calls that have been created and traded, spanning a range of strikes and maturities over time) has been remarkably successful: for instance, between 2008 and 2012, approximately 4.1 million Seagate options have been traded on the CBOE. In the week that followed the listing of Seagate options, also a new option series written on Sohu.com stock (ticker SOHU, a Chinese internet company listed on the NADSAQ) was introduced on the CBOE on Jan. 27, 2003. Yet—even though both companies seem to have remained healthy enough and both options sets have continuously traded since January 2003—the fate of the two option series has been very dissimilar, in the sense that between 2008 and 2012 less than only 600 thousands contracts written on Sohu.com have traded, roughly one-seventh the number of contracted traded on Seagate. What can explain this glaring difference between two option sets introduced at roughly the same time? Despite this latent and unexplained heterogeneity in the success of new option listings, to our knowledge there is no academic research that has empirically investigated the factors that affect the actual, *ex-post* realized investors' adoption rates *of newly listed option contracts*.

In greater detail, first, we examine the impact of information asymmetries on the success of option listings through a set of econometric tools, and controlling with multiple factors such as stock volatility, volume and market capitalization. Secondly, we track the dynamics of average *option* bid-ask spreads after inception, which gives us indications on the dynamics of the extent of information-based option trading in the aftermath of option introductions. Finally, and thirdly, the empirical literature (e.g., Danielsen and Sorescu, 2001; Skinner, 1990) has widely recognized that option listings affect the asymmetric nature of information flows in the underlying market. Therefore we also analyze the repercussions of option listings on information asymmetries, with a focus on the change of such measureable asymmetries between pre- and after-listing dates.

We use data from listings on the U.S. equity option markets. Option listings are common examples of financial innovations in which new securities (option contracts) are introduced into the market (see e.g., Massa, 2002). In fact, option listings represent a rather special kind of security design innovation in that the number of option contracts traded is endogenously determined by investors: in option listings, a set of standardized contracts written on the same underlying stock are made available for trading on an option exchange.[3] However, for each new option (series) there is no initial "established number" of contracts that should be traded. This is different from other types of offerings in which the number of contracts is exogenously determined by the issuer and price adjusts to bring supply and demand in equilibrium. As a result, in the standard case a failure of the offering (e.g., an equity IPO) simply means a lack of interest and under-subscription that is reflected in a low trading price, which may then be confounded by a myriad of complex, not-well-

[3] Differently from stock markets, where firms voluntarily apply to be listed, decisions to list options are made within the exchanges without any formal application by the stock-issuing company. For example, the bylaws of the CBOE include the criteria for options to be listed. These include share price, number of shareholders who are not insiders, and the trading volume of the underlying stock. Rule 5.4 lists the criteria that will cause the CBOE to stop listing options on a stock. The SEC also plays a role in determining the eligibility requirements for securities to be optioned, see e.g., http://www.sec.gov/rules.shtml.

understood pricing factors. On the opposite, in the case of option listings, we can judge the success or failure on the sheer basis of traded volume and open interest, regardless of the realized, observed price for the newly created contracts.

Studying the determinants of the option adoption process has also ramifications for our understanding of the effects and optimal design of innovation concerning financial securities. It is difficult to minimize the importance of a complete understanding of what determines the success of newly introduced derivatives. On the one hand, a traditional literature has emphasized that derivatives can improve market efficiency by lowering transaction costs (e.g., Merton, 1998), increasing the quality of information flows (e.g., Boehmer *et al.*, 2010; Cao, 1999; De Jong *et al.*, 2006; Kumar *et al.*, 1998), and by reducing the overall level of aggregate, systemic risk (e.g., Klemkosky and Maness, 1980; Darby, 1994). On the other hand, it is under everybody's eyes the fact that unchecked processes of "financialization" based on the introduction of sophisticated securities (e.g., the collateralized mortgage obligations that have been often blamed for the excesses of the U. S. sub-prime bubble, see Coval *et al.*, 2009) may generate instability and cause welfare losses. Additionally, knowledge of the factors that influence the success of option listings may have policy implications especially with regard to the inspiring criteria that should regulate how option exchanges select optioned stocks.

In this paper we consider option dollar-volume, option contract-volume, and open interest as alternative proxies for adoption levels of newly listed options (see Duffie and Jackson, 1989). In addition, we use different measures as proxies of information asymmetries affecting stocks, including estimated, implicit indicators that rely on a range of microstructure models (e.g., *PIN* and adjusted *PIN*, where *PIN* means "probability of informed trading") as well as plausible observable proxies. We obtain a number of important results. First, we show that, even when we control for the effects of lagged stock volume and volatility, an elevated level of the information asymmetry indicators (measured in the year that precedes a listing) results in a higher rate of option adoption among investors than it would otherwise be.[4] Our empirical results highlight that information asymmetries are significant in forecasting the success of a option listing. This positive relationship between the success of a listing and information asymmetries is consistent with few previous theoretical results (see e.g., Brennan and Cao, 1996; Vanden, 2008). Our result that measures of information asymmetries are key predictors of option listing success is robust to using different proxies for the asymmetries, to a range of control variables, and to using either parametric or non-parametric econometric methods.

Second, we find that in the period immediately following an option listing, the option relative bid-ask spread displays on average a low initial level, with a tendency to increase over time. Such a low starting level for illiquidity costs, followed by an upward trend are somewhat surprising because the early "life" of an option should be marked—as one may expect after all kinds of financial innovations—by a relatively high illiquidity, so that comparatively *high* and not low bid-ask spreads are expected.[5] Nevertheless, despite the bid-ask spread surely includes a liquidity/inventory component, it is important to take into account that the other component of the spread reflects premiums that option market makers charge to execute orders in the presence of a non-zero probability of trading with informed agents (see e.g., Bartram *et al.*, 2008). In fact, the low initial values observed for the option relative bid-ask spread after listing can be explained by a modest level of early participation by informed investors in the new option market. It is the low volumes in the market of newly listed options that discourages informed traders away and hence causes relatively low bid-ask spreads. This is sensible because the early stages after option listings are characterised by a reduced trading volume, where even small transactions are

[4] Controlling for stock volume and volatility is important because previous papers (see e.g., Mayhew and Mihov, 2004) have revealed that these are key factor to explain which stocks become optionable.

[5] As standard in the literature, the inverse of the relative bid-ask spread is often used as a measure of liquidity, see among others Amihud and Mendelson (1986) and Conroy *et al.* (1990).

noticeable and any trading activity by informed agents might be easily detected. As a result, informed investors will have incentives to hide their option trades by fragmenting them, or to simply wait for higher volumes, as reported by Chakravarty *et al.* (2004). Interestingly, the upward trend in the option relative bid-ask spreads is more pronounced for option listings characterized by strong information asymmetries. This is because these listings imply that more informed investors shall need to start trading progressively, at a measured pace, to hide their presence as volume slowly picks up (see Mayhew *et al.*, 1995).

Third, also using a control sample methodology designed to correct for the endogeneity of option flotation, we find that information asymmetries in the underlying stock market significantly decrease after option listings. This is caused by the fact that option trading is expected to improve the informational efficiency of the security market as whole, in the sense that option trades contribute to reveal private information (e.g., Chern *et al.*, 2008; De Jong *et al.*, 2006; Kumar *et al.*, 1998; Senchack and Starks, 1993). In particular, option trades accelerate the rate of disclosure of information from informed investors as a result of the newly observable market activity (as predicted by theoretical models since Diamond and Verrecchia, 1987, and Jennings and Starks, 1986). Moreover, option listings create space and incentives for additional information collection and dissemination which may improve the analysis and interpretation of the information revealed by informed agents through their trading, as implied by the theoretical analyses by Cao (1999) and Massa (2002). For instance, we find that the number of analysts increases significantly after option introductions, similarly to Skinner (1990). However, our results are stronger because they extend to and are based on fine microstructure measures of information asymmetries such as PIN, which have been used widely in the literature as a proxy for asymmetric information.

Moreover, and also in relation to asymmetric information reductions after option introduction, empirical evidence shows that the U.S. national system of options exchanges has become progressively more informationally efficient and better integrated with the underlying spot equity markets (see e.g., Battalio *et al.*, 2004). One of the important dimensions of efficiency in financial markets has long been identified with the reduction in information asymmetry between insiders and the general public of investors. Following a classical Grossman-Stiglitz's (1980) perspective, in this paper we ask whether stocks that are characterized—prior to option listing—by high degrees/likelihood of information asymmetries may enjoy greater chances of success (measured by option volume and open interest) when they are made optionable. Because a literature has argued on both theoretical and empirical grounds that trading in derivatives may be an important channel through which information is compounded into asset prices, such a success would also reduce the asymmetries *ex-post* and favour the overall efficiency of the financial system.

An earlier literature (see e.g., Mayhew and Mihov, 2004; Danielsen *et al.*, 2007) has analyzed a related phenomenon. These studies explore which factors are used by a option exchange to select a stock as an optioned one, and they the find that stock volume and volatility are the variables more relevant for exchanges to make this decision. However, this is only an *ex-ante* perspective on the phenomenon, given that these papers do not study the *ex-post*, effectively realized adoption rate (success) that follows a listing. To our knowledge, such an *ex-post*, realized perspective on option listing success is missing from the literature and of considerable importance to judge whether exchanges may effectively lure insiders to trade stock options, hence supporting the newborn derivative markets but also fostering the overall informational efficiency. Nevertheless, in a way, our analysis provides *ex-post* corroboration of the results in this early literature, by showing that option exchanges are on average "right" in terms of selecting optioned stocks using the factors in Mayhew and Mihov (2004) and Danielsen *et al.* (2007): stock liquidity and volatility are indeed good predictors of the actual, realized adoption rate after a listing. However, and differently, we show that *other factors* should be also taken into account to decide whether a stock can be used to introduce equity options; especially and as mentioned previously, factors related to levels of asymmetric information on the underlying market.

Our empirical analysis is also related to a number of theoretical studies, in which the introduction of derivatives in the presence of information asymmetries are jointly researched, although the emphasis of these papers is not specifically on the ex-post success of option listings. Brennan and Cao (1996) and Vanden (2008) present models in which information asymmetries are endogenously, positively related to option demand and volume. Cao (1999) finds that the introduction of derivatives could intensify the incentives to acquire additional information about the underlying asset payoffs. Massa (2002) develops a model with endogenous information acquisition when a derivative is introduced and where two types of agents exist, informed and uninformed investors. He shows that the incentives of the uninformed investors to purchase information depend on the market informational structure.[6]

Section 2 further discusses our empirical hypotheses in the light of the literature. Section 3 describes the data and the micro-structure indicators of information asymmetry. Section 4 presents our main findings. Section 5 documents that option bid-ask display an upward trend that can be explained by informed traders hesitating before trading in thin markets. Section 6 analyzes the reduction in information asymmetries that follows the listing of options. Section 7 concludes.

[6] A number of papers (e.g., Duffie and Jackson, 1989; Duffie and Rahi, 1995) have studied the design of derivatives in a theoretical perspective, but without any empirical analysis of the effects of information asymmetries on the rate of option adoption. In addition, a literature has investigated the factors that affect the ex-post realized adoption of futures contracts, for instance Nothaft *et al*. (1995) and Corkish *et al.* (1997). These studies find that stock volatility, liquidity, and market capitalization are key drivers. However, they do not study the interaction of asymmetric information and the adoption of derivative contracts.

2. Literature Review and Hypothesis Development

In this Section we develop three testable hypotheses that find their background and motivation in the literature on explanatory factors of option adoption. Previous empirical literature (see e.g., Anand and Chakravarty, 2007; De Jong *et al.*, 2006, and references therein) presents strong evidence of informed investors adopting trading strategies within option markets, and thus to try and maximize the profits from their private information. However, the existence of information asymmetries also causes differences of opinions among market participants. On the one hand, the general public of investors may wish to hedge the adverse effects of informed trading on their equity positions by trading options written on the stock. In this case, options markets will be perceived as venues in which uninformed investors try to shield themselves from informed investors' trades. On the other hand, informed investors may be eager for the stocks on which they have access to superior information to be optioned: options offer cheap ways in which private information may effectively be turned into profits. For this reason, we expect option listings to enjoy higher chances of ex-post realized success when the listings concern stocks that are characterized by pervasive information asymmetries. We therefore hypothesize the existence of a causal link between any information asymmetries plaguing the market for the underlying stock prior to options listing and the ex-post realized success of such listings:

Hypothesis 1: High prior (to listing) information asymmetries affecting the market of the underlying stock predicts a high rate of adoption of newly introduced options, even after controlling for any effects from simple differences in beliefs.

