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Information Spillover and Corporate Policies: The Case of Listed Options

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Abstract

Information production associated with derivatives markets is not a sideshow; rather, it has significantly positive spillover effects on an array of corporate decisions of underlying firms. Using a regression-discontinuity design based on exogenous variation in options availability as an instrument for changes in the information environment, we show that options introductions have causal effects on corporate policies on both sides of the balance sheet. Through improved information efficiency, options availability reduces the need for debt and payout, increases efficient investment, and yields superior innovation. We conduct two independent experiments demonstrating that our instrument's impact is not derived from alternative channels.

I. Introduction

Corporate finance theory posits that the quality of a firm's information environment affects its overall access to external funds in general (e.g., Stiglitz and Weiss (1981), Diamond (1985)) and specifically its access to equity financing (e.g.,

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Myers and Majluf (1984)). Information quality also directly and indirectly affects investment policy: directly because reduced information asymmetry reduces investment risk; and indirectly by making external capital more readily available. However, empirically testing the role of information quality remains challenging. Although a firm's information environment depends on many measurable features, including financial analyst coverage, institutional ownership, and disclosure quality, the variations in these features are hardly exogenous to firms' decisions. In this study, we examine the effects of information on corporate decision-making by exploring a unique shock to a firm's information environment. The shock entails one of the most significant innovations in capital markets: the introduction of exchange-listed options. Options introductions are an ideal research setting for this analysis because they are significant and frequent events, staggered over a long period of time, and are decided by the exchange without firm involvement. Although evidence shows that options market activity leads to higher price efficiency (e.g., Jennings and Starks (1986), Hu (2018)), evidence of the real impact of options introductions on corporate behavior is scant. By accounting for endogeneity concerns, our empirical analysis provides support for the notion that exchange-listed options mitigate informational frictions and have real effects on a wide array of corporate actions.

Given the increased informational efficiency associated with options introductions, one would expect investment and financing policies to respond to improvements in the informational efficiency induced by the options market. Theories of both adverse selection/signaling (e.g., Ross (1977), Meyers and Majluf (1984)) and agency (e.g., Jensen and Meckling (1976)) propose that options introductions should yield lower adverse selection costs and better contracting abilities, thereby allowing firms to access external capital markets more easily, increase reliance on equity financing, and invest more with fewer financing constraints. Although options are introduced to allow investors to hedge and/or take speculative positions, our results show that options introductions also have an unintended and important positive spillover effect on corporate decisions concerning both sides of the balance sheet. Our findings suggest that the improved information quality results in greater investment as well as increased and superior innovation; improved information quality reduces the need to use payout and debt to address agency and information frictions.

One of the unique features of our study is its ability to explore the causal impact of options listing on different corporate domains. Specifically, our empirical method combines a regression discontinuity design (RDD) analysis with a new instrumental variable (IV), which allows us to clearly identify the causal effect of options availability on firms' i) equity and debt issues, ii) financial leverage, iii) repurchase intensities and dividend payout, iv) investment intensities and quality, v) cash holdings, and vi) innovation activities. All our tests rely on the same factor (namely, options availability) to assess its impact on an array of corporate policies. Our holistic approach arguably sets a high hurdle, because any explanation of the evidence will require consistency across these domains. Our results show that the informational role played by the availability of options is consistent with the effect of agency and asymmetric information on corporate decisions (e.g., La Porta, Lopez-De-Silanes, Shleifer, and Vishny (1997), Derrien and Kecskes (2013)) as well as being consistent with the feedback effect from prices to corporate actions (e.g., Chen, Goldstein, and Jiang (2007)). Specifically, we find that when firms' equity is linked to listed options, firms raise more external capital, rely more heavily on equity financing, reduce leverage, reduce payout, invest in greater quantity, build larger cash reserves, and invest more efficiently.

The most significant challenge in analyzing the causal effect of options may be that options availability is not random, despite the important feature that options' introduction is initiated by the exchange rather than the firm, unlike initial listings of equity shares. The exchange is likely to choose stocks with specific characteristics such as high trading volume and high volatility to attract options traders. These considerations raise the concern of endogeneity from omitted variables that may affect both the options listing decision and variables of interest such as firms' financial and investment decisions. Prior literature has typically used options open interest and average moneyness to instrument options trading volume to study the effect of options trading (e.g., Roll, Schwartz, and Subrahmanyam (2009), Blanco and Wehereim (2017)). However, both of those IVs are outcomes of options traders' choices and can be subject to the endogeneity concern that the same confounding factors drive both options trading activity and a corporate covariate.

Building on these earlier efforts to identify the effects of options trading, we combine an RDD analysis with a new IV for options availability. Our IV exploits the Securities and Exchange Commission's (SEC) requirements for options listing under the Options Listing Procedure Plan (OLPP). Those rules require the underlying equity to have: i) a minimum public float of 7 million shares; ii) at least 2.4 million shares traded in the past 12 months; iii) at least 2,000 shareholders; and iv) a share price above \$7.50 prior to 2003 or \$3 thereafter. Based on these rules, our IV measures the eligibility of a firm for the treatment of options listings. This IV in essence exploits the random variations in the likelihood of having options due to either satisfying or failing to satisfy the regulatory requirements. Intuitively, our IVs should have the best ability to identify causal effects from options trading for firms around the regulatory threshold because these firms have similar characteristics. We, therefore, conduct an RDD analysis as our main empirical method.

To increase the power of the test and allow for conditional analysis, we also supplement the RDD analysis using all observations regardless of their distance to the regulatory thresholds in standard IV estimations. The results are qualitatively the same in the RDD and full sample IV analyses, although the economic magnitude of the full sample IV estimate is typically larger than that of the RDD analysis due to the inclusion of firms further away from the regulatory threshold. In all our analyses, we control for firm fixed effects (FEs) that remove unobservable time-invariant firm characteristics as well as year FE.

Of equal importance is the fact that we are able to support the validity of our instrument using two natural experiments. Meeting the regulatory requirements should only increase the likelihood of options listing when options markets exist. Therefore, the 1973 opening of the first public options market, the Chicago Board Options Exchange (CBOE), becomes an exogenous shock pertinent to our IV. Specifically, we find that although the IV has significant effects on the corporate policies of interest post-1973, its impact is muted before the initiation of public options trading.

Second, constructing the IV based on regulatory rules also enables the exploitation of exogenous changes in SEC rules. We focus on a major reform in 2003 of the options listing standards that makes low(er) price stocks eligible for options trading. We find that the treated firms, which became eligible only after the rule change, exhibit significant changes in their corporate policies relative to control (no change in eligibility) firms after the rule revision, whereas the two groups exhibit no significant differences before the rule revision. The results from these two natural experiments reinforce the causal impact of options trading on corporate policies.

We highlight several of our important findings below. First, both equity and debt issues significantly increase when firms' equity is linked to exchange-traded options. We also calculate the economic impact of our RDD and full sample IV results: they imply that options listing increases equity issues by 7% to 10.6% of the unconditional sample mean. At the same time, debt issues also increase by 4.7% to 10.2% of the unconditional sample mean. Thus, raising external capital using both equity and debt substantially increases as a result of options introduction.

Second, agency and adverse selection theories (e.g., Jensen and Meckling (1976), Myers (1984), Myers and Majluf (1984), and Zwiebel (1996)) also propose that listing options has a greater impact on equity than it does on debt due to equity's greater sensitivity to information effects. We find that debt-to-asset ratios become significantly lower (t-stat below -5) when firm equity is linked to exchange-traded options. The magnitude of this effect is noteworthy, and indicates a reduction of 2.3% to 3.3% in financial leverage relative to the sample mean.

Third, extant theory also predicts that the quality of a firm's information environment affects its incentives to distribute cash to investors. Specifically, when a firm experiences an options introduction, the reduced information asymmetry should reduce the value of a dividend payment both as a disciplinary device (e.g., Jensen (1986)) and a signaling device (e.g., Miller and Rock (1985), Michaely, Rossi, and Weber (2017)). Therefore, we expect dividends (and repurchases) to decrease after options are introduced. The results from RDD and IV estimators consistently support this prediction. Options introductions are associated with a decline in both repurchases and dividends. These effects are statistically significant in the full sample but less significant in the RDD sample. Taken together, our results suggest that options listings have a direct effect on leverage through an increased propensity to issue equity relative to debt, and an indirect and smaller effect on leverage through the effect on payout. Both the direct and indirect effects lead to lower financial leverage after options listings.

Because options trading improves the quality of information available to investors, the positive announcement returns associated with repurchases and the negative announcement returns associated with equity offerings should become less pronounced. Consistent with this prediction, we find that announcement returns of equity offerings and share repurchases are significantly lower in absolute magnitude when the firm's equity is linked to options. The magnitude of these effects is significant. For equity offerings, the 3-day announcement cumulative abnormal return (CAR) increases by 0.36%; for open market repurchases, it decreases by 0.31%. Given that the average CAR of equity offerings (repurchases) is around −1.77% (2.01%) in our sample, these effects are economically significant.

This evidence directly supports the claim that the costs of equity issuance and the benefits of equity buybacks are affected by listed options in a manner consistent with the impact of lower information costs when options are available.

When we turn to the asset side of the balance sheet, we conjecture that by reducing information asymmetry and relaxing financial constraints, options trading can lead to greater corporate investment (e.g., Myers and Majluf (1984), Chen et al. (2007)). We find that capital expenditures increase by 1.7% to 4.2% of the unconditional sample mean. Using the same line of reasoning, we also examine the impact of listed options on firms' innovation activity. The nature of firm innovation argues for this type of investment being most vulnerable to both adverse selection and agency conflicts. We find that firm innovation output significantly increases. In addition to the level of investment, investment efficiency should also increase as more efficient stock prices facilitate investor monitoring, thus mitigating the agency problem (e.g., Holmstrom and Tirole (1993), Zhu (2019)). Therefore, we expect firm investments to become increasingly responsive to growth opportunities. Indeed, we find that firm investment-q sensitivity becomes significantly larger when exchange-traded options are linked to the stock. Taken together, the evidence from the asset side of the balance sheet consistently supports the idea that listed options have a real, causal effect on firms' investment policies. The presence of listed options demonstrably yielded greater levels of investment and more efficient investment when measured in terms of both capital expenditures and innovation. The impact of listed options on the array of corporate decisions we examine is summarized in Figure 1.

After establishing the causal impact of options listings on corporate policies using the IV in an RDD analysis and the associated natural experiments, we also examine the intensive margin effects of options trading. Here, we extend the analysis not only to whether options are listed but also to the intensity of trading: the greater the trading intensity, the more information is revealed. We, therefore, investigate the effect of options trading volume and open interest on the same set of corporate policies. Our results support the conjecture that more intense options trading improves the information environment such that firms with higher options trading volumes and open interest raise more external capital, sustain lower debt-to-asset ratios, and sustain higher capital expenditures. Consistently, these firms also favor equity to debt and have lower payout.

Cross-sectional tests indicate that the effects of options listings on corporate policies are more pronounced when the firm's information environment is opaque. For example, we find that the effects of options introductions on corporate policies are consistently higher when analyst coverage of the underlying firm is low, when the probability of informed trading on the stock is high, or when the institutional ownership is low. We also find that firms in the early stage of their life cycle (small and young firms) tend to experience a larger impact from options trading. These results support our interpretation of the effects of options listings on firm behavior through an information channel.

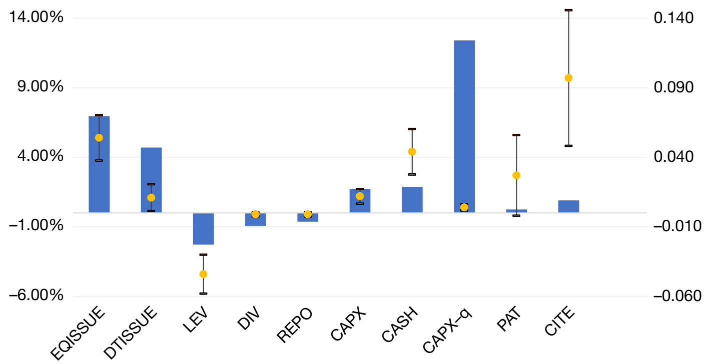
Our study contributes to a growing body of research that investigates the impact of changes in the information environment on aspects of firms' decision-making. For example, Derrien and Kecskes (2013) use changes in analyst coverage; Aghion, Van Reenen, and Zingales (2013), Crane, Michenaud, and Weston (2016),

FIGURE 1

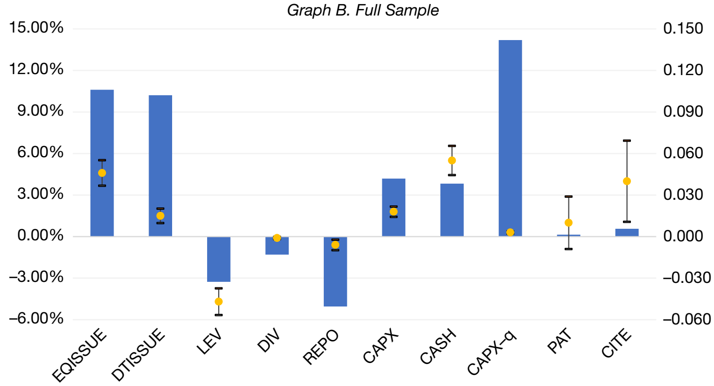
Summary of the Impact of Options Listing

Figure 1 presents the estimated coefficients on OP in the second-stage regressions of 2SLS together with the 90% confidence intervals (right axis) for all corporate policies we examine. We also plot the economic significance as the average percentage change in the corporate policy relative to the unconditional sample mean in blue bars (left axis).

Graph A. RDD Sample



Graph B. Full Sample



and Heath, Macciocchi, Michaely, and Ringgenberg (2022) use changes in the composition of institutions around index reconstitution; and Sufi (2009) and Tang (2009) use the certification effects of credit ratings. Although these studies primarily focus on shocks originating from financial intermediaries, our exploration of the impact of capital market innovations (i.e., introducing options) on a wide array of corporate decisions provides direct evidence of the importance of information spillover from capital markets to corporate actions. Our evidence, which is based on a large sample of events in readily observable public exchange markets, staggered over a long period, and derived from a variety of identification methods, demonstrates that information shocks from capital market innovation have a causal impact on corporate behavior. This result further highlights that the impact of options (perhaps the most important financial innovation in the last 50 years) extends beyond underlying price efficiency.[1] The reductions in both information

[1]In addition to price efficiency, prior studies have examined options' impact on the underlying price level (e.g., Conrad (1989)) and volatility (e.g., Skinner (1989), Mayhew and Mihov (2004)).

opacity and agency friction lead to economically impactful changes in firm behavior in the form of increased reliance on external financing, increased investment intensity and efficiency, increased investment in innovation, greater dependence on equity financing, and lower levels of payout. We, therefore, conclude that options markets are not a sideshow to corporate decisions.

II. Empirical Framework

A. Data, Sample Selection, and Variable Construction

We construct our sample by merging the CRSP, Compustat, and OptionMetrics databases for the period of 1996 to 2019, while requiring that firm-level data be available in CRSP and Compustat. We exclude financial and utilities firms as well as those with total assets or sales below 1 million USD in Compustat.

We identify options availability using a dummy variable, OP, which is equal to 0 if the stock has options price records in OptionMetrics during the corresponding firm-year, and 0 otherwise. To sharpen our identification of options effects, we focus on the 5 years before and after a firm is chosen by options exchanges for the first time during our sample period. For the same reason, we also exclude firms that have listed options throughout the study period. The final sample contains 6,050 firm-year observations from 1,065 unique firms classified as having options. The sample also includes 31,944 firm-years (from 4,049 unique firms) without options as the control group.

We use a comprehensive set of variables to measure corporate financial policies. All variables are calculated at the fiscal year-end unless otherwise specified. We measure equity issuance (EQISSUE_{it}) as the value of equity issued by firm i in year t divided by the book value of assets at the beginning of the year. Debt issuance (DTISSUE_{it}) is the long-term debt issued minus the long-term debt reduction in year t divided by total assets at the beginning of the year. A related variable of interest is the change in the financial leverage as a result of options introductions. We define financial leverage (LEV_{it}) as the book value of debt divided by the book value of assets. We also examine firm payout policies in detail. Equity repurchases (REPO_{it}) is the change in the number of shares outstanding divided by the number of shares at the beginning of the year. Finally, dividends (DIV_{it}) is calculated as the dividends paid divided by the book value of assets.

To measure firm investment, we use capital expenditures divided by the book value of assets (CAPX_{it}). We define corporate cash holdings (CASH_{it}) as cash and cash equivalents divided by total assets. Following the literature on firm innovation (e.g., Hall, Jaffe, and Trajtenberg (2005)), we use two patent-related variables to measure innovation outcome. For all successful patent applications submitted in a given year, PAT_{it} is the natural logarithm of 1 plus the number of patents, and CITE_{it} is the natural logarithm of 1 plus the average number of citations per patent.

Following the strand of empirical corporate finance literature including studies such as Shyam-Sunder and Myers (1999) and Frank and Goyal (2003), we control for the following differences across firms: firm size, estimated as the natural logarithm of total assets in book value ($ASSET_{it}$); asset tangibility, calculated as the ratio of fixed assets to total assets ($TANGIBILITY_{it}$); Tobin's q, calculated as the market value of equity plus the book value of debt divided by the book value of assets (MB_{it}); return on assets, defined as net income divided by the book value of assets (ROA_{it}); free cash flow (FCF_{it}), defined as earnings before interest and taxes plus depreciation and amortization divided by the book value of assets at the beginning of the year; and the number of years after the initial public offering (AGE_{it}). Panel A of Table 1 reports the summary statistics of these variables in the full sample as well as, separately, in the subsamples of firms with and without options. These univariate results imply that, on average, firms with options issue less equity but more debt and have higher financial leverage than those without options. Firms with options repurchase more equity and pay more dividends. These firms also invest at higher levels and generate more patents and citations. However, firms with options clearly differ from those without options. Consequently, one should use caution when interpreting these univariate comparisons. Firms with options typically have more assets, higher market valuations, higher returns on assets, and greater free cash flows. As such, our empirical analysis carefully addresses the selection by options exchanges using a combination of an IV approach, firm FE regressions, regression discontinuity analysis, and a propensity score matching (PSM) method.

B. Instrumental Variable Estimator

Our main identification of the causal effect from options trading takes advantage of the exogenous regulatory restriction in the options listing process. Unlike decisions concerning the listing of equity on an exchange (e.g., through an IPO), which are made by the company, options listing decisions are made by options exchanges without the underlying company's approval. Although an options listing event is an exogenous shock from the firm's perspective, options listing decisions are made by exchanges for profit maximization purposes and therefore are hardly random. As such, options listing events may reflect selection biases that hamper a clean identification of the options listing's treatment effects.

To overcome this selection bias, we draw causal inferences from tests based on an IV approach. Our instrument for the availability of listed options exploits the SEC's requirements that must be satisfied by the underlying stock to be eligible for options listing under the OLPP.[2] Eligibility for options listing entails the following requirements for each stock: i) a minimum public float of 7 million shares; ii) at least 2.4 million shares traded in the past 12 months; iii) at least 2,000 shareholders; and iv) the stock price must be above $7.50. These requirements provide an advantageous setting to study the treatment effect of options: if two firms fall on opposite sides of a regulatory threshold, the probability of each firm's stock being linked to listed options will be significantly different even when they have identical

[2]The OLPP is a national market system plan that describes the procedures to be followed by all the options exchanges in the United States in selecting underlying securities. See details at https://www.theocc.com/getmedia/198bfc93-5d51-443c-9e5b-fd575a0a7d0f/options_listing_procedures_plan.pdf.

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TABLE 1

Summary Statistics

Table 1 presents the firm-year observations merged from Compustat, CRSP, and OptionMetrics between 1996 and 2019 after excluding financial firms and utilities and firms with total assets or sales below 1 million dollars. We include 5 years of observations before and after each first-time options listing event identified in OptionMetrics. We exclude firms that always have listed options during the sample period. Panel A presents statistics for all firm-year observations. Panel B presents statistics for the sample of regression discontinuity design (RDD) analysis, where the firm is close to the regulation threshold for options listing eligibility, as detailed in Section II.C. EQISSUE is stock issuances minus equity repurchases divided by book assets at the beginning of the year. DTISSUE is long-term debt issuances minus changes in long-term debt divided by book assets at the beginning of the year. LEV is the book value of debt divided by book assets. REPO is the change in the number of shares outstanding divided by the number of shares at the beginning of the year. DIV is dividends divided by book assets at the beginning of the year. CAPX is capital expenditures divided by book assets at the beginning of the year. CASH is cash and cash equivalents divided by book assets at the beginning of the year. PAT is the natural logarithm of 1 plus the number of successful patent applications during the year. CITE is the natural logarithm of 1 plus the average number of citations per patent for all patents applied for during the year. ASSET is the natural logarithm of book assets. TANG is the ratio of fixed assets to book assets. MB is the market capitalization plus the book value of debt divided by the book value of equity plus the book value of debt. ROA is net income divided by book assets. FCF is earnings before interest and taxes plus depreciation and amortization divided by book assets at the beginning of the year. AGE is firm age (in years) since the initial public offering. PRICE is the average daily stock price during the firm-year. VOLUME is the total number of shares traded during the firm-year. FLOAT is the number of shares held by noninsiders. All variables are winsorized at the 5th and 95th percentiles.