Interestingly, if this causal link between information asymmetries and option listings were to be at work as our hypothesis 1 implies, then we expect not only the asymmetries to increase the chances of success of a listing, but to heavily affect also the way in which this success practically unfolds. Here one may naively expect that although a successful options market is characterized by high trading volume, this may follow a simple upward trend as the newly created market takes off. However, this conjecture fails to take into account the actual incentives of informed traders when it comes to operate in option markets. If these markets also exist to provide informed investors with a cheap way to turn information into profits, we know (see e.g., Lee and Yi, 2001) then they will require adequate volumes to hide their trades behind the flow motivated by hedging and liquidity demands. This is consistent with the hiding strategies of informed option traders reported by Anand and Chakravarty (2007).[7] The market makers intermediating the flow of trading in the newly created option markets will recognize this pattern of behaviour of informed agents. This should lead them to progressively *increase* the component of the bid-ask spread which provides protection against dealing with traders with superior information.[8] As a result, our notion of a successful market may have straightforward implications for volume and open interest, but not for measures of market liquidity such as the *option bid-ask spread*. Note that differently from earlier literature (e.g., Danielsen *et al.*, 2007), our conjecture concerns liquidity and quality in the options market after option listing, and not the effects of option listing on the underlying stock market. However, absent a formal

[7] Anand and Chakravarty (2007) show that informed agents should optimally apply stealth trading strategies by fragmenting their orders into small (medium) trades for low (high) volume contracts. In addition, Biais and Hillion (1994) and Easley et al. (1998b) describe how an informed trader will arbitrage between spot and derivatives market when selecting where to trade, on the basis of their comparative depth and liquidity, and the amount of leverage achievable with options. DeJong et al. (2006) find that insiders trade aggressively in both the stock and the option, and typically trade in the market that affords the most profitable trading opportunity. Lee and Yi (2001) find that the adverse selection component of the bid–ask spread for all trade sizes is greater on the Chicago Board Options Exchange than on the New York Stock Exchange, which suggests that option traders are more information-motivated than stock traders.

[8] Using intraday data for a sample of CBOE options, Lee and Yi (2001) have shown that large trades in the options market may be hardly anonymous, which might enable options market makers to screen large informed trades more effectively than in the stock market. This lack of anonymity in the options market will cause large investors with private information to behave differently than small investors. Kaul *et al.* (2004) have found that the adverse selection component of the underlying stock's spread explains a significant fraction of the option spread, i.e., that information asymmetries propagate from spot to derivatives markets.

theoretical model, we expect only weak and variable links between the success of a listing and the option bid-ask spread, as summarized by:

Hypothesis 2: There is no simple, linear relationship between the rate of adoption of newly introduced options on a stock and the post-listing liquidity of options as measured by (relative) bid-ask spread measures that reflect an adverse selection component.

One more type of feedback occurring after option listing is worth of investigation: the changes in information asymmetries after the listing date. Trading in the newly introduced options would increase the speed at which private information is incorporated into observed option and stock prices. As a result, measurable information asymmetries in the period following an option listing should decline, as in the theoretical analyses by Back (1992), Brennan and Cao (1996), Kraus and Smith (1996), and Vanden (2008):

Hypothesis 3: Option listings reduce the information asymmetries affecting the underlying stock.

Therefore, the causal link that our paper wants to emphasize has, after all, a "happy ending", if we assume that the goal of an efficient capital market is to compound all existing information into traded asset prices cheaply and quickly. The most resilient and successful option listings will concern underlying stocks plagued by strong information asymmetries. As a result, informed traders will progressively flock to the newly established option venue and do their best to extract the highest possible value from their information. As they do so, this makes the option market successful (because they generate volume and open interest), but also structurally not as liquid as initially "hoped" (and recorded, see Figure 1, section 5). Moreover, the process of price discovery that they trigger leads to a decline in the strength of the information asymmetries in the sense that in the underlying stock market there are less information-driven trades (see Faff and Hillier, 2005), while the cheap and effective mechanisms of the options market favour a faster compounding of information into prices (see Chakravarty *et al*., 2004). This last link of the chain has a classical Grossman and Stiglitz's (1980) flavour: informed traders are rewarded for their activity of acquiring information and taking it to the market; as they perform this role, they cause their superior information to depreciate and to be incorporated into prices. The only, to us rather major, difference in the story of this paper is that—as a result of this virtuous mechanism—in the end the financial system finds itself enriched of a new and useful informational conduit, the new option market.

Our paper also complements the analysis performed in three related papers by Roll, Schwartz, and Subrahmanyam (2009, 2010, henceforth RSS) and Choy and Wei (2012). RSS (2009) connects measures of intensity of option trading to the positive effects of informed trades in terms of reduced information asymmetries, more efficient asset valuations, and hence higher corporate investments, showing that firms with greater levels of options trading have higher valuations, as captured by Tobin's q. Interestingly, they also report that the effect of options trading on firm valuation is stronger in stocks with low analyst following and high values of PIN. Differently, from RSS (2009), in our paper we focus not the effects of option trading intensity on firms' behaviour conditioning for the strength of information asymmetries, but directly tract the effects of ex-ante asymmetries on option trading intensity, interpreted as a measure of ex-post success of option listings. RSS (2010) have investigated option trading volume relative to the volume in underlying stocks (the O/S ratio) and show that their ratio increases significantly in the few days around an earnings announcement, which implies that informed traders are most likely to trade in the derivatives markets. Although portions of our analysis rely on the assumption that informed traders may move from the stock to the option market when the latter is newly opened (see Pan and Poteshman, 2006), our empirical analysis is entirely devoted to volume—one of our measure of success of option listing—in the derivatives market, whereas O/S cannot be defined for periods preceding an option listing, and it remains characteristically modest (although it is increasing and time-varying) for newly listed option series. Finally, Choy and Wei (2012) have studied the motives of option trading in general (i.e., not only when new options are listed) but conditioned instead on earnings announcements, similarly to RSS (2010). They conclude that most trading derives from uninformed speculation and differences of opinion. Therefore, our focus is different since we conclude that at least in the case of new listings, *the underlying stock information asymmetries* play an important role in explaining the ex-post success of options. Moreover, our events are not represented by corporate announcements (scheduled or not) but instead by the very listings of options.

3. Data and Construction of Asymmetric Information Measures

We use daily data on equity options (calls and puts) traded on the entire U.S. option market from the OptionMetrics database covering the period January 4, 1996 to October 30, 2009.[9] The data base contains daily information, including closing bid and ask quotes, volumes traded, and open interest. We calculate a proxy for the option dollar-volume by multiplying the number of transactions for each option contract by its end-of-day quote midpoint, and then aggregate this amount across all option contracts written on the same underlying stock, across maturities and strikes. All options that were listed for the first

[9] Between 1973 and 1975, options were traded only on the Chicago Board Options Exchange (CBOE). In 1975 and 1976, option trading started on the American (Amex), Philadelphia, Midwest, and Pacific Stock Exchanges, but the Midwest exchange dropped out of the business in 1980. As of the end of our sample, options were also traded on the NYSE (ARCA), on the BATS exchange, on the BOX Exchange, on the International Securities Exchange (ISE), and on NASDAQ options circuit. Note that multiple listings of options are allowed and have become increasingly common after 1999. Since 1991 the SEC has also allowed the listing of options on securities other than common stock, such as preferred non-convertible stock, ADRs and index funds. However, in this paper we only consider optioned common shares of stock.

time in our sample period are selected and are the object of our investigation.[10] Because our goal consists of an analysis of the dynamics of the adoption process, in what follows we treat the listing date as the initial, "day zero" across all option listings, even though it is clear that these occur at different points over the calendar period 1996-2009. This means that our empirical analysis is performed in event time and not in calendar time, although controls for the state of the market and of the economy that reflect calendar time conditions will be employed. Additionally, some exclusionary criteria are applied. All options whose underlying stock are affected by company events that may influence the measures of adoption (i.e., option dollar-volume, option volume, and open interest) such as splits, mergers, spin-offs, new equity issues, right offerings, or warrant issuing in the year following day zero, are excluded.

The goal of our study is to analyze the role of asymmetric information in the option adoption process, and the behaviour of market participants when reacting to a new option listings. One key step therefore consists of the construction of measures of asymmetric information concerning the optioned stock. We resort to three alternative measures of asymmetric information. First, we calculate the probability of informed trading (*PIN*), which has been widely used in the literature since the seminal paper by Easley *et al.* (2002) (see e.g., Bharath *et al.*, 2009; Chan *et al.*, 2008; Roll *et al.*, 2009, for recent applications). Our second proxy of asymmetric information is the adjusted probability of informed trading (*AdjPIN*), an alternative measure of informed trading proposed by Duarte and Young (2009) to correct the fact that the standard *PIN* may often capture spurious liquidity effects. Finally, as a third measure of information asymmetries, we use the number of analysts which is a directly observable proxy for the level of asymmetric information (see Skinner, 1990). The use of the number of analysts is based on the premise that it should be

[10] Because multiple listings are possible, option may be introduced in one exchange even if the same (or at least, related) option contracts have been listed before in other exchanges. Therefore, we focus our attention exclusively on listing events in which the contracts are genuinely new and not already traded.

easy for the market as a whole to detect any private information in trades when many highly trained observers, such as analysts are, were to analyze market activity, which should lead to a reduction in information asymmetries.[11,12]

The *PIN* and the *AdjPIN* indices are computed for each stock and they reflect the probability that orders concerning the stock may reflect informed trading. They are calculated from two different microstructure models that we summarize in Appendix A (available in the supplementary material) for the sake of completeness. Trade and Quote (*TAQ*) transaction data concerning the stocks underlying the option listings are used to compute estimates of *PIN* and *AdjPIN*. Resorting to *TAQ* data to measure information asymmetries effectively restricts our listing sample to optioned NYSE stocks. Data from *TAQ* are filtered using the same criteria as in Huang and Stoll (1996) and Danielsen *et al.* (2007). For example, we omit trades and quotes if they are flagged as being in an out of time sequence or involve either an error or a correction; we omit quotes if either the ask or bid price is zero or less; we omit trades if the price or volume is not greater than zero. The *PIN* and *AdjPIN* measures are calculated both with reference to the year preceding and following the option listing.[13] Trades are classified as buys and sells following Lee and Ready's (1991) algorithm, as the likelihood functions presented in Appendix A use the daily number of buys and sells for each stock as data. Moreover, similarly to Easley *et al.* (2002), we exclude stocks for which we cannot find in TAQ at least 60 complete days of data concerning quotes and trades in the year prior to and the year following the listing of

[11] The use of the number of analysts is however also supported by Easley *et al.* (1998a) who state that: "(...) high analysts stocks face a lower probability of information-based trading (...)" (p. 200). Also Roll et al. (2009) supplement their analysis based on PIN with the (inverse of the) number of analysts.
[12] Another possible proxy for asymmetric information is the dispersion of analysts' forecasts. However, we do not focus on this measure because Barry and Jennings (1992) have shown that diversity of opinions among analysts can increase even though the level of asymmetric information objectively declines.
[13] The estimation of *PIN* and *AdjPIN* on a pre- and post-listing annual basis is due to the fact that the level of accuracy of the estimates decreases enormously when these are computed over shorter periods of time, see e.g., Easley *et al.* (2002). For the purposes of this paper, a year is defined as a period of 252 trading days that precede (follow) the listing, with day zero excluded from both the prior and post-periods.

options written on a given stock, so to obtain reliable *PIN* and *AdjPIN* estimates. This criterion also rules out from our analysis all stocks subject to recent initial public offerings.

We use Thomson's $I/B/E/S$ to extract data at monthly frequency concerning the total analyst following for each stock under investigation. We calculate the annual average of the monthly number of analysts publishing earnings forecasts for each of the newly optioned stocks; this estimate is produced for both the year before and the year after a listing. Because *PIN* and *AdjPIN* estimate the strength of the information asymmetries concerning a given stock and earlier literature has argued that low analyst following implies high information asymmetries, we also consider the inverse of the average of the number of analysts (*InvAnlst*) following a stock as a measure of information differentials alternative to *PIN* and *AdjPIN*.

Finally, the equity data required to compute the control variables used in this study (stock volatility, stock dollar-volume, and market capitalization) within our formal econometric analyses, are all obtained from the daily CRSP database and were available for all stocks for which new options were listed, as one would expect.