Panel A. Full Sample

	All Firms (N = 37,994)		OP = 0 (N = 31,944)		OP = 1 (N = 6,050)	
Variable	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Dependent Variables						
EQISSUE	0.062	0.204	0.061	0.201	0.070	0.222
DTISSUE	0.021	0.135	0.020	0.134	0.030	0.136
LEV	0.206	0.237	0.210	0.242	0.184	0.211
REPO	0.017	0.079	0.019	0.083	0.009	0.048
DIV	0.011	0.028	0.011	0.027	0.013	0.029
CAPX	0.061	0.085	0.059	0.083	0.074	0.096
CASH	0.204	0.250	0.192	0.241	0.263	0.288
PAT	4.204	67.771	3.287	52.233	9.046	120.051
CITE	4.562	23.817	4.334	24.373	5.764	20.589
Independent Variables						
ASSET	1,033	6,563	748	5,005	2,541	11,643
TANG	0.262	0.241	0.259	0.236	0.278	0.263
MB	1.907	1.584	1.856	1.582	2.176	1.566
ROA	0.023	0.234	0.013	0.238	0.071	0.199
FCF	0.000	0.239	−0.010	0.243	0.051	0.208
AGE	15.925	11.848	15.688	11.761	17.172	12.223
PRICE	3.128	4.831	2.408	4.167	6.929	6.141
VOLUME	16.050	50.307	10.260	42.315	46.622	72.998
FLOAT	3.782	9.524	3.308	9.436	6.287	9.597

Panel B. RDD Sample

	All Firm-Years (N = 15,028)		OP = 0 (N = 13,608)		OP = 1 (N = 1,420)	
Variable	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Dependent Variables						
EQISSUE	0.073	0.219	0.069	0.215	0.102	0.255
DTISSUE	0.022	0.134	0.023	0.136	0.012	0.113
LEV	0.182	0.217	0.185	0.217	0.156	0.219
REPO	0.015	0.071	0.016	0.072	0.010	0.054
DIV	0.010	0.026	0.010	0.026	0.009	0.025
CAPX	0.065	0.088	0.065	0.088	0.062	0.089
CASH	0.219	0.258	0.210	0.252	0.297	0.296
PAT	2.994	45.356	2.973	46.747	3.199	28.824
CITE	5.742	28.604	5.749	29.500	5.678	17.875
Independent Variables						
ASSET	565	4,510	507	3,485	1,121	9,931
TANG	0.249	0.228	0.248	0.226	0.252	0.252
MB	1.901	1.566	1.878	1.553	2.117	1.671
ROA	0.031	0.221	0.034	0.217	−0.001	0.255
FCF	0.007	0.231	0.010	0.228	−0.018	0.255
AGE	15.333	11.656	15.233	11.729	16.283	10.884
PRICE	2.625	4.032	2.481	3.844	3.996	5.324
VOLUME	10.232	34.350	7.915	29.644	32.432	59.367
FLOAT	2.444	5.991	2.222	5.789	4.572	7.331

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fundamental characteristics.[3] From that perspective, and given the intrinsic randomness characterizing whether firms meet all SEC options listing requirements, these institutional features provide an ideal setting in which to study the causal effects of listed options. Importantly, the minimum stock price required for options listing was reduced to $3 in Jan. 2003, which we exploit in a separate test.[4]

Moreover, the significance of the regulatory requirements is not limited to initial listing decisions but remains relevant on a continual basis. The requirements determine whether new options series on an underlying stock can be added after an initial listing. Therefore, for a given stock to be linked to exchange-listed options, it must meet specific regulatory requirements on an ongoing basis.[5]

For each firm-year observation, we first calculate the public float as the number of shares outstanding minus the number of shares held by insiders at the end of each quarter as well as the 12-month rolling-window trading volume for every trading day. Then, we compute the average public float across quarters and the average rolling-window trading volume and stock price across all trading days. In our main analysis, we use the concurrent initial listing requirements to construct the instrument. The options listing eligibility indicator, MEET_{it}, is equal to 1 if the firm-year observation satisfies all three listing requirements regarding stock price, trading volume, and public float.[6] Figure 1 also plots the number of eligible stocks (i.e., for whom MEET_{it} equals 1) in each year and shows that in a typical year during our period, the number of stocks eligible for an initial options listing is greater than the number of existing actual listings. However, this gap between potential and actual options listings markedly narrows as options gain popularity and the difference reverses in later years.[7]

We run 2-stage least squares (2SLS) regressions using the following specification:

$$Y_{it} = \alpha_{1t} + \alpha_{2i} + \beta \times \text{OP}_{it} + \Omega \times X_{it-1} + \in_{it}, \tag{1}$$

where OP is instrumented using MEET, Y is a spectrum of corporate policies we examine, and X is a set of firm characteristics.

A valid instrument must satisfy the relevance condition. The correlation between the options availability dummy (OP_{it}) and the options listing eligibility indicator (MEET_{it}) is 0.532 in our sample, thus indicating that the IV is closely

[3]Although options exchanges may list stocks that do not meet the requirements, listing options on those stocks requires special approval from the regulator and is costly to implement. Further, Mayhew and Mihov (2004) and Hu (2018) show that these exemptions are rare.

[4]See SEC Release No. 34-47190 (https://www.federalregister.gov/documents/2003/01/22/03-1347/self-regulatory-organizations-order-approving-proposed-rule-change-by-the-chicagoboard-options).

[5]The corresponding thresholds for continued listings are generally less stringent than those for initial listings. During our sample period, for continuous listing on options exchanges, a stock must have at least 6.3 million shares in public float and 1.8 million shares traded in the past 12 months; the other rules are the same as per initial listing.

[6]We do not apply the rule stipulating the number of shareholders because this information is not publicly available and shareholders typically register their ownership in street names. Although this empirical choice is likely to add noise to our IV, it does not bias our results toward spurious discovery.

[7]This is possible because relatively more stocks have become linked to options over the years and continued listing requirements are generally less stringent than those for initial listing.

related to the firm's actual options trading status. Indeed, among the 6,050 firm-year observations identified to have options in our full sample, MEET_{it} equals 0 for only 587 observations, or 9.71% of this subsample. This evidence suggests that the listing standards are binding and the instrument we use is relevant for studying the effects of options listings on corporate policies. We formally test this condition using the first-stage results of the 2SLS regressions in Section III.A.

A valid instrument must also satisfy the exclusion restriction in that its effect on the outcome variables occurs solely through the treatment, which in our case is options availability. Our instrument is created based on the joint effect resulting when a firm meets three distinct requirements on stock price, trading volume, and public float. We perform tests explicitly tackling the exclusion restriction using 2 natural experiments that exogenously affect the efficacy of our instrument. We conjecture that if our instrument has other channels for affecting corporate policies in addition to options availability, these effects should also exist when the link between the instrument and options availability is broken. Specifically, we use falsification tests in settings where the instrument has no effect through options trading to evaluate whether the exclusion restriction has been violated. Our first test utilizes the 1973 opening of the first public options market: the CBOE. Prior to the CBOE's initiating trading options in public markets, meeting the SEC's requirements regarding options listing should have no impact on corporate policies through options availability. For example, if our instrument is valid, whether the stock price is above or below $7.50 should not have any impact on corporate decision-making before options become available. If what we identify as options effects are a result of omitted correlated variables, we should find that they are correlated with the array of corporate policies we examine and that this effect should hold prior to 1973. Because the options database begins in 1996, we are unable to observe options trading status (OP_{it}) before that year. Therefore, we run firm FE regressions of corporate policies on the IV before and after 1973 and compare results for the two periods.

The second test concerns an exogenous shock from a change in the SEC's options listing standards. In Jan. 2003, the minimum stock price required for options listing was reduced from $7.50 to $3. This rule change exogenously expanded the set of stocks eligible for options trading. Specifically, stocks with prices between $3 and $7.50 became more likely to have options when these stocks were previously ineligible. This natural experiment enables the identification of the causal impact of meeting the SEC's price rule. We construct a sample of observations with prices below $7.50 from 1996 to 2019. We also require that these observations have trading volume and public float above the regulatory thresholds regarding options listing to ensure the test effect results solely from satisfying or not satisfying the price rule, which changes exogenously in the experiment. In this sample, firms with stock prices above $3 receive treatment of increased propensity of options listing when the rule changes, whereas those priced below $3 are the control firms that are unaffected by the reform. We then create an alternative instrument (i.e., treatment) based on whether the stock price is above $3 and below $7.50. If meeting the SEC's price rule affects corporate policies exclusively through options listing, we expect this alternative instrument to be effective only after 2003. Although the inference from this test may be limited to the role of the SEC's price

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rule, the test complements the other natural experiment of the CBOE opening. Other non-options-related events may influence the relation between the rule variables and corporate policies, but these events are unlikely to always coincide with exogenous changes in options listing rules.

Finally, regarding possible direct effects from any individual rule variable and not through options trading, we control for the three listing rule variables in all our IV estimations. This approach enables each listing requirement factor to potentially affect corporate policies directly while ensuring that our IV exclusively reflects a discrete effect on the availability of listed options, which stems from the three thresholds being jointly met during the corresponding firm-year. In this specification, a violation of the exclusion restriction will occur only if the discrete effect stemming from this joint meeting of the three requirements directly affects corporate policies. This inference appears unlikely.

C. Regression Discontinuity Analysis

The regulatory requirement for options listing should have the largest impact on the treatment for firms close to the threshold, because firms around these cutoff points have similar characteristics but distinct probabilities of being selected for options listing. Our instrument can sharply identify the local treatment effect of having options due to a firm meeting the regulatory requirements compared with another firm that has similar characteristics but falls on the other side of the threshold. Note that options exchanges do not list all eligible stocks. Therefore, we apply a fuzzy regression discontinuity (RD) design to evaluate the causal effects of options around the cutoff points of eligibility.

We follow Cattaneo, Titiunik, and Vazquez-Bare (2020) and Wong, Steiner, and Cook (2013) to tackle the fuzzy RD problem using multiple running variables. Specifically, the eligibility is based on three variables: stock price, trading volume, and public float. We focus on observations that have at least one of the three running variables falling in a small region near the cutoff $[C - \mu, C + \mu]$, where C is the regulatory cutoff point and μ is close to 0. The other running variables do not need to be on the boundary, but they must be greater than C so that they do not obstruct treatment assignment once the marginal variable crosses the cutoff. We normalize the three rule variables so we can use the same values of C and μ in sampling. Thus, for every firm-year observation, we divide the stock price, trading volume, and public float by the corresponding regulatory cutoff value. A firm is eligible for options listing if all the normalized rule variables are greater than 1.

We use 2SLS to estimate the fuzzy RD effects, again using MEET to instrument OP, controlling for the three normalized running variables as well as firm and year FE. We choose μ equal to 0.6 using the robust bandwidth selections suggested by Calonico, Cattaneo, Farrell, and Titiunik (2017). Considering the listing threshold for stock prices is set to \$3 after 2003, the bandwidths we choose reflect narrow price regions of \$1.20–\$4.80. We use an alternative μs of 0.5 and 0.7 in the Supplementary Material, which generate largely the same results, thus indicating that the conclusions were not sensitive to the bandwidth choice.

Panel B of Table 1 describes this sample in the RD analysis, which includes 1,420 firm-years with options and 13,608 firm-years without options. The differences between the independent variables such as total assets are notably smaller than those from the full IV sample in Panel A, thereby suggesting that the observations close to the regulatory threshold of options listing are indeed more similar and the RD analysis most closely represents a randomized experiment. Therefore, we use the RD analysis as our main empirical strategy in studying the causal impact of options trading on corporate policies. We also aid the inference using estimation in the full IV sample.

We believe the IV we use, combined with the RDD setting, enables us (and future research) to more cleanly examine the direct impact of options listing on various corporate policies. To test robustness, we also use a firm FE panel setting, which rules out explanations based on unobserved time-invariant firm characteristics and PSM, where treated and control firms are matched on important observables. Both methods have their advantages. Whereas firm FE analysis has potential application in larger cross sections, the PSM method compares the paired treated and control firms during the same time window and thus allows for nonlinearities. We, therefore, conduct these analyses as well. As we describe later in the article, the results from these analyses reinforce our main conclusions stemming from the IV analysis. They are described in more detail in the Supplementary Material.

III. Results

A. First-Stage Results

Our identification strategy of the options listing effects utilizes an IV: the eligibility of options listing. We first check whether our instrument satisfies the relevance condition. In Table 2, we report the first stage output of our 2SLS IV estimation in both the RD and full samples. Throughout our analysis, we cluster standard errors at the firm level. Looking at the RD sample first, with firm and year FE included, the univariate estimation result in column 1 of Table 2 shows that the estimated coefficient on MEET_{it} is 0.118 with a t-statistic of 12.90. We include the three rule variables, namely, stock price, trading volume, and public float, as well as all the control variables in the second-stage regressions for corporate financial and investment policies in columns 2 and 3 of Table 2, respectively. The coefficient on MEET_{it} decreases slightly to 0.094 (t-stat = 11.18) in column 2 and 0.093 (t-stat = 11.08) in column 3. Meanwhile, the coefficients on stock price and trading volume are positive and significant, thereby suggesting that options exchanges prefer stocks with high prices and large trading volumes. Many firm characteristics also have significant coefficients in the first-stage regressions, thus indicating that options listings are more likely for large and mature firms and firms with high market valuations. The F-statistics in columns 2 and 3 exceed 400, thereby suggesting that MEET is a strong instrument for OP. When we turn to the full sample results in columns 4–6, we find the results are qualitatively the same, except that the magnitude and statistical significance of the relation between our instrument and the actual treatment become larger.[8]

[8]There are several reasons why the estimated coefficient on MEET may appear to be small relative to what one might expect if it predicted option listings perfectly. First, both MEET and OP are measured

TABLE 2

First-Stage Estimation for Options Availability

Table 2 reports the first-stage results of the 2SLS regressions. The dependent variable (OP) is an options availability dummy that equals 1 if the stock appears in OptionMetrics in that year, and 0 otherwise. The instrumental variable (MEET) is an options listing eligibility dummy that equals 1 if the stock satisfies all options listing requirements of the SEC during the year, and 0 otherwise. PRICE is the average daily stock price during the firm-year in natural logarithmic form. VOLUME is the total number of shares traded during the firm-year in natural logarithmic form. FLOAT is the number of shares held by noninsiders in natural logarithmic form. ASSET is the natural logarithm of book assets. TANG is the ratio of fixed assets to book assets. MB is the market-to-book value of assets ratio, calculated as the market capitalization plus the book value of debt divided by book assets. ROA is the return on assets, calculated as net income divided by book assets. FCF is free cash flow, calculated as earnings before interest and taxes plus depreciation and amortization divided by book assets. AGE is firm age (in years) since the initial public offering. Both firm and time fixed effects are included. Standard errors are clustered by firm because options listing occurs at the firm level. Corresponding *t*-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	RDD Sample (1996–2019)			Full Sample (1996–2019)		
	OP	OP	OP	OP	OP	OP
	1	2	3	4	5	6
$MEET_{it}$	0.118**	0.094**	0.093**	0.210**	0.143**	0.142**
	(12.90)	(11.18)	(11.08)	(20.72)	(15.66)	(15.48)
$FLOAT_{it}$		0.000	0.000		−0.001*	−0.001*
		(0.05)	(0.03)		(−2.47)	(−2.43)
$PRICE_{it}$		0.004+	0.004+		0.011**	0.011**
		(1.92)	(1.84)		(7.85)	(7.76)
$VOLUME_{it}$		0.002**	0.002**		0.001**	0.001**
		(7.42)	(7.46)		(11.40)	(11.48)
$ASSET_{it-1}$		0.092**	0.089**		0.097**	0.094**
		(11.13)	(11.16)		(18.79)	(18.58)
$TANG_{it-1}$		0.014			0.006	
		(0.33)			(0.22)	
MB_{it-1}		0.005*	0.005*		0.008**	0.008**
		(2.09)	(2.08)		(5.21)	(5.18)
ROA_{it-1}		−0.038+			−0.052**	
		(−1.81)			(−4.77)	
FCF_{it-1}			−0.008			−0.008
			(−0.55)			(−0.93)
AGE_{it-1}		0.030*	0.030*		0.031**	0.031**
		(2.48)	(2.54)		(5.84)	(5.82)
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
F-statistic		423.397	480.373		1,228.397	1,227.475
No. of obs.	15,028	15,028	15,028	37,994	37,994	37,994
R^2	0.116	0.187	0.186	0.163	0.261	0.260

This result indicates that the likelihood firm equity is linked to exchange-listed options increases significantly when its stock meets all SEC requirements for options listing. The economic significance of this relation is large, implying an increased likelihood of having options at about 0.09 in the RD sample and about 0.14 in the full sample, even after controlling for all rule variables and firm characteristics. These effects reflect almost 100% of the unconditional frequency of options availability of only 0.09 in the RD sample and 0.16 in the full sample. This evidence strongly supports the claim that meeting the SEC thresholds for

with errors because we cannot pinpoint the options listing dates exactly or account for the threshold requirement for number of shareholders due to data limitations. Second, because we track stocks for some years post-options listing and options delisting requirements are less stringent than listing requirements; and thus, the predicted relation between MEET and OP is naturally attenuated as a result. Last, the options exchange may exercise discretion when selecting stocks for options listing even if all requirements are met or request an SEC exemption from listing requirements when they are not met. While this added uncertainty attenuates the predicted effect of MEET on OP, it also limits the possibility that firms may meet the listing threshold via manipulation, a concern typical in this type of experimental setting centered on regulatory thresholds.

options listing is relevant to explaining the variations in options availability in our sample (i.e., our IV meets the relevance condition).

Finally, the fact that the three conditions (public float, volume, and price) for options listing must be met on a continual basis suggests that manipulating the listing requirement variables for the purpose of increasing/decreasing the chance of listing for a particular stock is highly unlikely. Therefore, a firm's eligibility for having options has a low probability of being endogenous to the desires of either its management or investors. Nonetheless, in Table A3 in the Supplementary Material, we conduct a robustness test to address the concern that firms may intentionally use stock splits and reverse splits to affect their eligibility for options listing. We find no impact of stock splits on our results.

B. Options Listing and Financial Policies

Theory predicts that with the increased information quality, both adverse selection and agency frictions decrease. Consequently, as discussed earlier, one of the main predictions of the effect of options listing on financing policies is that firms are able to rely more heavily on external financing. With this in mind, we start by examining the issuance activities of debt and equity. The impact on equity issuance is reported in column 1 of Table 3. The dependent variable is the value of equity issued scaled by total assets; the main independent variable is the existence of options on the underlying stock, instrumented by the eligibility of options listing. Our main RD specification result in Panel A of Table 3 suggests that for firms on the regulatory boundary of options listing requirements, within-firm, and controlling for year FE, the introduction of options has a positive and significant impact on equity issuance. The coefficient on the options availability dummy (OP) indicates that options introductions lead to a statistically significant increase of 5.4 percentage points in equity issues with a *t*-statistic of 5.46. Because the causal impact of options trading is identified using exogenous variations in options listing eligibility, to gauge the economic impact, we consider both the increased options listing probability due to meeting the requirement in the first stage and the corporate policy response to options trading in the second stage of the 2SLS regressions. For ease of discussion, we also scale the economic impact by the unconditional sample mean. For equity issuances, passing the regulatory thresholds increases the options listing likelihood by 0.094 in the first-stage regression (column 2 of Table 2); the estimated average causal effect on equity issuance due to meeting these requirements is therefore 0.51% of total assets (0.094×0.054). Given that the average ratio of equity issuance to total assets is 7.3 percentage points in the RD sample, the economic magnitude of this effect is significant and represents about 7% of the unconditional sample mean.

After establishing the legitimacy of causal interpretation using the RD analysis and the instrument, we feel more comfortable applying our analysis to the full sample using the same instrument. The results are reported in Panel B of Table 3. Although the potential effects from missing variables in the first-stage regression may increase when we move away from the narrow bandwidth around the regulatory threshold, the full sample IV analysis provides a validation test that allows for a more granular subsequent investigation. Consistent with the result in Panel A, we

TABLE 3

Options Availability and Financial Policy

Table 3 reports the second-stage results of the 2SLS regressions that examine the impact of options trading on underlying firms' financial policies. Panel A reports results for the RDD sample and Panel B reports the results for the full sample. OP is instrumented using the options listing eligibility dummy MEET. The dependent variables include EQISSUE, calculated as equity issuances minus equity repurchases divided by book assets at the beginning of the year; DTISSUE, calculated as long-term debt issuances minus changes in long-term debt divided by book assets at the beginning of the year; LEV, which represents the book leverage of assets, calculated as the book value of debt divided by book assets; REPO, calculated as the change in the number of shares outstanding divided by the number of shares at the beginning of the year; and DIV, or the dividend ratio, calculated as dividends divided by book assets at the beginning of the year. The control variables include ASSET, calculated as the natural logarithm of book assets; TANG, calculated as the ratio of fixed assets to book assets; MB, which represents the market-to-book value of assets ratio, calculated as the market capitalization plus the book value of debt divided by the book value of equity plus the book value of debt; and ROA, which represents the return on assets, calculated as net income divided by book assets. PRICE is the average daily stock price during the firm-year in natural logarithmic form. VOLUME is the total number of shares traded during the firm-year in natural logarithmic form. FLOAT is the number of shares held by noninsiders in natural logarithmic form. Standard errors are clustered by firm because options listing occurs at the firm level. Corresponding *t*-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