As a result of the merging of all these data sets as well as of the exclusionary criteria listed above, we obtain a final sample of 891 option listings for which we can estimate appropriate *PIN*, *AdjPIN*, and *InvAnlst* measures for both the year prior to and the year following the listings, over our 1996-2009 sample. Table 1 reports summary statistics for the key variables in this paper: $DVlm_{OP,1Y}$, $Vlm_{OP,1Y}$, and $OInt_{OP,1Y}$ are the averages of the daily option dollar-volume, option volume (number of traded contracts), and open interest, respectively, in the year after the option listing; $BAre_{OP,1Y}$ is the average of the relative bid-ask spread for the option during the year following the listing. PIN_{0Y}, $AdjPIN_{0Y}$, and $InvAnlst_{0Y}$ are the *PIN* estimate, the *AdjPIN* estimate, and the inverse of the average number of analysts, referring to the year prior to listing; PIN_{1Y}, $AdjPIN_{1Y}$, and $InvAnlst_{1Y}$ are the *PIN*, the A*djPIN*, and the inverse of the average number of analysts, concerning the year following the listing.[14] Interestingly, Table 1 shows preliminary evidence concerning a potential relation between changes in the strength of information asymmetries and new option listings: The mean of the *PIN* estimates, the *AdjPIN* estimates, and the inverse number of analysts all strongly decline after the listing date, by −25.05%, −23.52%, and −24.24%, respectively. Simple tests for differences in means (proportions, under the assumption of constant variance) reveal that all these changes in information asymmetry indices are strongly statistically significant. For instance, the cross-sectional average of *AdjPIN* declines from 0.17 in the year prior to listing, to 0.13 in the year that follows the listing; the cross-sectional median for *AdjPIN* declines from 0.16 in the year prior to listing, to 0.13 in the year that follows the listing. Estimates of the decline in measured information asymmetries are even stronger in the case of *PIN* and *InvAnlst*. In section 5, we further test the significance of these changes using formal econometrics, different sub-samples, and conditioning for known covariate factors.

[Insert Table 1 here]

4. Key Empirical Findings

We use the option dollar-volume, the option volume, and the open interest as measures of adoption for newly listed options. The option dollar-volume and option volume are selected because these are the measures of success tracked by option exchanges when they assess a listing of new derivative securities (see Duffie and Jackson, 1989).[15] In addition, we use open interest as a measure of the success of a listing because in this market the number of contracts is established in an endogenous process based on

[14] Appendix B, available in the supplementary material, reports the cross-sectional distribution of the parameter estimates for the micro-structure models described in Appendix A which are used to calculate *PIN* and *AdjPIN* (see Tables AI and AII).

[15] Quoting Mayhew and Mihov (2004): "Presumably, the main objective for an exchange is to maximize long-term profits for its members. In practice, according to industry sources, this amounts to listing those options expected to generate the highest trading volume." (p. 450)

investors' demands. Therefore, after an option listing, the open interest truly represents the willingness of investors to participate in and trade the newly listed securities.

The first objective of our study is to examine whether factors related to the asymmetric information in the underlying assets can predict the *ex-post, realized* success of equity option listings. As an initial step, we use simple regressions estimated to test the significance of the determinants of option adoption. In the regressions, the dependent variables are the adoption levels (i.e., $DVlm_{OP,1\,Y}$, $Vlm_{OP,1\,Y}$, and $OInt_{OP,1\,Y}$), while the explanatory variables are the asymmetric information measures. The asymmetric information measures are the probability of informed trading ($PIN_{0\,Y}$), the adjusted probability of informed trading ($AdjPIN_{0\,Y}$), and the inverse of the number of analysts ($InvAnlst_{0\,Y}$) in the year prior to the listing date. Further control variables (again, all calculated with reference to the year prior to listing) are included to coincide with the same "ex-ante" factors that Mayhew and Mihov (2004) identify as the main predictors of stock selection by exchanges for option listing: stock dollar-volume, distinguishing between its long-term ($DVlm_{S,252,0\,Y}$) and short-term ($DVlm_{S,21,0\,Y}$) components (daily averages using the previous 252 and 21 trading days, respectively); stock volatility, distinguishing between its long-term ($SDev_{S,252,0\,Y}$) and short-term ($SDev_{S,21,0\,Y}$) components (the annualized standard deviation of stock daily log returns over the 252 and 21 trading days, respectively); and total stock market capitalization ($Size_{0\,Y}$) for the year prior to listing.[16] As discussed in sections 1 and 2, the inclusion of these control factors is important for us, because we want to show that information asymmetries predict ex-post realized option success even when the standard ex-ante factors—stock trading volume and volatility—are taken into account. Therefore, our formal analysis is described by three simple linear models:

[16] Results are not sensitive to defining the explanatory variables with reference to the period that goes between 252 and 20 trading days before the option introduction, as in Danielsen *et al.* (2007). As they argue in their paper, the exclusion of the 20 trading sessions before option trading starts is sensible if the goal of the study is to isolate which factors are actually taken into account by options exchanges that rule over which stocks should be optioned, which is however not our objective here.

$$log\big(DVlm_{OP,1Y}\big) = \phi + \beta_1 Asymmetry_{0Y} + \gamma_1 log\big(DVlm_{S,252,0Y}\big) + \gamma_2 \frac{DVlm_{S,21,0}}{DVlm_{S,252,}}$$
$$+\gamma_3 SDev_{S,252,0Y} + \gamma_4 \frac{SDev_{S,21,0Y}}{SDev_{S,252,0Y}} + \gamma_5 log(Size_{\,0Y}) + \varepsilon, \tag{1a}$$

$$log(Vlm_{OP,1Y}) = \phi + \beta_1 Asymmetry_{0Y} + \gamma_1 log(DVlm_{S,252,0Y}) + \gamma_2 \frac{DVlm_{S,21,0Y}}{DVlm_{S,252,0Y}}$$
$$+\gamma_3 SDev_{S,252,0Y} + \gamma_4 \frac{SDev_{S,21,0Y}}{SDev_{S,252,0Y}} + \gamma_5 log(Size_{\,0Y}) + \varepsilon, \tag{1b}$$

and

$$log(OInt_{OP,1Y}) = \phi + \beta_1 Asymmetry_{0Y} + \gamma_1 log(DVlm_{S,252,0Y}) + \gamma_2 \frac{DVlm_{S,21,0}}{DVlm_{S,252,0}}$$
$$+\gamma_3 SDev_{S,252,0Y} + \gamma_4 \frac{SDev_{S,21,0Y}}{SDev_{S,252,0Y}} + \gamma_5 log(Size_{\,0Y}) + \varepsilon \tag{1c}$$

where the variable $Asymmetry_{0Y}$ is identified with (a transformation of) either *PIN*, *AdjPIN*, or *InvAnlst*, ε is a random (measurement) error, and $log(\cdot)$ is the natural logarithm.[17] The short term components of liquidity and volatility are expressed in terms relative to the long-term components, to separate the two effects, i.e., $DVlm_{S,21,0Y}/DVlm_{S,252,0Y}$ and $SDev_{S,21,0Y}/SDev_{S,252,0Y}$ as in Mayhew and Mihov (2004).

The estimated coefficients from the models in equations (1a)-(1c) are reported in Table 2. The results in Table 2 are consistent with an *ex-ante* selection policy by option exchanges that consists in introducing option contracts only/principally written on stocks with high volume and high volatility, as already found by Mayhew and Mihov (2004) and Danielsen *et al.* (2007). Table 2 shows that also the *ex-post* levels of option adoption are positively and significantly related to stock volume and volatility in the year prior to listing. The long-term and short-term components of the stock dollar-volume are positively and significantly related to all the option adoption measures (fourth and fifth columns,

[17] Because *PIN*, *AdjPIN*, and *InvAnlst* are either constructed as or are scaled to range between zero and one, their values are logistically transformed before correlating them with other variables.

respectively, in Table 2). The economic magnitude of such effects are indeed large: for instance, holding all other factors the same and assuming that information asymmetries are measured by *AdjPIN*, a 1% increase in measured *stock* volume in the year prior to listing, increases post-listing *option* dollar-volume by 0.83%. A 1% increase in the measured short-term relative *stock* dollar-volume component starting from the cross-sectional average of this value raises increases post-listing *option* dollar-volume by 0.41%. However, only the long-term component of stock volatility is positively and significantly related to option adoption (sixth column in Table 2); the relationship between listing success and the short-term component of stock volatility is instead negative but insignificant (seventh column in Table 2).[18] The negative coefficients observed for the short-term component of stock volatility are however consistent with the results in Mayhew and Mihov (2004) who find evidence of a tendency in option exchanges to list options in periods when there is a decreasing stock volatility, probably a proxy for quiet market states. This finding shows that Mayhew and Mihov's results concern not only how option exchanges select stocks to become optioned, but also the fact that the exchanges are rightly using these factors to single stocks out, because volume and volatility are also precisely estimated and economically meaningful predictors of actual listing success.

[Insert Table 2 here]

We expect the underlying stock market volume to be an important predictor of ex-post realized option success because stock option exchanges are likely to form rational expectations. These are member-owned organizations in which listing decisions are made by the members whose profits are an increasing function of trading activity. If the exchanges are on average right in their choice, they will introduce new option contracts that are ex-post successful, in the sense that a strong volume in the spot market is correctly

[18] The economic effect of a 1% increase in long-term stock return volatility is however rather sizeable, a 26.7% increase in dollar-volume for options when information asymmetries are measured by *AdjPIN*.

anticipated to move to generate (or even, spill over to, in a positive-sum-game) high volume in the derivatives market. Similarly, we expect that the underlying stock return volatility will predict the ex-post realized option success. This occurs because for high volatility stocks, new information hits financial markets at a faster rate thus creating a higher potential price movements that can be traded on and/or hedged by taking appropriate positions in derivatives.

However, stock volume and volatility do not represent the '*whole story*', as far as ex-post adoptions are concerned, which is one of the main contributions of our paper. Table 2 also shows that the success of a listing is positively and significantly related to prior information asymmetries, even when one controls ex-post for Mayhew and Mihov's factors. Table 2, columns 1-3 all show evidence of a positive and significant relationship between *PIN, AdjPIN*, and *InvAnlst* (in the year prior to listing) and adoption levels (in the year following listings). The economic effects are large: for instance, holding all other factors constant, a 1% increase in *AdjPIN* (starting from its cross-sectional mean prior to listing, in Table 1) leads to an increase in post-listing option dollar-volume of 4.39%; the comparables values of the estimated elasticities for *PIN* and *InvAnlst* are 6.11 and 0.87, respectively. Results are similar as far as the effects on contract volume and open interest are concerned. For instance, a 1% increase in *AdjPIN* (starting from its cross-sectional mean prior to listing) leads to an increase in post-listing option contract volume of 4.39% and in open interest of 13.00%. Note that in the light of the recorded declines (see Table 1) in the values of *AdjPIN* (−25.05%), *PIN* (−23.52%), and *InvAnlst* (−24.24%) between prior and post-listing dates, an experiment based on a 1% change in these indicators is extremely realistic. Therefore, Table 2 supports the existence of significant, positive effects of information asymmetries on the adoption of new option contracts, consistently with our hypothesis 1.

Table 2 also shows that the total market capitalization of the underlying stock prior to listing shows a negative relationship with option success, although with insignificant coefficients (eighth column). This negative association is also coherent with our argument that information asymmetries are positive predictors of the success of equity option listings: larger firms normally suffer from lower information asymmetries because they receive more attention from analysts and regulators (see e.g., Bhushan, 1989), as they tend to be more mature and successful firms which tend to benefit from effective mechanisms to provide information disclosure to investors (see, e.g., Diamond and Verrecchia, 1991). Should this be the case, then our indicators of information asymmetries in the first three columns prove to do a good job in picking up most of the effects on listing. This is also consistent with Mayhew and Mihov's (2004) result that firm size was an important determinant of listing through the 1980s, but became unimportant in the 1990s (our data start in 1996, when Mayhew and Mihov's end), and with the findings in Danielsen *et al.* (2007) that the market value of equity has a negative relation with the likelihood of option listing, consistent with smaller stocks being favoured as optionable candidates.