Panel A. RDD Sample

	EQISSUE	DTISSUE	LEV	DIV	REPO
	1	2	3	4	5
OP_{it}	0.054**	0.011+	−0.044**	−0.001	−0.001
	(5.46)	(1.88)	(−5.19)	(−1.20)	(−1.05)
$ASSET_{it-1}$	−0.074**	−0.021**	0.066**	−0.003**	0.001*
	(−9.90)	(−5.76)	(11.37)	(−3.13)	(2.54)
$TANG_{it-1}$	0.127**	0.055**	0.247**	−0.002	−0.004**
	(6.19)	(3.42)	(9.19)	(−0.60)	(−2.75)
MB_{it-1}	0.022**	0.006**	0.015**	0.001*	0.000
	(5.25)	(2.84)	(4.56)	(2.49)	(−0.99)
ROA_{it-1}	−0.247**	−0.024	−0.099**	0.006	0.007
	(−8.23)	(−1.44)	(−4.22)	(1.31)	(1.46)
AGE_{it-1}	0.001+	0.000	−0.002**	0.000	0.000
	(1.87)	(0.50)	(−3.92)	(−0.14)	(0.01)
$FLOAT_{it}$	0.002*	0.001*	−0.001	0.000	0.000
	(2.29)	(2.05)	(−1.51)	(−0.73)	(−0.81)
$PRICE_{it}$	0.006**	0.003**	−0.005**	0.001**	0.000
	(5.88)	(3.92)	(−3.78)	(3.96)	(−0.56)
$VOLUME_{it}$	0.000	0.000	0.000+	0.000	0.000
	(1.46)	(0.08)	(1.67)	(1.57)	(1.41)
Firm FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
No. of obs.	15,028	15,028	15,028	15,028	15,028
R^2	0.097	0.025	0.239	0.082	0.020

Panel B. Full Sample

	EQISSUE	DTISSUE	LEV	DIV	REPO
	1	2	3	4	5
OP_{it}	0.046**	0.015**	−0.047**	−0.001+	−0.006**
	(8.24)	(4.73)	(−8.05)	(−1.95)	(−2.62)
$ASSET_{it-1}$	−0.040**	−0.015**	0.042**	0.000	0.002
	(−11.40)	(−8.12)	(10.85)	(−0.45)	(1.21)
$TANG_{it-1}$	0.071**	0.026**	0.229**	−0.001	−0.003
	(5.93)	(3.15)	(10.40)	(−0.52)	(−0.65)
MB_{it-1}	0.025**	0.005**	0.012**	0.002**	−0.001
	(10.80)	(4.09)	(5.43)	(6.47)	(−0.76)
ROA_{it-1}	−0.209**	−0.041**	−0.112**	0.004+	−0.022**
	(−13.37)	(−4.08)	(−7.46)	(1.68)	(−3.91)
AGE_{it-1}	−0.000	−0.000+	−0.002**	0.000	−0.001**
	(−0.13)	(−1.67)	(−4.96)	(1.14)	(−5.30)
$FLOAT_{it}$	0.000	0.001**	0.000	0.000	−0.001**
	(0.42)	(2.74)	(−0.21)	(−0.40)	(−4.66)
$PRICE_{it}$	0.001*	0.002**	−0.003**	0.001**	0.000+
	(2.21)	(6.95)	(−4.22)	(6.22)	(1.70)
$VOLUME_{it}$	0.000	0.000	0.000**	0.000	0.000**
	(−0.49)	(−0.09)	(3.91)	(−0.79)	(3.01)
Firm FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
No. of obs.	37,994	37,994	37,994	37,994	37,994
R^2	0.059	0.020	0.183	0.089	0.008

find that the introduction of options has a positive and significant impact on equity issuance, as the coefficient on the instrumented OP is positive (0.046) and highly significant (*t*-stat = 8.24) in the full sample in Panel B. The estimated effect of options listings in the full sample is of similar magnitude to that in the RD sample. Given the first-stage result in column 5 of Table 2, meeting the options listing requirement is expected to increase equity issues by 0.66% of total assets (0.046 × 0.143).

An improved information environment allows firms to raise more capital through debt issues in addition to equity issues. We report the impact on debt issues in column 2 of Table 3, where the dependent variable is long-term debt issued minus long-term debt reduction scaled by total assets. The RD analysis in Panel A of Table 3 and full sample IV estimation in Panel B of Table 3 indicate a strong and positive impact of options availability on debt issues. The RD coefficient on OP (0.011, *t*-stat = 1.88) indicates that debt issuance increases by 0.1% of assets (0.094 × 0.011) after listing options due to meeting the SEC's options listing requirements on the margin. This causal impact has even stronger support from the full sample IV estimation, with a *t*-statistic of 4.73 on the instrumented OP in Panel B. The estimated economic impact of listing options is also slightly larger at 0.21% of total assets (0.015 × 0.143). Given the average debt issue-to-asset ratio of about 2 percentage points in the RD and full samples, these economic effects indeed reflect increases of 5% (RDD) and 10% (full sample) in the debt issuance intensity relative to the unconditional sample mean. The results in the first 2 columns of Table 3 indicate that firms are able to use external capital markets (both equity and debt) more easily when the underlying firm's equity becomes linked to exchange-traded options.

Another important implication is that equity issuance activity is more sensitive to changes in the information environment than is to debt issuance. Indeed, the results of both RD and full sample IV estimations indicate that equity issues are more responsive to options listings than are debt issues (columns 1 and 2 in Panels A and B of Table 3). The differential sensitivity of debt and equity to options listings should also result in a change in the debt-to-equity ratio, which we report in column 3 of Table 3. The dependent variable is the debt-to-assets ratio. The RD estimation result shows the clear impact of options trading in reducing financial leverage because the coefficient on the instrumented OP has a *t*-statistic of −5.19. This result implies that when options are listed as a result of meeting regulatory requirements, the treated firm's leverage falls by 0.41 percentage points (−0.044 × 0.094). Compared to the average leverage of 18.2 percentage points in the RD sample, this economic impact, which represents a reduction in financial leverage of 2.3%, is also significant. The full sample IV regression results consistently suggest that options availability reduces the debt-to-assets ratio by 0.67 percentage points (−0.047 × 0.143) with a *t*-statistic of −8.05 on the coefficient on OP. Consistent with prior literature (Frank and Goyal (2003)), we find that financial leverage is positively related to firm size and asset tangibility and negatively related to Tobin's Q and profitability.

We note two reasons that firms' payouts may be affected by the introduction of options and the associated decrease in information asymmetry. First, the direct effect of the reduction in information asymmetry, whether due to agency or

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signaling, suggests that dividends will be lower. Second, because firms are proactive in reducing leverage and because reducing payout also effectively reduces leverage, firms should reasonably reduce payout in addition to issuing more equity. We examine equity repurchases and dividends (both scaled by total assets) as dependent variables in columns 4 and 5 of Table 3, respectively. The RDD results in Panel A and the full sample IV results in Panel B paint a consistent picture that listing options reduces firm payout to shareholders in terms of dividends and repurchases, as the estimated coefficients on OP are all negative. These effects are statistically significant for repurchases (t-stat = −1.95) and dividends (t-stat = −2.62) in the full sample but weaker in the RD sample. In terms of economic significance, the RDD and full sample IV results suggest that when options are listed as a result of meeting regulatory requirements, share repurchase intensity is reduced by 0.63% to 5.05% and dividend payout is reduced by about 1%.

In summary, we find consistent evidence across multiple dimensions of firm behavior that the presence of listed options has an economically significant impact on firms' financial decisions. We present these effects' statistical and economic significance in Figure 1. Consistently across the RDD and full samples, options availability results in firms raising more external capital, especially in the form of equity, and paying fewer dividends, repurchasing fewer shares, and reducing their financial leverage. These results support the notion that as the introduction of options reduces information asymmetry and improves the underlying stock price efficiency, the cost of capital decreases, and particularly for equity. Furthermore, firms raise more capital and rely more on equity than on debt, which is consistent with the theoretical literature.

C. Market Response to Announcements of Financing Activities

We expect that increased information quality resulting from options availability will affect not only corporate policies but also the market perception of those policies given that their information content of those policies should change in the presence of listed options. Specifically, in an environment of reduced information asymmetry between firms and investors, the stock market reaction to the corresponding announcements of corporate actions should decrease in absolute magnitude. To conduct these tests, we collect information on the timing of corporate announcements. The nature of the test (market reaction) and data availability dictate that we focus on seasoned equity offerings (SEOs), repurchases, and announcements of dividend changes.

We use the SDC database for SEOs and open market repurchase announcements and CRSP for cash dividend announcements. We exclude SEOs with primary offerings of less than 1% of the market capitalization. Similarly, we exclude repurchase announcements with an authorized buyback value below 1% of the market capitalization. For dividend announcements, we focus on the dollar change in the announced dividend relative to the last cash dividend payout to gauge its informational value. After merging the announcement data with our panel sample reported in Panel A of Table 1, we obtained 928 SEOs, 2,601 repurchases, and 5,042 dividend changes announcements between 1996 and 2019.

For each announcement, we calculate CAR on day −1 to day 1 relative to the day of announcement. The risk-adjusted returns are estimated using the Fama and French (1993) factors plus a momentum factor in 63 trading days surrounding the event from day −31 to day 31. In line with existing studies (e.g., Hovakimian and Hu (2020) for SEOs, and Grullon and Michaely (2004) and Manconi, Peyer, and Vermaelen (2019) for repurchases), we find that the announcements of SEOs are associated with negative market reactions and that repurchase announcements induce positive reactions. Specifically, the 3-day CAR averages a negative 1.77% for SEOs and a positive 2.01% for repurchases. These figures provide a baseline for evaluating the effects of listed options on the corresponding announcement returns. We expect options trading to make the SEOs' CAR less negative and the repurchases' CAR less positive. Because dividend changes can be either upward or downward, we flip the sign of CAR for dividend decreases. As a result, the average CAR of 1.13% in our sample reflects the market impact of dividend changes.

We examine the impact of options availability on these announcement CARs by regressing CAR on the OP dummy instrumented using MEET in Table 4. For SEOs, we follow the literature mentioned previously and control for the size of the offering (SEO_SIZE), calculated as the number of shares issued divided by the existing number of shares. For open market repurchases, we control for the size of repurchase programs (REPO_SIZE), calculated as the announced number of shares

TABLE 4

Options Availability and Market Response to Financing Activities

Table 4 examines the impact of options trading on the market response to equity-related corporate actions, where OP is instrumented using the options listing eligibility dummy MEET. The sample includes corporate announcements in the merged sample of the SDC, Compustat, CRSP, and OptionMetrics databases between 1996 and 2019 after excluding financial firms and utilities. The dependent variables in the regressions are the abnormal cumulative returns (CAR) from days −1 to 1 relative to the announcement date. Columns 1–3 report the results for the event samples of seasoned equity offerings, open market repurchases, and dividend changes, respectively. For observations of dividend decreases, the sign of CAR is reversed to be consistent with the use of ΔDIV. MKTCAP is the market capitalization in billions of dollars. MB is the asset market-to-book ratio. RET_12M is the stock return for the 12 months preceding the announcement date. SEO_SIZE is the number of newly issued shares as a percentage of existing common equity. REPO_SIZE is the number of shares announced in the repurchase program as a percentage of existing shares outstanding. ΔDIV is the absolute change in the announced dividend relative to the last dividend payout scaled by the stock price on the day before the announcement. Corresponding *t*-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	$SEO\_CAR_{-1,1}$	$REPO\_CAR_{-1,1}$	$\Delta DIV\_CAR_{-1,1}$
	1	2	3
OP	1.006*	−1.143**	−0.320**
	(2.04)	(−5.53)	(−3.00)
MKTCAP	−0.008	−0.182*	−0.170*
	(−0.08)	(−2.05)	(−2.00)
MB	−0.063	0.067	0.203*
	(−0.30)	(0.31)	(2.30)
RET_12M	−0.385+	−1.922**	0.559*
	(−1.74)	(−3.93)	(2.34)
SEO_SIZE	−4.497*		
	(−2.01)		
REPO_SIZE		0.015**	
		(5.97)	
ΔDIV			0.388**
			(4.03)
Industry FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
No. of obs.	928	2,601	5,042

in the program divided by the number of existing shares. For dividend changes, we control for the absolute value of the announced dividend change scaled by the stock price on the day before the announcement (ΔDIV). In all models, we include market capitalization, market-to-book ratio, and the stock return in the 12 months preceding the announcement as well as year and industry FE.