Moreover, to check whether the results are sensible to details of our logistic regression framework, we also use a matched sample analysis of option listings, with the objective of testing through a different econometric approach the positive relationship between asymmetric information and the success of option listings. To this purpose, we match the quartile of option listings for which there is evidence of strong information asymmetries on the underlying stock in the year prior to listing, with the quartile of option listings that have low prior levels of asymmetric information. In practice, we follow Easley *et al.* (1998b) and select pairs of option listings that differ in prior underlying asymmetric information as much as possible but that are as similar as possible according to other matching criteria. Such a control sample methodology is designed to correct for the endogeneity of option listing. To achieve the goal of maximizing the difference in the informational dimension, the pairs are selected in the upper and lower quartiles of the cross-sectional distribution of *PIN*, *AdjPIN*, and *InvAnlst*, respectively, computed as usual with reference to the year prior to the listing date. Three matching criteria are simultaneously used: the industry to which the stock/firm belongs; the average daily stock dollar-volume over the 252 trading days before listing; and the volatility of the underlying stock returns estimated as the annualized daily standard deviation over the 252 trading days prior to listing.[19] Once we have built the two matched samples, we apply the paired-sample sign test and the Wilcoxon signed-rank test to option adoption rates measured as option dollar-volume, contract volume, and open interest in the year *following* option listings. Therefore, if high information asymmetries prior to listing positively affect the listing success, the upper quartile of listings ranked by *PIN*, *AdjPIN*, and *InvAnlst* (i.e., the quartile with options with the strongest prior asymmetric information) should display the highest values of the adoption indices in the first year of the option "life".

In Table 3, Panels A and B, the matched pairs from the lower and upper quartiles of option listings ranked by *PIN* and *AdjPIN* are used to test the differences in option success. The null hypothesis of no differences in adoption is tested against the one-sided alternative of higher adoption rates in the upper quartile vs. the lower quartile. Table 3, Panel A (Panel B) shows that out of 152 (141) pairs of matched listings, 95 (89), 92 (92), and 93 (86) of the pairs have larger levels of dollar-volume, contract volume, and open interest, respectively, in the group of listings characterized by elevated prior information asymmetries.[20] In addition, in Panel A (Panel B), the median of the percentage differences

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[19] For instance, suppose we want to match an option listing i with an option listing j. Option listings i and j have underlying stock dollar-volumes $DVlm_{S,i}$ and $DVlm_{S,j}$, and underlying stock volatilities $SDev_{S,i}$ and $SDev_{S,j}$, respectively. Given a listing i, when possible our algorithm selects j in the same industry as i by further minimizing the sum of the absolute relative differences between stock volumes and return volatilities, i.e., $\min_j[|(DVlm_{S,i} - DVlm_{S,j})/ DVlm_{S,i}| + |(SDev_{S,I} - Sdev_{S,j})/SDev_{S,i}|]$. Later in this section we also match firms on the basis of their size.

[20] The number of option listings in the lower and upper quartiles is 222 (891/4 = 222.75). However the number of pairs matched from the lower and upper quartiles is less than 222 because the matching criteria

in option dollar-volume, contract volume, and open interest between listings from the upper vs. the lower quartiles are 152.1% (200.7%), 103.0% (186.7%), and 163.3% (238.4%), respectively. These are very large disparity. The null hypothesis of no differences can be rejected at the 1% level using both tests (paired-sample sign and Wilcoxon signed-rank tests), both measures of asymmetric information, and all adoption measures. Consequently, the rejection of the null hypothesis of equal adoption rates independently of the strength of the underlying information asymmetries in Table 3, Panels A and B, implies that we can also reject the null that prior information asymmetries do not affect the success of option listings.

[Insert Table 3 here]

Table 3, Panel C shows results similar to those in Panels A and B, although in this case the lower and upper quartiles are selected sorting the cross-section of listings on the basis of the inverse of the number of analysts following the underlying stocks. In Panel C, the null hypothesis of no difference in the adoption rates is tested against the one-sided alternative of faster adoption in the upper quartile group than in the lower quartile. We find that out of 146 pairs of matched listings 99, 97, and 96 of the pairs have higher dollar-volume, contract volume, and open interest in the upper quartile group than in the lower quartile. The null hypothesis of no difference in the rate of listing success is also rejected at the 1% level using both tests and all adoption measures.

However, because our analysis has focussed on listings using as its base date the official start of negotiations, one may object that Table 3 may be affected by considerable non-synchronicity, in the sense that many or even most of the pairs of listings may implicitly compare options that have been newly introduced in the market at very different

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do not always allow perfect matching. The main matching constraint is that the two extreme quartiles contain different numbers of listings in different industries, and therefore this criterion reduces the number of pairs of listings that we are able to analyze. In addition, we impose a constraint by which the maximum absolute total relative difference used in the matching process cannot exceed 40% for the stock dollar-volume and return volatilities, i.e., we prevent poorly matching pairs to be formed.

times. Therefore we have reproduced in Table 4 a similar matched sample analysis as in Table 3, but this time making sure that the matched pairs are not separated in calendar time by more than 252 trading days.[21] Table 4 shows that the effects of prior information asymmetries on option listing success are even stronger after including this additional time-window constraint when building the matched pairs. The null hypothesis of no differences in adoption rates is again rejected at the 1% level, against the one-sided alternative of finding higher rates in the upper quartile group using both tests, the three measures of asymmetric information, and all adoption measures.

[Insert Table 4 here]

Roll *et al.* (2009) have reported a strong correlation between firm size and option trading intensity, regardless of the recent listing of options. It may therefore be surprising that firm size may not affect adoption success in our regressions in Table 2. To gain a deeper insight on the role played by size, in Table C1 we have repeated (collected in Appendix C, available in the supplementary material) the exercise in Table 3 when the matching reflects an additional constraint by which matched pairs must both have similar market capitalization calculated with reference to the year prior to listing. This ensures that the effects of information asymmetries on ex-post option adoption are measured without confounding effects coming from the very different size and popularity of different firms. The results we find are qualitatively similar to those in Table 3. For instance, when we form quantiles on the basis of *AdjPIN*, out of 129 pairs of matched listings, 80, 83, and 81 of the pairs have larger levels of dollar-volume, contract volume, and open interest, respectively, in the group of listings characterized by elevated prior information asymmetries; the median of the percentage differences in option dollar-volume, contract volume, and open interest between listings from the upper vs. the lower quartiles are

[21] For example, suppose we match an option listing *i* with an option listing *j*. We use the same matching criteria as in Table 3, but if option *i* is listed on day *t*, option *j* is now required to have been listed over the period [t-252, t+252] for the pair to be matched.

195%, 192%, and 210%, respectively. The null hypothesis of no differences (i.e., of a zero median change) can be rejected at the 1% level using both tests (paired-sample sign and Wilcoxon signed-rank tests), both measures of asymmetric information, and all adoption measures.

Because our 1996-2009 sample includes rather heterogeneous stock market phases and in particular the tech bubble (1997-2001) and the financial crisis (2008-2009) periods, and markets have manifested a rather anomalous behaviour during these periods, we also repeated the analysis in Table 3 excluding these years (i.e., only using data for the residual 7 years). On the one hand, this imposes a cost in terms of reducing the number of observations. However, if our matching methodology in Table 3 were unable to pick any differential dynamics over time, such an additional exercise would reveal them. Table C2 (collected in Appendix C) show results that confirm our hypothesis 1. For instance, when we form quantiles on the basis of *AdjPIN*, out of 58 pairs of matched listings, 39, 38, and 40 of the pairs have larger levels of dollar-volume, contract volume, and open interest, respectively, in the group of listings characterized by elevated prior information asymmetries; the median of the percentage differences in option dollar-volume, contract volume, and open interest between listings from the upper vs. the lower quartiles are 188%, 147%, and 140%, respectively. The null hypothesis of no differences can be rejected at the 1% level using both tests, both measures of asymmetric information, and all adoption measures.

The results in Tables 3 and 4 and these supplementary tests are all consistent with the empirical evidence presented in Table 2 that stronger information asymmetries characterizing the underlying stock predict a greater success—as measured by the strength of adoption—for any options written on the same stock, when these are eventually listed. This is consistent with our key hypothesis 1. An implication of these findings for the actual, *ex-post* dynamics of option trading and open interest is that option exchanges should also take prior asymmetric information measures into account when *ex-ante* making listing decisions.

5. Relative Bid-Ask Spreads and Informed Trading Activity in Option Markets after New Listings

In our analyses we have found that that the *option* relative bid-ask spread, $BAre_{OP}$, starts at low initial levels but subsequently displays a tendency to progressively increase.[22] This pattern of the $BAre_{OP}$ is presented in Figure 1, which shows the evolution of the cross-sectional daily average (for the complete sample of all listings) of $BAre_{OP}$ and the option dollar-volume in each month over the first year after listing. The low starting level of $BAre_{OP}$ is particularly interesting because in the early "life" of a (set of) option contracts these are normally characterized by substantial illiquidity, as one would expect of all newly created securities. Therefore, high and not low $BAre_{OP}$ values should be expected because the inverse of the relative bid-ask spread is a standard liquidity proxy in the literature (see e.g., Amihud and Mendelson, 1986).

[Insert Figure 1 here]

We conjecture that the initial low levels of $BAre_{OP}$ can be explained by a modest level of informed option trading activity in the early stages after option listings. On the one hand, even though it is well known that bid-ask spreads contain an inventory/liquidity component, the spread also reflects a component caused by the possibility of informed traders inflicting losses to market makers (see e.g., Copeland and Galai, 1983; Glosten and

[22] The relative bid-ask spread (*BAre*) is defined as *BAre* = (Ask Price – Bid Price)/(0.5(Ask Price + Bid Price)). Danielsen et al. (2007) report that in an ex ante perspective, stocks would be optioned when they have a high liquidity as measured also by low relative bid-ask spread. However our empirical result here concerns the dynamics of options' relative bid-ask spreads, not the underlying stocks'. However, if we were to assume that liquidity in the two markets shares a common component (as shown, for instance, by Kaul *et al.*, 2004), Danielsen et al. find that the market quality of the stock is not guaranteed to improve after option listing, which is consistent with our findings for the dynamics of *BAre* in the options market.

Milgrom, 1985); such a component is therefore an increasing function of any information asymmetries. This asymmetric information component has been characterized in models that capture the adverse selection problem faced by market makers.[23] On the other hand, immediately after the listing date, option contracts tend to be characterized by moderate volume at best, which only slowly grows over time (see Figure 1). Therefore, in the early stages after a listing, informed trades could be easily detected. As a result, rational informed investors are deterred from trying to hide their trades by fragmenting them and decide instead to optimally wait for volumes to pick up. Consequently, one may indeed see $BAre_{OP}$ start out low and progressively increase as volumes—hence the possibility to fragment informed trades in sequences of smaller, hard-to-detect orders as characterized by Anand and Chakravarty (2007)—pick up.

To corroborate this conjecture, we use a two-pronged strategy, in Tables 5 and 6. In a first step, we analyze matched samples to test the significance of the growth in $Bare_{OP}$ over time (see Table 5, Panel A), using the following logic. We calculate equally-weighted (as in Easley *et al.*, 1998b, and Pan and Poteshman, 2006) cross-sectional averages of relative bid-ask spreads with reference to both the first and the thirteenth months after listing, matched in pairs of $BAre_{OP}$ values. Because the calculation of spreads across different option classes requires a weighting scheme across series. We later test the robustness of our conclusions for at-the-money options only. Paired-sample sign and the Wilcoxon signed-rank tests are then applied to test whether relative bid-ask spreads increase significantly between the first month and the initial thirteenth months over the option 'life'. The null hypothesis of no change in $BAre_{OP}$ is tested against the one-sided alternative of an increase in $BAre_{OP}$. Table 5 Panel A, shows that out of the 891 option

[23] The *PIN* and the *AdjPIN* estimates obtained from underlying stock data using the microstructure model in Appendix A, available in the supplementary material, rely exactly on these arguments. An alternative proxy for the probability of informed trading could be given by *PIN* or *AdjPIN* estimated from intra-day option data. However, it would be difficult to obtain reliable *PIN* and *AdjPIN* estimates based on options data because of the modest trading activity that usually characterizes the immediate post-listing periods.

listings in our sample, 689 display a widening $BAre_{OP}$ when one compares the thirteenth- with the first-month after listing; the median increase is a stunning 40.59%, with the median relative bid ask spread growing from 0.32 just after listing to 0.44 after 13 months from listing. The null hypothesis of no difference is formally rejected at 1% under both types of tests, i.e., paired-sample sign and the Wilcoxon signed-rank tests.

[Insert Table 5 here]

Furthermore, to check for potential biases in the results in Table 5 due to variations in market conditions after listings, we use a further control group of 891 equity options with at least three years of market activity after their listing dates. The spirit of the test is to verify that long after the listing, the relative bid ask spread stabilizes as the flow of access of informed traders to the options market does (or, alternatively, as information asymmetries disappear, see section 6 and hypothesis 3). Therefore each of the newly listed contract series in Table 5 is matched to a "seasoned" stock option. The matching is performed using the same criteria used in Tables 3 and 4: the underlying stock industry, the underlying stock return volatility, and underlying stock volume. In Table 5 Panel B, no significant increasing (or decreasing) patterns for option relative bid-ask spreads are found for the control sample of seasoned listings. In fact, the change in $BAre_{OP}$ between the first and thirteenth month following the listing date for the baseline sample of listings (Table 5, Panel A) is strongly significantly larger than the percentage change observed for the seasoned controls (data from Table 5, Panel B), using both paired-sample sign and the Wilcoxon signed-rank tests. In addition, in an unreported test, we verify that the percent change in $BAre_{OP}$ for option listings in Tables 5 Panel A is significantly higher than the changes in $BAre_{OP}$ for the control group (Tables 5 Panel B).

The calculation of average and median spreads requires a weighting scheme across option series and in Table 5 we have used a simple but intuitive equally-weighted scheme.

In particular, controlling for the fact that additional option series may be added over time, thus increasing the equal weighted spread may pose problems. To avoid these issues, in Table C3 we have repeated (see Appendix C, available in the supplementary material) the exercise in Table 5 when the relative bid-ask spread is calculated using only at-the-money (*ATM*) one-month to maturity option contracts, which are also traditionally considered the most liquid and actively traded option class.[24] The results in Table 5 are confirmed by this additional exercise: for instance, out of 891 *ATM* listings, 543 display a widening $BAre_{OP}$ when one compares the thirteenth- with the first-month after listing; the median increase is a stunning 31.39%, with the median relative bid ask spread growing from 0.18 just after listing to 0.23 after 13 months from option listing. The null hypothesis of no difference is formally rejected at 1%. Of course, these relative spreads are much smaller than those reported in Table 5 because they concern highly liquid and traded *ATM* contracts. Moreover, as in panel B of Table 5, the increase in spreads stops occurring in the time intervals marked by months 14-36.

Additionally, we perform a matched sample analysis to evaluate the impact of any fragmentation strategies implemented by informed option traders on the changes in relative bid-ask spread ($\Delta Bare$). In Table 6, we match the group of option listings for which we have evidence of strong information asymmetries on the underlying asset in the year prior to option listings with listings that have instead low prior levels of asymmetric information (i.e., listings in the upper vs. lower quartiles of *PIN*, *AdjPIN*, and *InvAnlst*, respectively). We use the same matching criteria and apply to the resulting sample of matched pairs sign- and signed-rank tests to test the null of no change in $\Delta Bare$. The intuition is that only when an option listing is successful, informed investors may progressively ripe the advantages of growing volumes to hide their trades. However, this is

[24] Given that there are not always exactly at-the-money one-month option contracts, we calculate *BAre* by simple linear interpolation using the four contracts around the one-month time-to-maturity and with closest strike price to the underlying asset.

exactly what a rational market maker would anticipate, thus ending up widening the relative bid-ask spreads over time.

[Insert Table 6 here]

Table 6 shows that the listings characterized by previously high levels of information asymmetries imply larger changes in the option relative bid-ask spread than the listings with low prior asymmetric information. For instance, Table 6, Panel A, shows that out of 152 pairs of matched option listings, 89 of the pairs have larger *ΔBare* values in the quartile characterized by elevated prior information asymmetries, as ranked by *PIN*. The null hypothesis of no differences in *ΔBare* can be rejected with p-values always inferior to 5%, using both tests and all measures of asymmetric information. The results reported in Table 6 are consistent with our conjecture that informed option traders would wait for sufficient volume to flood the market before implementing their typical stealth strategies. Hence option listings concerning stocks plagued by strong information differentials would only eventually attract a large number of informed option traders who however would avoid trading in the immediate aftermath of listing. In Table C4 (in Appendix C, available in the supplementary material) we have also implemented these tests with reference to *ATM* contracts only and find qualitatively similar results, thus ruling out that it may be our equal-weighting scheme the main driver of results concerning the dynamics of bid-ask spreads over time.Finally, in Table C5 (in Appendix C) we have repeated the same analysis as in Table 6 when we include an additional constraint by which matched pairs are formed under the additional restriction of the any pair of listings having occurred within 252 trading days (similarly to the time-window constraint already applied in Table 4). The goal of this further restriction is to prevent results to be driven by paired samples characterized by large disparities in the calendar dates of the listings. The null hypothesis of no difference in *ΔBare* is again rejected with p-values well below a 5% size, against the one-sided alternative of larger $\Delta Bare$ values in the upper vs. the bottom quartiles, and using both tests and all measures of asymmetric information.

6. The Effect of Option Listings on Information Asymmetries

As already conjectured in section 2, the ex-post dynamics triggered by the introduction of new option may eventually record a "happy ending". In spite of the fact that optioned stocks characterized by stronger information asymmetries score the greater success and that this may initially happen at the expenses of the liquidity in the derivatives market, the introduction of option-style derivative contracts end up improving market quality, in the sense that markets may become increasingly informationally efficient. In our empirical tests, we find that information asymmetries *in the underlying stock market* decline after options are newly listed. Such a decline derives from a learning process by the uninformed agents that are now able to exploit the visibility of the trades by the informed investors in options to infer their private information.[25] Moreover, after an option listing, it is customary that multiple standardized option contracts on the same underlying asset are simultaneously introduced (for instance, along the strike and maturity dimensions), so that there are strong incentives for all investors towards collecting increasing amounts of information concerning the underlying asset's payoffs (see e.g., Cao, 1999; Massa, 2002). Therefore, option listings ought to induce an increase in the number of market analysts following the underlying stock (e.g., Skinner, 1990) as well as a decline in the fraction of trades that are backed by private information, as the latter asymptotically vanishes (or, its profitability declines) in a market in which option trades contribute to the efficient dissemination of company news. In fact, the tendency for the number of analysts to

[25] The main cognitive mechanism followed by uninformed investors may consist of a learning-by-observing process which assumes that agents do not live in an isolated environment and in which, on the opposite, their surroundings represent a source of additional knowledge (e.g., see Bikhchandani *et al.*, 1998; DeLong and DeYoung, 2007).

increase is also instrumental to further reductions in information asymmetries, as the scrutiny of news and trades offered by skilled professional facilitates the detection of any private information. This process of declining in information asymmetries may be presumed to be driven by options markets, because the literature has emphasized that it may be easier to uninformed investors to learn from informed trading when there are option securities than without them. For instance, the high leverage that characterizes option markets and the corresponding low margin requirements are particularly attractive to informed investors (see e.g., Back, 1992; Biais and Hillion, 1994; Pan and Poteshman, 2006), thus many empirical papers (e.g., Danielsen and Sorescu, 2001) have concluded that option trades affect the asymmetric nature of information stocks and flows also in the equity market, and additionally Cao and Wei (2010) find evidence that information asymmetry is greater for options than for the underlying stock (implying that agents with information find the options market a more efficient venue for trading).

Similarly to section 4, we use *PIN* and *AdjPIN* to analyze any changes in information asymmetries that follow option listings. For each newly listed option, we compare two estimates of *PIN* and *AdjPIN*: the values estimated over the year that precedes the listing date (PIN_{0Y} and $AdjPIN_{0Y}$) and the values estimated over the year following the listing (PIN_{1Y} and $AdjPIN_{1Y}$). The paired-sample sign and the Wilcoxon signed-rank tests are then applied to the different *PIN* and *AdjPIN* estimates. In Tables 7 and 8, the null hypothesis of no change in the asymmetry measures is tested against the one-sided alternative of a decrease after the listing. The two tests are applied to a few alternative sub-samples (in descending order in the table): the complete sample of listings; the listings in the lower and upper quartiles as sorted by $DVlm_{OP,1Y}$; the listings in the lower and upper quartiles by $Vlm_{OP,1Y}$; and listings in the lower and upper quartiles by $OInt_{OP,1Y}$. Reporting results sorted by quartiles of measures of adoption helps to test whether it is the newly admitted trading in options that causes the decline in information asymmetries. If that is the case, the decline in the latter should be stronger the higher is the success of a listing, i.e., for listings in the upper quartiles of $DVlm_{OP,1Y}$, $Vlm_{OP,1Y}$, and $OInt_{OP,1Y}$. Table 7 concerns *PIN* while Table 8 concerns *AdjPIN*. In both tables, we find that information asymmetries concerning optioned stocks substantially decline after options are listed. This means that the data fail to reject our hypothesis 3 from section 2. For instance, Table 7, first row (Table 8, first row) shows that out of the 891 option listings in our complete sample, 532 (558) have lower *PIN* (*AdjPIN*) values after the listing date, with a median percentage change of -22.5% (-20.0%). Moreover, in both cases the null hypothesis of no differences in *PIN* and *AdjPIN* estimates before and after option listings is rejected with p-values smaller than 5% using paired-sample sign and Wilcoxon signed-rank tests and for all quartiles.

[Insert Table 7 here]

[Insert Table 8 here]

However, in Tables 7 and 8, we also note that—as expected—the listings in the higher quartiles of success (independently of how this is measured) imply larger effects compared to option listings in the lower quartiles. For instance, in Table 8, last two panels, *AdjPIN* declines by a median 14.9% in the case of the bottom post-listing open interest quartile vs. a much larger 30.6% decline in the case of the upper post-listing open interest quartile, which is more than double. The higher impact recorded in the upper quartile listings are consistent with a learning-by-observing hypothesis à la DeLong and DeYoung (2007): in the upper quartiles built sorting by adoption rates, market activity following a listing is, by construction, stronger than for other quartiles, so that large amounts of private information may be revealed through trading in newly listed options. This is related to recent findings by Roll *et al.* (2009) concerning the trading of seasoned options: they report that "liquidity-attracts-liquidity" so that highly traded options increase firm value because, besides completing markets, they stimulate informed trades and therefore informational efficiency.

Furthermore, to check for potential biases in the results in Tables 7 and 8 due to variations in market conditions after listings are decided, we use a further control group of 891 equity options with at least three years of market activity after their listing dates. The spirit of the test is to verify that long after the listing, the reduction effects on measurable information asymmetries become negligible, or at least weak, which is also consistent with the learning-by-observing hypothesis: initial effects from learning from the trading environment ought to be stronger than steady-state effects. Therefore each of the newly listed contract series in Tables 7 and 8 is now matched to a "seasoned" stock option. The matching is performed using the same criteria used in Tables 3 and 4.

Tables 9 and 10 reproduce the analysis in Tables 7 and 8 using the control group of seasoned options. Table 9 and 10 show that there are no significant reductions in information asymmetries concerning the underlying stocks for options in this control group. For example, Table 9, first row (Table 10, first row) shows that out of 891 equity options in the complete control group, 458 (463) have lower *PIN* values (*AdjPIN* values) after the listing date but with a median change of only -1.35% (-1.59%). The null hypothesis of no changes in *PIN* and *AdjPIN* after the introduction date cannot be rejected at conventional significance levels for all sub-groups based on quartiles and using both parametric and nonparametric tests. In addition, in unreported tests, we verify that the percentage change in *PIN* and *AdjPIN* of matched listings in Tables 7 and 8 over the year prior to and the year following listing are significantly more negative than the same percentage changes in *PIN* and *AdjPIN* of their controls (Tables 9 and 10), also in this case using tests at a 5% size. Therefore, Tables 9 and 10 strengthen earlier evidence that changes in information asymmetries are fundamentally related to option listings and not to the mere fact that options are traded, independently of their recent introduction.

[Insert Table 9 here]

[Insert Table 10 here]

Additionally, similarly to Damodaran and Lim (1991) and Skinner (1990), we find that the number of analysts significantly increases after option listings, which is the evidence presented in Table 11. In Table 11, the paired-sample sign and the Wilcoxon signed-rank tests are applied to matched pairs based on the inverse of the number of analysts following (i.e., publishing earnings forecasts) on average the stock during the year prior ($InvAnlst_{0Y}$) and following ($InvAnlst_{1Y}$) the listing date. The null hypothesis of no difference in the (inverse of the) number of analysts is tested against the one-sided alternative of a decline (which would imply that $InvAnlst_{1Y}$ increases over $InvAnlst_{0Y}$). Table 11 shows that out of 891 listings in the complete sample, 610 imply a lower value for *InvAnlst* (i.e., more analysts follow the stock) after the listing, with a median percentage change of -22.4%. The null hypothesis of no differences for all the quartile sub-samples is always rejected at the 1% level using both tests. Like Tables 7 and 8, also Table 11 implies that information asymmetries abate after options are listed, both directly if we take $InvAnlst_{Y}$ as an index of such asymmetries and indirectly, as the growth in the number of analysts producing public news reports has been shown to help tame any information differences (see Tetlock, 2010). Also in Table 11, such an increase in the number of analysts is stronger the more successful listings are.