For equity issuances, the estimated coefficient on OP is 1.006 with a *t*-statistic of 2.04 (column 1 of Table 4), thus indicating that the CAR of firms with options is less negative and is thus perceived by the market to contain less information. The first-stage result shows a coefficient on 0.36 on the IV in this sample. Taken together, options listing due to meeting the regulatory requirement increases the CAR by 36 basis points (1.006 × 0.36). The average SEO_CAR is −1.77% in our sample; therefore, options trading is estimated to offset 20% (0.36%/1.77%) of the negative market response to SEO announcements.

We also find a significant effect from options on repurchase announcement CAR in column 2 of Table 4. The coefficient on OP is −1.143 with a *t*-statistic of −5.53. With a first stage coefficient on 0.272 on the IV, options trading reduces the repurchase announcement CAR by 31 basis points, or 15% relative to its mean (2.01%). The analysis of market response to dividend changes in column 3 of Table 4 also shows that options availability reduces the market impact of corporate announcements. When we use listing eligibility to instrument options availability, the 2SLS regression result shows that the coefficient on OP is significantly negative (−0.32, *t*-stat = −3), thus indicating that the market response to dividend changes falls by 11 basis points on average, which is equivalent to 10% of the sample average market impact (1.13%). The evidence presented in this table shows that corporate announcements related to equity (issuance, repurchases, and dividend changes) elicit more muted stock market reactions post options listing. This finding indicates that the presence of options results in greater information efficiency even before those events are announced.

In summary, we find that post options listing, firms use more external equity to finance their operations, reduce their leverage, repurchase fewer shares, and pay lower dividends. Meanwhile, the stock market response to these corporate actions becomes significantly smaller at the time of public announcements. Both sets of results (firms' actions and the market reaction to them) support the premise that the presence of listed options results in less information asymmetry and in stock prices that contain more information.

D. Options' Impact on Investment Policies

Having examined the impact of listed options on corporate financial policies, we now turn to how corporate decisions affect the asset side of the balance sheet, namely, investment, cash retention, and innovation. The first important prediction we make is that better information quality will result in increased investment. We test this prediction in column 1 of Table 5. The coefficient on OP in Panel A of Table 5 indicates that, on average, options availability increases the ratio of capital expenditures to total assets by 1.2 percentage points with a *t*-statistic of 3.78 in the RD sample. Because meeting the options listing requirements increases the likelihood of listing by 0.093 in the first-stage regression (column 3 of Table 2), the

TABLE 5

Options Availability and Investment Policies

Table 5 reports the second-stage results of the 2SLS regressions that examine the impact of options trading on underlying firms' investment policies. Panel A reports results for the RDD sample and Panel B reports the results for the full sample. OP is instrumented using the options listing eligibility dummy MEET. The dependent variables include CAPX, or firm investment calculated as capital expenditures divided by book assets at the beginning of the year; CASH, calculated as cash and cash equivalents divided by book assets at the beginning of the year; PAT, or the number of patents calculated as the natural logarithm of 1 plus the number of successful patent applications during the year; and CITE, or patent citations calculated as the natural logarithm of 1 plus the average number of citations per patent for all patents applied for during the year. The patent data are obtained from the extended data set in Kogan, Papanikolaou, Seru, and Stoffman (2017). PRICE is the average daily stock price during the firm-year in natural logarithmic form. VOLUME is the total number of shares traded during the firm-year in natural logarithmic form. FLOAT is the number of shares held by noninsiders in natural logarithmic form. Standard errors are clustered by firm because options listing occurs at the firm level. Corresponding *t*-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

Panel A. RDD Sample

	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5
OP_{it}	0.012**	0.044**	−0.009**	0.027	0.097**
	(3.78)	(4.44)	(−2.86)	(1.54)	(3.28)
$ASSET_{it-1}$	−0.012**	−0.094**	−0.008**	0.005	−0.016
	(−6.93)	(−12.87)	(−6.06)	(0.36)	(−0.96)
MB_{it-1}	0.005**	0.025**	0.003+	0.004	0.01
	(5.21)	(6.20)	(1.93)	(0.79)	(1.14)
FCF_{it-1}	0.025**	0.046+	0.020**	−0.101*	−0.035
	(4.08)	(1.93)	(3.15)	(−2.55)	(−0.65)
AGE_{it-1}	0.000	−0.001	0.000	0.002	0.004+
	(1.19)	(−1.17)	(−0.44)	(1.46)	(1.94)
$FLOAT_{it}$	0.000+	0.002+	0.000	0.000	−0.001
	(1.81)	(1.70)	(1.40)	(−0.11)	(−0.19)
$PRICE_{it}$	0.002**	0.009**	0.002**	0.001	−0.003
	(5.77)	(7.07)	(5.69)	(0.23)	(−1.21)
$VOLUME_{it}$	−0.000+	0.001**	0.000	0.001*	0.000
	(−1.76)	(3.68)	(−0.41)	(2.49)	(0.27)
$OP_{it} \times MB_{it-1}$			0.004**		
			(3.00)		
Firm FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
No. of obs.	15,028	15,028	15,028	15,028	15,028
R^2	0.100	0.277	0.121	0.107	0.081

Panel B. Full Sample

	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5
OP_{it}	0.018**	0.055**	−0.006**	0.010	0.040*
	(7.99)	(8.59)	(−5.22)	(0.87)	(2.25)
$ASSET_{it-1}$	−0.009**	−0.061**	−0.002**	0.045**	0.051**
	(−7.73)	(−15.86)	(−3.21)	(6.50)	(5.58)
MB_{it-1}	0.004**	0.027**	0.003**	0.010**	0.021**
	(6.56)	(10.73)	(4.88)	(3.25)	(4.25)
FCF_{it-1}	0.034**	0.085**	0.031**	−0.057**	−0.056*
	(8.05)	(5.59)	(7.24)	(−2.99)	(−1.98)
AGE_{it-1}	0.000	−0.001**	0.000	−0.001	0.000
	(0.99)	(−2.82)	(−0.80)	(−0.64)	(−0.46)
$FLOAT_{it}$	0.000**	0.000	0.000*	−0.002**	−0.002**
	(3.12)	(−0.19)	(2.19)	(−4.31)	(−3.33)
$PRICE_{it}$	0.001**	0.003**	0.001**	0.002	0.000
	(5.21)	(4.08)	(7.09)	(1.47)	(0.17)
$VOLUME_{it}$	−0.000**	0.000	0.000	0.001**	0.000*
	(−2.62)	(0.51)	(0.16)	(4.84)	(2.40)
$OP_{it} \times MB_{it-1}$			0.003**		
			(6.14)		
Firm FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
No. of obs.	37,994	37,994	37,994	37,994	37,994
R^2	0.034	0.204	0.134	0.101	0.070

estimated economic impact of options eligibility is an increase in CAPX of about 0.11% of the total assets or 1.7% of the sample mean. Turning to the full sample IV estimator, we find that the 2SLS results in Panel B of Table 5 consistently indicate a strong and positive impact from options trading on investment because the coefficient on OP is again positive and highly significant (0.018, t-stat = 7.99), which implies that options listings that occur due to meeting regulatory requirements increases firm investment by about 4.2%.

Options introductions also allow firms build larger cash reserves for future investment because the resulting reduction in information asymmetry reduces agency issues and the adverse selection associated with larger cash reserves. We examine this implication regarding corporate cash retention policies in column 2 of Table 5, where the dependent variable is cash and cash equivalents scaled by total assets. Consistent with our conjecture, the options availability dummy OP has positive and significant coefficients in the RD analysis (β = 0.044, t-stat = 4.44) and full sample IV estimation (β = 0.055, t-stat = 8.59), which indicates that firms hold significantly more cash after options listings. This effect is economically significant. For example, the average treatment effect identified by the RD (full sample) result implies an increase in cash holdings by 1.9% (3.8%) of the sample mean due to meeting the options listing requirements.

Another important prediction regarding the real side of the balance sheet is that options introductions can also result in more efficient investment and heighten the responsiveness of capital expenditures to market valuations because the improved underlying stock price efficiency after options listing can facilitate managers' learning from the stock price and thus align managers' incentives. Following Chen et al. (2007) and Zhu (2019), we test the impact of options listings on firms' investment sensitivity to market valuations by estimating the following model:

$$\text{(2)} \qquad \begin{aligned} \text{CAPX}_{it} = {} & \alpha_{1t} + \alpha_{2i} + \beta_1 \times \text{OP}_{it} + \beta_2 \times \text{MB}_{it-1} + \beta_3 \times \text{OP}_{it-1} \\ & + \beta_4 \times \text{OP}_{it-1} \times \text{MB}_{it-1} + \Omega \times X_{it-1} + \in_{it}. \end{aligned}$$

In this setting, the coefficient β_2 reflects the baseline investment sensitivity to (lagged) market valuations in the absence of listed options. For our purposes, we are most interested in assessing whether the availability of options affects the sensitivity of firms' investment decisions to market valuations, which the coefficient β_4 of OP and MB (both lagged) should capture. The RD estimator in column 3 in Panel A of Table 5 shows clear support for this conjecture, as the coefficient on the interaction of MB_{t-1} and instrumented OP_{t-1} is significantly positive (0.004, t-stat = 3.00). This effect from options trading increases the investment sensitivity to stock market valuation because the standalone MB_{t-1} also has a positive and significant coefficient. These results suggest that the average investment-q sensitivity rises by 12.4% (0.004 × 0.093) relative to the baseline investment-q sensitivity, as identified by the coefficient on MB_{t-1} (0.003). The full sample regression result in Panel B is consistent with the RD result, and the interaction has a coefficient of 0.003 with a t-statistic of 6.14.

Finally, we evaluate the impact of options listings on a special type of investment (firm innovation activities) given its critical role in long-term growth.

Options listings should improve innovation outcomes due to more efficient stock prices reducing information asymmetry and better aligning managers' incentives. We use the number of patents and average patent citations (expressed in natural logarithm form) to measure the quantity and quality of innovation and report results in columns 4 and 5 of Table 5, respectively. The results indicate that the introduction of options significantly causes the number of citations to increase. The estimated coefficients on options availability, as instrumented by listing eligibility, are positive in the RD analysis (Panel A) and full sample IV estimation (Panel B) with strong statistical significance (t-stat > 2.2). However, the effects on the number of patents are consistently positive in both analysis but statistically insignificant.