[Insert Table 11 here]

Finally, we have repeated the same analysis as in Table 11, but using our control group of seasoned equity options already defined for the purpose of computing Tables 9 and 10. Table C6 in Appendix C emphasizes once more that there is no significant increase in the number of analysts of seasoned equity options long after the initial listing date. Table C6 shows that out of 891 matched, seasoned listings in our complete sample, 461 have a lower level in the *InvAnlst* after the listing date, but with a median change of a puny −1.77%. The null hypothesis of no differences in *InvAnlst* cannot be rejected for all quartiles sub-samples using any of the tests. Moreover, in additional analyses, we observe that the percentage change in the inverse number of analysts of matched option listings (statistics from Table 11) between the year prior to and the year following the listing is significantly more negative than the same value for their controls (statistics from Table C6) at a 1% level.

7. Conclusions

Option listings represent one often seen (one would say, routine) case of financial innovation in which completely new derivatives securities are introduced into the market for the first time by option exchanges. Consequently, understanding the option adoption process has enormous importance to both policy (i.e., normative) and positive perspectives, also because a deeper insight into the process may lead to a better understanding of the dynamics of the success and/or failure of even more sophisticated derivative securities. In addition, knowledge of the factors that affect the success of newly listed options is of extreme relevance to option exchanges, which are in charge of selecting optionable stocks and are likely to do so with a view to long-run profit maximization.

Differently from earlier literature that has focussed on the *ex-ante* decision (by option exchanges) to select stocks for listing option contracts, our study has examined the determinants of the actual, *ex-post* success of option listings, and particularly the role that asymmetric information plays in affecting the adoption process. We use data from option listings on the U.S. equity option markets over a relatively long period of time, 1996-2009. Our first and crucial result is that, using different proxies for information asymmetries common to the microstructure literature, high information asymmetries affecting the underlying stock prior to listing result in an ex-post higher rate of adoption. Importantly, information asymmetry measures remain a key predictor of new options' success, even after controlling for factors that the earlier literature indicated as responsible for the exchanges' choice to option certain stocks and not others.

We also find that option listings reduce asymmetric information. Moreover, we obtain empirical evidence to support a view in which informed option traders slowly enter the newly created markets because of their need to exploit sufficient volumes to "hide" their informed trades. As a result—because it reflects the probability that market makers perceive of dealing with informed investors—the option relative bid-ask spread is observed to be initially low and to progressively increase as an option market takes off. This counterintuitive result (one would naively expect bid-ask spreads to narrow as volume picks up) appears to be stronger for options written on stocks characterized by large information differentials, which are however also the option listings that are mostly likely to be successful.

Finally, the econometric approach used in our paper is simple and intuitive because we limit ourselves to use standard regression analysis and, even more frequently, matched-pair tests of differences in medians for the quantities of interest. However, it is clear that more sophisticated and (possibly) more powerful econometric techniques may allow us to expand our study to other issues that remain to be addressed. For instance, apart from the need to generalize our results to other, more complex (such as over-the-counter) derivatives, an exploration of whether there are windows of opportunity for exchanges to optimally time the introduction of new option contracts on the basis of the underlying asymmetries in information has been left for future research.

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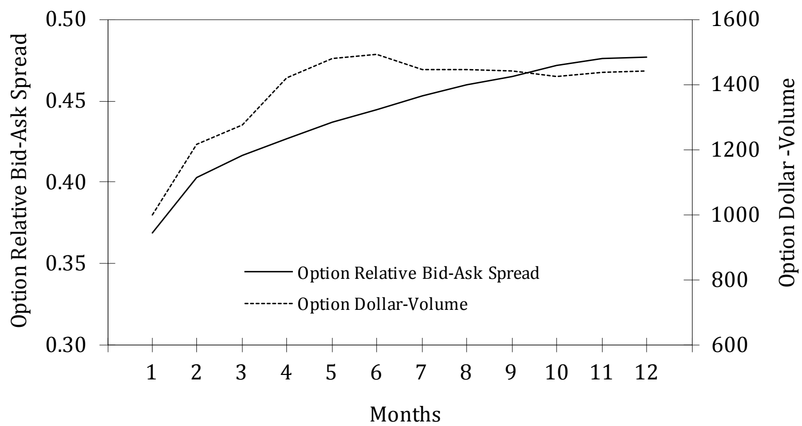

Figure 1. Evolution of the option dollar-volume and the option relative bid-ask spread. The figure presents the evolution of the cross-sectional mean in each month of the average for the daily option dollar-volume and the relative option bid-ask spread in the 12 months following the listing date. The relative bid-ask spread (*BAre*) is defined as *BAre* = (*Ask Price* – *Bid Price*)/(0.5(*Ask Price* + *Bid Price*)).

Table 1
Summary Statistics

The table contains cross-sectional statistics of the main variables used in the study. $DVlm_{OP,1Y}$, $Vlm_{OP,1Y}$, and $OInt_{OP,1Y}$ are the averages of the daily option dollar-volume, option contract volume, and open interest, respectively, in the first year after the option listing. $BAre_{OP,1Y}$ is the average of the option relative bid-ask spread in the year following the listing date, where *BAre* = (*Ask Price* – *Bid Price*)/(0.5(*Ask Price* + *Bid Price*)). PIN_{0Y}, $AdjPIN_{0Y}$, and $InvAnlst_{0Y}$ are the *PIN* and *AdjPIN* estimates, and the inverse of the average of the number of analysts, respectively, for the year prior to option listing. PIN_{1Y}, $AdjPIN_{1Y}$, and $InvAnlst_{1Y}$ are the *PIN* and A*djPIN* estimates, and the inverse function of the average of the number of analysts, respectively, for the year immediately after option listing.

Variable	Mean	Median	Std. Dev.	Skewness	Kurtosis	Min.	Max.	Obs.
$DVlm_{OP,\,1Y}$	1335.79	124.44	31064.85	35.57	832.56	0.28	9012509.12	891
$Vlm_{OP,\,1Y}$	498.72	74.40	7212.01	29.51	880.45	0.38	214713.33	891
$OInt_{OP,\,1Y}$	8835.03	1833.14	108539.05	29.41	876.35	25.51	3228950.54	891
$BAre_{OP,\,1Y}$	0.44	0.42	0.15	0.69	3.81	0.11	1.11	891
PIN_{0Y}	0.20	0.19	0.06	0.43	2.91	0.08	0.37	891
PIN_{1Y}	0.15	0.14	0.04	0.84	3.98	0.08	0.27	891
$AdjPIN_{0Y}$	0.17	0.16	0.05	0.24	4.37	0.02	0.33	891
$AdjPIN_{1Y}$	0.13	0.13	0.03	0.56	5.42	0.03	0.23	891
$InvAnlst_{0Y}$	0.33	0.27	0.23	1.31	4.25	0.04	1.00	891
$InvAnlst_{1Y}$	0.25	0.20	0.17	2.37	9.86	0.04	1.00	891

Table 2

Regression Analysis of the Impact of Different Factors on Ex-Post Option Adoption Rates

The table reports regressions of measures of success of new and recently listed stock options on a range of explanatory factors. Panel A, B, and C present the estimated coefficients of equations (14a), (14b), and (14c), respectively. $Ln(\cdot)$ is the natural logarithmic function. $DVlm_{OP,1Y}$,$Vlm_{OP,1Y}$, $OInt_{OP,1Y}$, PIN_{0Y}, $AdjPIN_{0Y}$, and $InvAnlst_{0Y}$ are defined in Table 1. Since PIN, $AdjPIN$, and $InvAnlst$ range between zero and one, their estimates are logistically transformed before being used as explanatory variables in the regressions. The other variables used are: underlying stock volume, distinguishing between long-term ($DVlm_{S,252,0Y}$) and a short-term ($DVlm_{S,21,0Y}$) components, which are calculated as the average daily stock dollar-volume using the 252 and 21 trading days preceding the listing date, respectively; the underlying stock return volatility, distinguishing between long-term ($SDev_{S,252,0Y}$) and short-term ($SDev_{S,21,0Y}$) components, calculated as the annualized standard deviation of daily log returns over the 252 and 21 trading days preceding the listing date, respectively; and the distinguishing between stock market capitalization ($Size_{0Y}$) calculated with reference to the year to listing. ***, **, and * denote significance at 1%, 5%, and 10%, respectively (t-statistics are in parentheses).

PIN_{0Y}	$AdjPIN_{0Y}$	$InvAnlst_{0Y}$	$Ln\ (DVlm_{S,252,0Y})$	$DVlm_{S,21,0Y}/DVlm_{S,252,0Y}$	$SDev_{S,252,0Y}$	$SDev_{S,21,0Y}/SDev_{S,252,0Y}$	$Ln\ (Size_{0Y})$	Const.	Obs.	R^2
Panel A: Dependent Variable $Ln\ (DVlm_{OP,1Y})$										
0.45			0.91	0.40	1.29	-0.26	-0.22	-9.28	891	0.41
(2.02)**			(11.73)***	(10.76)***	(3.91)***	(0.97)	(1.03)	(9.83)***		
	0.49		0.83	0.41	1.18	-0.31	-0.15	-10.24	891	0.43
	(2.72)***		(11.83)***	(11.23)***	(3.16)***	(2.78)	(0.77)	(11.31)***		
		0.83	1.04	0.48	0.71	-0.46	-0.08	-11.51	891	0.48
		(3.16)***	(12.16)***	(11.02)***	(3.63)***	(1.81)*	(0.85)	(11.20)***		
Panel B: Dependent Variable $Ln\ (Vlm_{OP,1Y})$										
0.43			0.85	0.43	0.22	-0.39	-0.25	-7.01	891	0.37
(2.51)**			(11.24)***	(9.02)***	(2.63)***	(0.85)	(1.54)	(6.28)***		
	0.49		0.71	0.35	0.12	-0.33	-0.12	-8.35	891	0.42
	(2.88)***		(10.66)***	(10.04)***	(3.21)***	(0.92)	(1.45)	(8.24)***		
		0.81	0.86	0.41	0.19	-0.50	-0.16	-7.74	891	0.36
		(3.21)***	(10.38)***	(9.74)***	(3.39)***	(1.39)	(1.84)*	(7.82)***		
Panel C: Dependent Variable $Ln\ (OInt_{OP,1Y})$										
0.33			0.69	0.35	0.04	-0.35	-0.20	-2.64	891	0.31
(2.66)***			(9.16)***	(10.47)***	(2.52)**	(0.32)	(1.18)	(3.95)***		
	0.32		0.62	0.29	0.07	-0.32	-0.17	-3.04	891	0.33
	(2.94)***		(9.58)***	(8.82)***	(2.76)***	(0.55)	(1.33)	(3.21)***		
		0.79	0.78	0.36	0.05	-0.43	-0.13	-3.53	891	0.35
		(3.44)***	(10.08)***	(9.45)***	(3.28)***	(0.41)	(1.68)*	(3.91)***		

Table 3

Matched Sample Analysis of Option Listings to Assess the Effect of Asymmetric Information on Option Adoption Rates

The table reports a matched sample analysis of option listings in relation to the effects of prior information asymmetries on the adoption rate during the first year after listing. PIN_{0Y}, $AdjPIN_{0Y}$, $InvAnlst_{0Y}$, $DVlm_{OP,1Y}$, $Vlm_{OP,1Y}$, and $OInt_{OP,1Y}$ are defined in Table 1. In Panels A, B, and C the matched pairs consist of listings in the lower and upper quartiles using *PIN*, A*djPIN*, and *InvAnlst*, respectively to sort listings, all calculated with reference to the year prior to listing. The matching criteria for the pairs of listings are the underlying stock industry, the underlying stock return volatility, and the underlying stock dollar-volume. *OAdp* is the measure of option adoption using either $DVlm_{OP,1Y}$, $Vlm_{OP,1Y}$, or $OInt_{OP,1Y}$; while $OAdp_{LQ}$ ($OAdp_{UQ}$) is the measure of option adoption for the listings in the lower (upper) quartile ranked by measures of asymmetric information. The table reports both the paired-sample sign test and the Wilcoxon signed-rank test which are applied to pairs of option listings in relation to adoption levels. The null hypothesis of no difference in rates of option adoption is tested against the one-sided alternative of larger $OAdp_{UQ}$ values than $OAdp_{LQ}$. For instance, the first row of Panel shows that out of 152 pairs of matched listings, 95 have higher adoption levels using as proxy $DVlm_{OP,1Y}$ in the upper quartile than in the lower quartile of options ranked by *PIN* values in the year prior to the listing date. [a], [aa], and [a] denote significance at 1%, 5%, and 10%, respectively, for the paired-sample sign test. [bbb], [bb], and [b] indicate significance at 1%, 5%, and 10%, respectively, for the Wilcoxon signed-rank test.

OAdp	Obs.	#($OAdp_{LQ}$ < $OAdp_{UQ}$)	%($OAdp_{LQ}$ < $OAdp_{UQ}$)	Median $OAdp_{LQ}$	Median $OAdp_{UQ}$	Median % Change ($OAdp_{UQ}/OAdp_{LQ}$-1)
	Panel A: Differences in Option Adoption (Year Following Opt. Listings) between the Groups in the Lower and Upper Quartiles by the PIN_{0Y} (Year Prior to Opt. Listings)					
$DVlm_{OP,1Y}$	152	95	62.50%	72.83	172.80	152.05%[aaa,bbb]
$Vlm_{OP,1Y}$	152	92	60.53%	49.76	104.28	102.99%[aaa,bbb]
$OInt_{OP,1Y}$	152	93	61.18%	1324.41	2796.76	163.29%[aaa,bbb]
	Panel B: Differences in Option Adoption (Year Following Opt. Listings) between the Groups in the Lower and Upper Quartiles by the $AdjPIN_{0Y}$ (Year Prior to Opt. Listings)					
$DVlm_{OP,1Y}$	141	89	63.12%	58.39	188.73	200.68%[aaa,bbb]
$Vlm_{OP,1Y}$	141	92	65.25%	43.78	113.39	186.69%[aaa,bbb]
$OInt_{OP,1Y}$	141	86	60.99%	1182.11	2953.81	238.43%[aaa,bbb]
	Panel C: Differences in Option Adoption (Year Following Opt. Listings) between the Groups in the Lower and Upper Quartiles by the $InvAnlst_{0Y}$ (Year Prior to Opt. Listings)					
$DVlm_{OP,1Y}$	146	99	67.81%	71.13	160.27	131.07%[aaa,bbb]
$Vlm_{OP,1Y}$	146	97	66.44%	51.47	95.16	121.06%[aaa,bbb]
$OInt_{OP,1Y}$	146	96	65.75%	1462.78	2294.50	89.54%[aaa,bbb]

Table 4

Matched Sample Analysis of Option Listings to Assess the Effect of Asymmetric Information on Option Adoption Rates under a Time-Window Constraint

The table reports a similar matched sample analysis as in Table 3, but in this case the matching reflects an additional constraint by which matched pairs must both belong to a time-window of 252 trading days from the baseline listing. For instance, in the case of the matching of a listing *i* with a listing *j*, we use the same matching criteria as in Table 3, but if the listing *i* has occurred on day *t*, the listing *j* must fall in the interval [t-252, t+252]. PIN_{0Y}, $AdjPIN_{0Y}$, $InvAnlst_{0Y}$, $DVlm_{OP,1Y}$, $Vlm_{OP,1Y}$, and $OInt_{OP,1Y}$ are defined in Table 1. In Panels A, B, and C the matched pairs consist of option listings in the lower and upper quartiles using *PIN*, A*djPIN*, and *InvAnlst*, respectively, all calculated in the year prior to listing. *OAdp* is the level of option adoption using either $DVlm_{OP,1Y}$, $Vlm_{OP,1Y}$, or $OInt_{OP,1Y}$; while $OAdp_{LQ}$ ($OAdp_{UQ}$) is the measure of option adoption for option listings in the lower (upper) quartile. The table reports both paired-sample sign test and Wilcoxon signed-rank test which are applied to the pairs of listings in relation to adoption levels. The null hypothesis of no difference in adoption rates is tested against the one-sided alternative of $OAdp_{UQ}$ exceeding $OAdp_{LQ}$. For instance, the first row of Panel A shows that out of 41 pairs of matched listings, 33 implied higher adoption rates using as a proxy $DVlm_{OP,1Y}$ in the upper than in the lower quartiles ranked by *PIN* values in the year prior to listing. [a], [aa], and [a] denote significance at 1%, 5%, and 10%, respectively, for the paired-sample sign test. [bbb], [bb], and [b] indicate significance at 1%, 5%, and 10%, respectively, for the Wilcoxon signed-rank test.

OAdp	Obs.	#($OAdp_{LQ}$ < $OAdp_{UQ}$)	%($OAdp_{LQ}$ < $OAdp_{UQ}$)	Median $OAdp_{LQ}$	Median $OAdp_{UQ}$	Median % Change ($OAdp_{UQ}/OAdp_{LQ}$-1)
	Panel A: Differences in Option Adoption (Year Following Opt. Listings) between the Groups in the Lower and Upper Quartiles by the PIN_{0Y} (Year Prior to Opt. Listings)					
$DVlm_{OP,1Y}$	41	33	80.49%	69.31	174.94	168.17%[aaa,bbb]
$Vlm_{OP,1Y}$	41	30	73.17%	51.36	107.63	111.49%[aaa,bbb]
$OInt_{OP,1Y}$	41	32	78.05%	1258.71	2900.44	172.91%[aaa,bbb]
	Panel C: Differences in Option Adoption (Year Following Opt. Listings) between the Groups in the Lower and Upper Quartiles by the $InvAnlst_{0Y}$ (Year Prior to Opt. Listings)					
$DVlm_{OP,1Y}$	38	29	76.32%	50.56	197.26	208.34%[aaa,bbb]
$Vlm_{OP,1Y}$	38	31	81.58%	40.21	118.40	193.83%[aaa,bbb]
$OInt_{OP,1Y}$	38	27	71.05%	1068.03	3262.49	245.48%[aaa,bbb]
	Panel C: Differences in Option Adoption (Year Following Opt. Listings) between the Groups in the Lower and Upper Quartiles by the $InvAnlst_{0Y}$ (Year Prior to Opt. Listings)					
$DVlm_{OP,1Y}$	37	33	89.19%	67.16	167.83	140.36%[aaa,bbb]
$Vlm_{OP,1Y}$	37	29	78.38%	47.48	98.53	129.33%[aaa,bbb]
$OInt_{OP,1Y}$	37	31	83.78%	1390.10	2340.01	95.82%[aaa,bbb]

Table 5

Increase in the Option Relative Bid-Ask Spread After Option Listings

The table presents a matched sample analysis of the effects of option listings on changes in the option relative bid-ask spread for the baseline sample of options (Panel A) and for a control group (panel B). The relative bid-ask spread (*BAre*) is defined as *BAre* = (*Ask Price* – *Bid Price*)/(0.5(*Ask Price* + *Bid Price*)). $BAre_{OP,1M}$ ($BAre_{OP,13M}$) is the average relative bid-ask spread in the first (thirteenth) month after listing. In Panel A, the matched pairs concern the same option listing but *BAre* is measured in different time periods (in the first and thirteenth months after listing). The table reports both the paired-sample sign test and the Wilcoxon signed-rank test applied to the change from $BAre_{OP,1M}$ to $BAre_{OP,13M}$. The null hypothesis of no change in $BAre_{OP}$ is tested against the one-sided alternative of $BAre_{OP,13M}$ being larger than $BAre_{OP,1M}$. For instance, the first row shows that out of 891 pairs of matched $BAre_{OP}$ values in the complete sample of option listings, 689 have larger $BAre_{OP}$ levels in the thirteenth month than in the first month. In Panel B, the control group is created by selecting a seasoned equity option (with at least three years of market activity subsequent to its listing) to match each listing in the original sample. The matching criteria are defined in Table 7. All the statistics in Panel B are calculated using data from the equity options in the control group and have structure similar to Panel A. [a], [aa], and [a] denote significance at 1%, 5%, and 10%, respectively, for the paired-sample sign test. [bbb], [bb], and [b] indicate significance at 1%, 5%, and 10%, respectively, for the Wilcoxon signed-rank test.

Diff. in Option Relative Bid-Ask Spreads between the First and Thirteenth Months Following Option Listings Using Multiple Sub-Samples

	Obs.	#($BAre_{OP,1M}$ < $BAre_{OP,13M}$)	%($BAre_{OP,1M}$ < $BAre_{OP,13M}$)	Median $BAre_{OP,1M}$	Median $BAre_{OP,13M}$	Median % Change ($BAre_{OP,13M}$/$BAre_{OP,1M}$-1)
	Panel A: Complete Sample of Equity Options					
$BAre_{OP}$	891	689	77.33%	0.32	0.44	40.59%[aaa,bbb]
	Panel B: Control Group of Equity Options					
$BAre_{OP}$	891	441	49.49%	0.47	0.46	-2.09%

Table 6

Matched Sample Analysis of the Impact of Options Listings on the Change in Option Relative Bid-Ask Spreads

The table presents a matched sample analysis of the effects of option listings and stealth strategies by informed traders on the change in the option relative bid-ask spread. $BAre_{OP,1M}$ ($BAre_{OP,13M}$) is the average of the option relative bid-ask spread in the first (thirteenth) month after option introduction. The change in the option relative bid-ask spread is defined as $\Delta Bare = BAre_{OP,13M}/BAre_{OP,1M.}$ PIN_{0Y}, $AdjPIN_{0Y}$, and $InvAnlst_{0Y}$ are defined in Table 1. In Panels A, B, and C the matched pairs consist of listings in the lower and upper quartiles built using *PIN*, A*djPIN*, and *InvAnlst*, respectively which are calculated in the year prior to the listing date. The matching criteria to form pairs of option listings are the underlying stock industry, the underlying stock return volatility, and the underlying stock dollar-volume. $\Delta Bare_{LQ}$ ($\Delta Bare_{UQ}$) is the $\Delta Bare$ for the option listings in the lower (upper) quartiles ranked by measures of asymmetric information. The table reports both the paired-sample sign test and the Wilcoxon signed-rank test applied to the change in the option relative bid-ask spread. The null hypothesis of no difference in $\Delta Bare$ is tested against the one-sided alternative of a positive difference between $\Delta Bare_{UQ}$ and $\Delta Bare_{LQ}$. For instance, the first row shows that out of 152 pairs of matched option listings, 89 have higher $\Delta Bare$ in the upper quartile than in the lower quartile of options ranked by *PIN* values in the year prior to the listing date. [a], [aa], and [a] denote significance at 1%, 5%, and 10%, respectively, for the paired-sample sign test. [bbb], [bb], and [b] indicate significance at 1%, 5%, and 10%, respectively, for the Wilcoxon signed-rank test.

$\Delta BAre = BAre_{OP,13M}/BAre_{OP,1M}$	Obs.	#($\Delta BAre_{LQ} < \Delta BAre_{UQ}$)	%($\Delta BAre_{LQ} < \Delta BAre_{UQ}$)	Median $\Delta BAre_{LQ}$	Median $\Delta BAre_{UQ}$	Median % Change ($\Delta BAre_{UQ}/\Delta BAre_{LQ}$-1)
	Panel A: Differences in Option Adoption (Year Following Opt. Listings) between the Groups in the Lower and Upper Quartiles by the PIN_{0Y} (Year Prior to Opt. Listings)					
$\Delta BAre$	152	89	58.55%	1.32	1.49	19.32%[aa,bb]
	Panel B: Differences in Option Adoption (Year Following Opt. Listings) between the Groups in the Lower and Upper Quartiles by the $AdjPIN_{0Y}$ (Year Prior to Opt. Listings)					
$\Delta BAre$	141	84	59.57%	1.22	1.43	23.02%[aa,bbb]
	Panel C: Differences in Option Adoption (Year Following Opt. Listings) between the Groups in the Lower and Upper Quartiles by the $InvAnlst_{0Y}$ (Year Prior to Opt. Listings)					
$\Delta BAre$	146	86	58.90%	1.25	1.38	17.94%[aa,bb]

Table 7

Reduction in *PIN* After Option Listings

The table presents a matched sample analysis of the effects of option listings on changes in the *PIN* measure after the listing date. PIN_{0Y}, PIN_{1Y}, $DVlm_{OP,1Y}$, $Vlm_{OP,1Y}$, and $OInt_{OP,1Y}$ are defined in Table 1. The matched pairs contain *PIN* estimates from the year before and the year after option listings. The table reports results for both the paired-sample sign test and the Wilcoxon signed-rank test applied to measure the change from PIN_{0Y} to PIN_{1Y}. The null hypothesis of no change in *PIN* is tested against the one-sided alternative of PIN_{1Y} being inferior to PIN_{0Y}. Both tests are applied to alternative quartile sub-samples. For instance, the first row shows that out of 891 pairs of matched *PIN* values in the complete sample, 532 have smaller *PIN* after listings than in the year prior to the listing. [a], [aa], and [a] denote significance at 1%, 5%, and 10%, respectively, for the paired-sample sign test. [bbb], [bb], and [b] indicate significance at 1%, 5%, and 10%, respectively, for the Wilcoxon signed-rank test.