In summary, we find that in addition to the impact on financial policies, options trading also has comprehensive effects on the asset side of the balance sheet. Figure 1 also presents these effects in terms of both their statistical and economic significance. Specifically, options trading allows the underlying firms to make larger investments, retain more cash, and improve the quality of their innovation. We also find that, in terms of investment-q sensitivity, options availability improves firms' investment efficiency. Overall, these results establish the positive spillover effect of options listings on firms' investment and financing decisions, which is consistent with both the agency and adverse selection theories.

E. Exclusion Restriction: Further Evidence of the Causal Relation

The validity of our instrument (options listing eligibility) is crucial to our identification of the effects of options trading. We established its relevance in the first-stage results of the 2SLS regressions in Table 2. In this subsection, we use two natural experiments to further examine the other, more elusive criterion of IV, namely, the exclusion restriction. Although the exclusion restriction could not be empirically proven, we believe that the results of these experiments justify the interpretation of our findings as a causal impact of options listing on corporate policies. The premises of the experiments are straightforward: the exclusion restriction is predicated on the claim that the instrument is correlated with the dependent variables only through options listing and not, for example, through firms' growth, size, maturity, investors, clientele, and so forth. If this claim does not hold, and the instrument explains the dependent variable not due to listing options but for other reasons (e.g., other correlated variables), it should have been equally valid prior to the existence of public options markets. This logic is the basis for our first experiment, detailed below. In the second experiment, we exploit a major change in the SEC's rule regarding the minimum stock price. Although the logic of these experiments applies to both the RD and full samples, we focus our tests on the full sample due to sample size constraints on the RD sample.[9]

[9]The RD sample is already restricted by the small bandwidth around the regulatory threshold. For example, as shown in Table 1, the RD restriction excludes 77% of the treatment observations from the full sample. Further sampling restriction renders a sample that is too small and can greatly reduce the power of the test.

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1. Establishing the Exclusion Restriction Using the Introduction of Public Options Markets

The first natural experiment is the emergence of public options markets that exogenously makes our instrument relevant. Meeting the SEC's options listing standards should only increase the likelihood of having listed options when options exchanges exist. Essentially, our instrument based on options listing eligibility should become relevant only after options exchanges are established. On the other hand, if our instrument violates the exclusion restriction, its impact on corporate policies should exist in periods prior to the existence of options exchanges. Therefore, the opening of the CBOE in 1973 is useful for examining whether a structural break occurs in the relation between our instrument and corporate policies. To conduct this test, we merge Compustat and CRSP data between 1961 and 1985.[10] To ensure that we test the same group of firms in this event, we exclude firms that exit the sample before or enter the sample after 1973. After removing firms with missing or negative assets and removing financial firms and utilities, we are left with 9,823 firm-years in the pre-CBOE period from 1961 to 1972 and 22,578 firm-years in the post-CBOE period from 1974 to 1995. For observations after 1973, we construct a dummy IV, HMEET, to indicate whether the firm-year satisfies the concurrent options listing rules regarding the stock price, trading volume, and public float.[11] For the period before 1973, we use the initial set of regulatory rules established by the SEC in 1973 to create HMEET; however, passing these thresholds does not signify any material effect related to the options market.

Because we do not observe the actual treatment of options listing before OptionMetrics began to track options data in 1996, we cannot run 2SLS regressions. Instead, we run FE regressions of corporate policies on HMEET in the two subperiods while controlling for the same covariates as well as firm and time FE. Table 6 reports the results. We first look at the pre-CBOE period between 1961 and 1972 in Panel A of Table 6. The results show that HMEET has statistically nonsignificant relations with all of the corporate policies in this period. These results clearly indicate that the instrument does not have the same impact on any of the corporate policies of interest prior to the establishment of options exchanges.

We next turn to the post-CBOE period from 1974 to 1985 in Panel B of Table 6. We find that the results in the post-CBOE period are largely consistent with our main findings. Specifically, HMEET has positive and significant coefficients for both equity and debt issues, and capital expenditures. Additionally, its coefficients are significantly negative on the debt-to-asset ratio, dividend payout, and repurchases. However, the results on cash levels, investment-q sensitivities, and two patent variables are consistently positive but statistically nonsignificant. That these results are slightly weaker than our main findings may result from the fact that we do not observe the actual treatment of options listing. The FE regressions of

[10]Using a shorter event window of only 5 years before and after the CBOE opening generates qualitatively similar results.

[11]Specifically, before 1982, the SEC required a minimum stock price of $10, total trading volume of 2 million shares in the past 12 months, and public float of 8 million shares for stocks to become eligible for option listing. The rules were then amended to those we use in the main sample in 1982.

TABLE 6

A Natural Experiment Using the CBOE Opening Event

Table 6 performs a test on the exclusion condition of our instrumental variable for options availability. The sample includes firms that exist in Compustat both before and after the CBOE opening in 1973. Table 6 reports the OLS regression results with firm and year fixed effects included. HMEET is calculated using the SEC's initial rules regarding the stock price, trading volume, and public float for the period before 1973. After 1973, the dummy HMEET is calculated using the concurrent options listing rules. Panels A and B report the results before and after the event, respectively. We only report the results for the variables of interest while controlling for the same variables as in Tables 3 and 5. Standard errors are clustered by firm because options listing occurs at the firm level. Corresponding *t*-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

Panel A. Before CBOE Opening: 1961–1972

	EQISSUE	DTISSUE	LEV	DIV	REPO	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5	6	7	8	9	10
$HMEET_{it}$	0.002	0.003	−0.007	−0.001	−0.005	0.005	0.002	0.006	0.043	−0.043
	(0.72)	(0.96)	(−1.38)	(−1.18)	(−0.96)	(1.02)	(0.53)	(0.77)	(1.20)	(−1.09)
$HMEET_{it} \times MB_{it-1}$								−0.001		
								(−0.13)		
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	9,823	9,823	9,823	9,823	9,823	9,823	9,823	9,823	9,823	9,823

Panel B. After CBOE Opening: 1974–1985

	EQISSUE	DTISSUE	LEV	DIV	REPO	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5	6	7	8	9	10
$HMEET_{it}$	0.009**	0.012**	−0.042**	−0.002**	−0.039**	0.010**	0.002	0.004	0.027	0.005
	(5.54)	(3.82)	(−8.97)	(−4.82)	(−3.79)	(3.94)	(0.60)	(0.63)	(1.44)	(0.16)
$HMEET_{it} \times MB_{it-1}$								0.004		
								(1.10)		
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	22,578	22,578	22,578	22,578	22,578	22,578	22,578	22,578	22,578	22,578

corporate policies on the instrument identify the intention-to-treat (ITT) effect in this setting, and these effects can be smaller than the treatment-on-treated (TOT) effects, as identified by the 2SLS regressions in our main results. In the early years of options trading, options exchanges have greater freedom in selecting stocks to list, given the large pool of eligible firms without options, which can further reduce the ITT effect of the instrument. Nonetheless, the sharp contrast of HMEET's relations with corporate policies before and after the CBOE opening supports our instrument's validity in terms of its compliance with the exclusion restriction.

2. Establishing Exclusion Restriction Using a Change in Listing Requirement

Our instrument, MEET, is a function of three listing requirements: i) price, ii) volume, and iii) public float. To further probe the possible impact of meeting the options listing standards through non-options channels (e.g., correlated omitted variables), the second experiment exploits a major change in the SEC's rule regarding the minimum stock price. In Jan. 2003, the minimum stock price required for options listing was reduced from \$7.50 to \$3. Therefore, among firms originally ineligible for options listing because their stock price was below \$7.50, those with a stock price above \$3 became eligible in 2003. These firms constitute the treatment group in our experiment. The rule change did not affect other ineligible firms whose

TABLE 7

Another Natural Experiment Using the SEC's 2003 Rule Change

Table 7 examines the effects of options trading on corporate policies using a change in the SEC's options listing standards, which occurred in 2003 and reduced the required price minimum from $7.50 to $3. The sample includes firm-years in the 5 years before and after the 2003 (i.e., 1998–2008). More specifically, we run a difference-in-differences analysis for firms with stock prices between $3 and $7.5, while firms with stock prices between $3 and $7.5 at the time of the rule change are treated (TREAT = 1) and firms with prices below $3 are the control group. POST is a time dummy equal to 1 for 2004–2008. We only report the results for the variables of interest while all regressions control for the same variables as in Tables 3 and 5. Standard errors are clustered by firm because options listing occurs at the firm level. Corresponding *t*-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	EQISSUE	DTISSUE	LEV	DIV	REPO	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5	6	7	8	9	10
TREAT × $POST_{it}$	0.031**	0.008+	−0.039**	−0.001**	−0.092**	0.012**	0.042**	0.008*	0.005	0.018+
	(4.10)	(1.65)	(−5.69)	(−3.57)	(−4.59)	(5.73)	(5.67)	(2.46)	(0.40)	(1.67)
TREAT × $POST_{it}$ × MB_{it-1}								0.002		
								(1.60)		
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	16,525	16,525	16,525	16,525	16,525	16,525	16,525	16,525	16,525	16,525

prices were below $3, therefore, we use these firms as the control group.[12] If the effects of options listing eligibility are obtained through the availability of options, we expect corporate policies between the treatment and control groups to be significantly different only after the rule change. To sharpen our identification of the impact of the minimum stock price rule change, we focus on those firms that meet the other two options listing requirements, namely, the trading volume and size of the public float, between 1998 and 2008 (i.e., 5 years before and 5 years after the rule change). For these firms, only the stock price determines options listing eligibility, and thus the rule change should have the largest impact. We construct a dummy, TREAT, to indicate whether the stock price is above $3 (i.e., treatment firms), and an event dummy, POST, to indicate the number of years after the rule change in 2003.

We perform DD estimations of the effects of the rule change on corporate policies and report the results in Table 7. Our focus is on the interaction of TREAT and POST, which identifies the treatment effect of the increased probability of listing options. We find that most of our previous findings also hold in this restricted sample. Specifically, we find that options trading significantly reduces financial leverage, repurchases, and dividend payout while increasing equity and debt issues, corporate investment intensity, cash holdings, and patent citations. The estimated coefficients are statistically significant at the 10% level or better in 8 models out of the 10 tests. The estimated effects on investment-q sensitivity and the number of patents are in the same direction as our main findings but statistically insignificant.

In summary, the evidence in both natural experiments significantly reduces the likelihood that our IV (i.e., eligibility for options listing) can affect the corporate policies of interest through alternative channels other than options trading. We show

[12]Firms with stock prices above $7.50 were also not affected by the rule change as they had always been eligible for option listing. We find that using these firms as an alternative control group leads to the same conclusion regarding the validity of option listing eligibility as an instrument for option availability.

that when the link between the instrument and options availability is absent, the instrument has either no effect or effects that counter our main results. These results indicate that the instrument satisfies the exclusion condition for the IV estimation and supports the causal impact of options trading on corporate policies.