Differences in *PIN* between the Year Prior to and the Year Following Option Listings Using Multiple Sub-Samples						
	Obs.	#($PIN_{0Y} > PIN_{1Y}$)	%($PIN_{0Y} > PIN_{1Y}$)	Median PIN_{0Y}	Median PIN_{1Y}	Median % Change (PIN_{1Y}/PIN_{0Y} - 1)
			Complete Sample			
PIN	891	532	59.71%	0.19	0.14	-22.47%[aaa,bbb]
			Lower Quartile by $D\,Vlm_{OP,1Y}$			
PIN	222	125	56.31%	0.17	0.15	-10.61%[aa,bb]
			Upper Quartile by $D\,Vlm_{OP,1Y}$			
PIN	222	144	64.86%	0.23	0.15	-36.35%[aaa,bbb]
			Lower Quartile by $Vlm_{OP,1Y}$			
PIN	222	124	55.86%	0.17	0.14	-10.82%[aa,bb]
			Upper Quartile by $Vlm_{OP,1Y}$			
PIN	222	142	63.96%	0.21	0.15	-34.70%[aaa,bbb]
			Lower Quartile by $OInt_{OP,1Y}$			
PIN	222	127	57.21%	0.17	0.14	-14.66%[aa,bb]
			Upper Quartile by $OInt_{OP,1Y}$			
PIN	222	147	66.22%	0.23	0.14	-35.97%[aaa,bbb]

Table 8

Reduction in *AdjPIN* After Option Listings

The table presents a matched sample analysis of the effects of option listings on changes in the *AdjPIN* measure after the listing date. The matched pairs contain *AdjPIN* estimates from the year before and the year after option listings. The table reports results for both the paired-sample sign test and the Wilcoxon signed-rank test applied to the change from $AdjPIN_{0Y}$ to $AdjPIN_{1Y}$. The null hypothesis of no change in *AdjPIN* is tested against the one-sided alternative of $AdjPIN_{1Y}$ being inferior to $AdjPIN_{0Y}$. Both tests are applied to alternative quartile sub-samples. For instance, the first row shows that out of 891 pairs of matched *PIN* values in the complete sample, 532 have smaller *AdjPIN* after listings than in the year prior to the listing. [a], [aa], and [a] denote significance at 1%, 5%, and 10%, respectively, for the paired-sample sign test. [bbb], [bb], and [b] indicate significance at 1%, 5%, and 10%, respectively, for the Wilcoxon signed-rank test.

Differences in *Adj PIN* between the Year Prior to and the Year Following Option Listings Using Multiple Sub-Samples

	Obs.	#($AdjPIN_{0Y}$ > $AdjPIN_{1Y}$)	%($AdjPIN_{0Y}$ > $AdjPIN_{1Y}$)	Median $AdjPIN_{0Y}$	Median $AdjPIN_{1Y}$	Median % Change ($AdjPIN_{1Y}/AdjPIN_{0Y}$ - 1)
			Complete Sample			
AdjPIN	891	558	62.63%	0.16	0.13	-19.96%[aaa,bbb]
			Lower Quartile by $D\ Vlm_{OP,1Y}$			
AdjPIN	222	125	56.31%	0.15	0.13	-12.02%[aa,bb]
			Upper Quartile by $D\ Vlm_{OP,1Y}$			
AdjPIN	222	149	67.12%	0.18	0.12	-30.22%[aaa,bbb]
			Lower Quartile by $Vlm_{OP,1Y}$			
AdjPIN	222	129	58.11%	0.14	0.13	-14.78%[aaa,bb]
			Upper Quartile by $Vlm_{OP,1Y}$			
AdjPIN	222	152	68.47%	0.18	0.12	-32.17%[aaa,bbb]
			Lower Quartile by $OInt_{OP,1Y}$			
AdjPIN	222	134	60.36%	0.15	0.13	-14.91%[aaa,bbb]
			Upper Quartile by $OInt_{OP,1Y}$			
AdjPIN	222	150	67.57%	0.18	0.12	-30.64%[aaa,bbb]

Table 9

Reductions in *PIN* After the Listing Date in the Control Group

The table presents a matched sample analysis of the effects of option listings on changes in the *PIN* measure after the listing date for a control group of stock options. The control group is created by selecting a seasoned equity option (with at least three years of market activity subsequent to its listing) to match each listing in the original sample. The matching is performed using: the underlying stock industry, the underlying stock return volatility (annualized standard deviation of daily returns over the 252 trading days prior to listing), and underlying stock volume (mean of daily dollar-volume over the 252 trading days prior to listing). All the statistics in this table are calculated using data from the equity options in the control group. PIN_{0Y}, PIN_{1Y}, $DVlm_{OP,1Y}$, $Vlm_{OP,1Y}$, and $OInt_{OP,1Y}$ are defined in Table 1. The matched pairs consist of *PIN* estimates from different time periods for each matching seasoned stock option (in the year before and the year after the listing date). The table reports both the paired-sample sign test and the Wilcoxon signed-rank test applied to the change from PIN_{0Y} to PIN_{1Y}. The null hypothesis of no change in *PIN* is tested against the one-sided alternative of PIN_{1Y} being inferior to PIN_{0Y}. Both tests are applied to alternative quartile sub-samples. For instance, the first row shows that out of 891 pairs of matched *PIN* values in the complete sample, 458 have smaller *PIN* after the listing date than in the previous year. [a], [aa], and [a] denote significance at 1%, 5%, and 10%, respectively, for the paired-sample sign test. [bbb], [bb], and [b] indicate significance at 1%, 5%, and 10%, respectively, for the Wilcoxon signed-rank test.

Differences in *PIN* between the Year Prior to and the Year Following Option Listings Using Multiple Sub-Samples						
	Obs.	$\#(PIN_{0Y} > PIN_{1Y})$	$\%(PIN_{0Y} > PIN_{1Y})$	Median PIN_{0Y}	Median PIN_{1Y}	Median % Change $(PIN_{1Y}/PIN_{0Y} - 1)$
	Complete Sample					
PIN	891	458	51.40%	0.14	0.13	-1.35%
	Lower Quartile by $DVlm_{OP,1Y}$					
PIN	222	114	51.35%	0.14	0.13	-1.34%
	Upper Quartile by $DVlm_{OP,1Y}$					
PIN	222	109	49.10%	0.14	0.14	1.58%
	Lower Quartile by $Vlm_{OP,1Y}$					
PIN	222	116	52.25%	0.14	0.14	-2.29%
	Upper Quartile by $Vlm_{OP,1Y}$					
PIN	222	107	48.20%	0.14	0.15	2.13%
	Lower Quartile by $OInt_{OP,1Y}$					
PIN	222	118	53.15%	0.13	0.13	-1.88%
	Upper Quartile by $OInt_{OP,1Y}$					
PIN	222	110	49.55%	0.15	0.15	2.06%

Table 10

Reductions in *AdjPIN* Levels after the Listing Date: Control Group

The table presents a matched sample analysis of the effects of option listings on changes in the *AdjPIN* measure after the listing date for a control group of stock options. The control group is created by selecting a seasoned equity option (with at least three years of market activity subsequent to its listing) to match each listing in the original sample. The matching criteria are defined in Table 7. All the statistics in this table are calculated using data from the equity options in the control group. $AdjPIN_{0Y}$, $AdjPIN_{1Y}$, $DVlm_{OP,1Y}$, $Vlm_{OP,1Y}$, and $OInt_{OP,1Y}$ are defined in Table 1. The matched pairs consist of *AdjPIN* estimates from different time periods for each matching seasoned stock option (in the year before and the year after the listing date). The table reports both the paired-sample sign test and the Wilcoxon signed-rank test applied to the change from $AdjPIN_{0Y}$ to $AdjPIN_{1Y}$. The null hypothesis of no change in *AdjPIN* is tested against the one-sided alternative of $AdjPIN_{1Y}$ being inferior to $AdjPIN_{0Y}$. Both tests are applied to alternative quartile sub-samples. For instance, the first row shows that out of 891 pairs of matched *AdjPIN* values in the complete sample, 458 have smaller *AdjPIN* after the listing date than in the previous year. [a], [aa], and [a] denote significance at 1%, 5%, and 10%, respectively, for the paired-sample sign test. [bbb], [bb], and [b] indicate significance at 1%, 5%, and 10%, respectively, for the Wilcoxon signed-rank test.

Differences in *Adj PIN* between the Year Prior to and the Year Following Option Listings Using Multiple Sub-Samples

	Obs.	#($AdjPIN_{0Y}$ > $AdjPIN_{1Y}$)	%($AdjPIN_{0Y}$ > $AdjPIN_{1Y}$)	Median $AdjPIN_{0Y}$	Median $AdjPIN_{1Y}$	Median % Change ($AdjPIN_{1Y}/AdjPIN_{0Y}$ - 1)
	Complete Sample					
AdjPIN	891	463	51.96%	0.13	0.13	-1.59%
	Lower Quartile by $D\,Vlm_{OP,1Y}$					
AdjPIN	222	117	52.70%	0.13	0.13	-1.89%
	Upper Quartile by $D\,Vlm_{OP,1Y}$					
AdjPIN	222	107	48.20%	0.14	0.14	0.77%
	Lower Quartile by $Vlm_{OP,1Y}$					
AdjPIN	222	113	50.90%	0.13	0.12	-1.03%
	Upper Quartile by $Vlm_{OP,1Y}$					
AdjPIN	222	110	49.55%	0.14	0.14	0.46%
	Lower Quartile by $OInt_{OP,1Y}$					
AdjPIN	222	115	51.80%	0.13	0.12	-1.98%
	Upper Quartile by $OInt_{OP,1Y}$					
AdjPIN	222	108	48.65%	0.14	0.14	1.12%

Table 11

Increase in Analysts Following after Option Listings

The table presents a matched sample analysis of the effects of option listings on changes in the inverse of the average number of analysts following stocks after the listing date. $InvAnlst_{0Y}$, $InvAnlst_{1Y}$, $DVlm_{OP,1Y}$, $Vlm_{OP,1Y}$, and $OInt_{OP,1Y}$ are defined in Table 1. The matched pairs concern the inverse of the average number of analysts from the year before and the year after option listings. The table reports both the paired-sample sign test and the Wilcoxon signed-rank test applied to the change in $InvAnlst_{0Y}$ from $InvAnlst_{1Y}$ (i.e., a measure of the increase in the number of analysts after listing). The null hypothesis of no change in the inverse of the number of analysts is tested against the one-sided alternative of $InvAnlst_{1Y}$ being inferior to $InvAnlst_0$. Both tests are applied to alternative quartile sub-samples. For instance, the first row shows that out of 891 pairs of matched values of the inverse number of analysts in the complete sample, 610 display a lower *InvAnlst* after the option listing than in the previous year. [a], [aa], and [a] denote significance at 1%, 5%, and 10%, respectively, for the paired-sample sign test. [bbb], [bb], and [b] indicate significance at 1%, 5%, and 10%, respectively, for the Wilcoxon signed-rank test.

Differences in the *InvAnlst* between the Year Prior to and the Year Following Option Listings Using Multiple Sub-Samples						
	Obs.	#($InvAnlst_{0Y}$ > $InvAnlst_{1Y}$)	%($InvAnlst_{0Y}$ > $InvAnlst_{1Y}$)	Median $InvAnlst_{0Y}$	Median $InvAnlst_{1Y}$	Median % Change ($InvAnlst_{1Y}/InvAnlst_{0Y}$-1)
	Complete Sample					
InvAnlst	891	610	68.46%	0.27	0.20	-22.38%[aaa,bbb]
	Lower Quartile by $D\,Vlm_{OP,1Y}$					
InvAnlst	222	137	61.71%	0.21	0.19	-13.48%[aaa,bbb]
	Upper Quartile by $D\,Vlm_{OP,1Y}$					
InvAnlst	222	162	72.97%	0.32	0.20	-37.03%[aaa,bbb]
	Lower Quartile by $Vlm_{OP,1Y}$					
InvAnlst	222	138	62.16%	0.22	0.20	-12.34%[aaa,bbb]
	Upper Quartile by $Vlm_{OP,1Y}$					
InvAnlst	222	155	69.82%	0.30	0.19	-31.17%[aaa,bbb]
	Lower Quartile by $OInt_{OP,1Y}$					
InvAnlst	222	139	62.61%	0.23	0.20	-12.50%[aaa,bbb]
	Upper Quartile by $OInt_{OP,1Y}$					
InvAnlst	222	150	67.57%	0.28	0.18	-28.48%[aaa,bbb]

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