F. Cross-Sectional Variation

The results to this point provide evidence that options trading directly impacts a wide range of firm policies, which is consistent with the role of options trading in reducing information asymmetry. To further ascertain the role of information, we examine how options listings affect firms with different degrees of information asymmetries. The channel we identify (the reduction in information asymmetry) posits that this shock to the information environment should have greater effects on firms with considerable information asymmetry. In addition, we also examine the conditional effects based on the firm life cycle because the literature has shown that firms' information asymmetry can also depend on their current stage of the life cycle (e.g., Grullon, Michaely, and Swaminathan (2002), DeAngelo, DeAngelo, and Stulz (2006)).

We rely on three proxies for quality of the firm information environment: financial analyst coverage, the probability of informed trading (PIN) from Easley, Kiefer, O'Hara, and Paperman (1996), and institutional ownership. These firm-level characteristics plausibly reflect three separate aspects of the information environment. First, analyst coverage affects the quality of information available to all investors, which will reduce information asymmetries between firms and investors as well as among investors. The impact of options trading on stocks with fewer analysts and who produce less information (e.g., Hong and Kacperczyk (2010), Merkley, Michaely, and Pacelli (2017)) is expected to be more pronounced. Specifically, options are expected to have a more significant effect on corporate decisions when few analysts follow the underlying firm.

We use institutional holdings as our second proxy for information asymmetry. Prior studies have established that institutions are likely to possess more information than individual stockholders and to be superior monitors (e.g., Shleifer and Vishney (1986)). Therefore, the benefits of reduced information asymmetry due to options trading may be lower for these firms. We expect that the real impact of options trading should be weaker for firms with high institutional holdings. We measure institutional ownership as the number of shares held by institutional investors as recorded in Thomson Reuters 13F database divided by the total number of shares outstanding.

The third proxy, PIN, reflects how actively investors acquire and exploit private information. Stocks with high PIN values are more prone to information asymmetry (e.g., Bharath, Pasquariello, and Wu (2009)). We expect that the effects of options trading would be larger for firms with high PIN values.

For firm life cycle, we use firms' equity market capitalization (i.e., the total number of common shares outstanding times the number of shares at the end of the year) and age (i.e., the number of years since the first data entry in Compustat) as proxies because mature firms tend to be old and large in the cross section we use. We then repeat our full sample IV estimations after augmenting the models with the

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information environment quality and firm life cycle measures as well as their interactions with the instrumented OP. We conjecture that the effects of options listings should be weaker when firms have more analyst coverage or higher institutional holdings, or when they are larger or older. This hypothesis implies that the sign of the coefficient on the interaction term of OP and analyst coverage (institutional ownership, size, and age) should be the opposite of that on the standalone OP. Similarly, the effects of options listings should be stronger for firms with larger PIN values, and the interaction term of OP and PIN should have the same sign as that on OP.

Panels A–E of Table 8 report the results of this analysis for analyst coverage, institutional ownership, and PIN as well as the life cycle proxies (i.e., equity market capitalization and age). Across the board, the 2SLS estimation results in Panel A of Table 8 are in line with our prediction. In particular, and consistent with our main findings, the presence of an options market linked to the firm equity leads to reduced firm leverage and stock repurchases as well as increased equity and debt issues, investment, cash holdings, and innovation outcomes. Critically, and consistent with our arguments, these effects become progressively weaker when firms have greater analyst coverage, as indicated by the coefficients on the interaction terms almost always having the opposite sign and being statistically significant at conventional probability levels in all cases except in that for debt issuance. Panel B of Table 8 shows that the effects of options trading are negatively associated with institutional ownership. Specifically, the interaction of OP and institutional ownership has significant coefficients whose signs are opposite to those on OP for equity issues, debt issues, book leverage, repurchases, investment, cash holdings, and patent citations. We do not find significant coefficients on the interactions for dividends, investment-q sensitivities, and number of patents. The evidence in Panel B of Table 8 shows that the baseline effects from options trading are generally larger for firms whose stocks have high PIN values. In particular, we find that the interaction terms between PIN and OP either amplify or completely absorb the previously documented baseline effects of listed options. Therefore, when the degree of information asymmetry among investors is high, the presence of an options market linked to the firm equity will lead to greater reductions in firm leverage and stock repurchases as well as to greater increases in security issues, investment intensity, cash holdings, and investment-q sensitivity. Our investigation of the conditional effects based on the firm life cycle in Panels D and E also shows results that are consistent with our expectations. Specifically, the effects of options listings are stronger for small (young) firms because the interactions of the OP dummy and firm assets (age) generally have statistically significant coefficients with signs opposite to those on OP in Panel D (E).

Overall, the results in Table 8 support the prediction that the effects of options listings on corporate decisions depend on the firm's information environment and maturity such that the effects are amplified in environments that are more opaque (i.e., with low analyst coverage and institutional ownership), when the amount of private information is high in capital markets (i.e., with high PIN values), and when the firm is less mature (i.e., are small and young). This outcome in turn supports the idea that firms are more likely to benefit from gains in the informational efficiency of underlying prices stemming from financial innovation and options listings,

TABLE 8
Cross-Sectional Variation in the Effects of Options Availability

Table 8 examines the cross-sectional variation in the effects of options trading on the firm's information environment. We repeat the IV estimations in Tables 3 and 5 and add an interaction of the options trading dummy (OP) and proxies for firm information quality to the model. In all tests, OP is instrumented using the listing eligibility dummy MEET. In Panel A, the proxy for information is analyst coverage, where ANALYST is calculated as the natural logarithm of 1 plus the number of analysts following the stock. In Panel B, the proxy for information is the proportion of shares outstanding held by institutional investors (HOLDINGS). In Panel C, the proxy for information is the probability of informed trading (PIN) as in Easley et al. (1996), estimated using daily order flow data in a year. In Panel D, the proxy for information is the firm size measured by total assets. In Panel E, the proxy for information is the firm age from the Compustat database. The same control variables as in the previous analyses are included in all regressions but omitted here for brevity. Standard errors are clustered by firm because options listing occurs at the firm level. Corresponding *t*-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

Panel A. Analyst Coverage

	EQISSUE	DTISSUE	LEV	DIV	REPO	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5	6	7	8	9	10
OP_{it}	0.047**	0.010*	−0.041**	−0.001	−0.012**	0.018**	0.052**	0.018**	0.013	0.052+
	(6.09)	(2.04)	(−5.15)	(−1.15)	(−3.55)	(5.69)	(5.83)	(4.58)	(0.76)	(1.80)
$OP_{it} \times ANALYST_{it}$	−0.016**	−0.001	0.006+	−0.001*	0.005**	−0.004**	−0.019**	−0.003+	−0.026**	−0.045**
	(−4.84)	(−1.02)	(1.81)	(1.96)	(3.12)	(−2.68)	(−4.90)	(−1.92)	(−3.13)	(−3.27)
$OP_{it} \times MB_{it}$								0.000		
								(0.43)		
$OP_{it} \times MB_{it-1} \times ANALYST_{it}$								−0.001		
								(−1.53)		
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	29,228	29,228	29,228	29,228	29,228	29,228	29,228	29,228	29,228	29,228

Panel B. Institutional Holdings

	EQISSUE	DTISSUE	LEV	DIV	REPO	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5	6	7	8	9	10
OP_{it}	0.028**	0.010**	−0.026**	−0.001+	−0.007**	0.014**	0.028**	0.013**	0.011	0.047*
	(5.72)	(3.21)	(−5.15)	(−1.85)	(−3.12)	(6.97)	(5.00)	(5.75)	(0.90)	(2.46)
$OP_{it} \times HOLDINGS_{it}$	−0.056**	−0.005	0.041**	0.001	0.013**	−0.022**	−0.060**	−0.019**	−0.026	−0.082+
	(−5.22)	(−0.72)	(3.09)	(0.48)	(2.69)	(−5.27)	(−4.66)	(−3.89)	(−0.91)	(−1.88)
$OP_{it} \times MB_{it}$								0.000		
								(0.92)		
$OP_{it} \times MB_{it-1} \times HOLDINGS_{it}$								−0.001		
								(−1.08)		
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No	No
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	30,447	30,447	30,447	30,447	30,447	30,447	30,447	30,447	30,447	30,447

Panel C. PIN

	EQISSUE	DTISSUE	LEV	DIV	REPO	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5	6	7	8	9	10
OP_{it}	−0.027**	−0.003	0.024**	−0.001	0.004	−0.002	−0.022*	0.007*	0.043*	0.045
	(−3.20)	(−0.67)	(2.81)	(−1.60)	(1.58)	(−0.80)	(−2.37)	(1.96)	(2.27)	(1.26)
$OP_{it} \times PIN_{it}$	0.301**	0.084+	−0.298**	0.001	−0.048+	0.084**	0.248**	0.016	−0.314+	−0.179
	(4.10)	(1.96)	(−4.15)	(0.23)	(−1.89)	(3.32)	(3.07)	(0.63)	(−1.94)	(−0.60)
$OP_{it} \times MB_{it}$								−0.004**		
								(−3.61)		
$OP_{it} \times MB_{it-1} \times PIN_{it}$								0.029**		
								(3.24)		
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	21,565	21,565	21,565	21,565	21,565	21,565	21,565	21,565	21,565	21,565

(continued on next page)

TABLE 8 (continued)

Cross-Sectional Variation in the Effects of Options Availability

Panel D. Firm Size

	EQISSUE	DTISSUE	LEV	DIV	REPO	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5	6	7	8	9	10
OP_{it}	0.213**	0.038**	−0.135**	−0.005*	−0.024	0.043**	0.212**	0.039**	0.021	0.095+
	(10.51)	(3.98)	(−8.78)	(−2.27)	(−1.34)	(7.29)	(10.27)	(5.46)	(0.71)	(1.84)
$OP_{it} \times ASSET_{it-1}$	−0.035**	−0.005**	0.019**	0.001*	0.002	−0.005**	−0.033**	−0.004**	−0.002	−0.012
	(−10.51)	(−3.21)	(7.64)	(2.10)	(0.60)	(−5.92)	(−9.94)	(−3.84)	(−0.50)	(−1.45)
$OP_{it} \times MB_{it}$								0.002		
								(1.26)		
$OP_{it} \times MB_{it-1} \times ASSET_{it-1}$								−0.001+		
								(−1.85)		
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	37,994	37,994	37,994	37,994	37,994	37,994	37,994	37,994	37,994	37,994

Panel E: Firm Age

	EQISSUE	DTISSUE	LEV	DIV	REPO	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5	6	7	8	9	10
OP_{it}	0.072**	0.019**	−0.063**	−0.003**	−0.012**	0.028**	0.069**	0.027**	0.007	0.051+
	(7.58)	(3.49)	(−6.61)	(−2.60)	(−2.87)	(7.63)	(7.00)	(6.34)	(0.38)	(1.81)
$OP_{it} \times AGE_{it-1}$	−0.003**	−0.001**	0.002**	0.0001**	0.000**	−0.001**	−0.003**	−0.001**	0.001*	0.001
	(−7.18)	(−2.99)	(5.50)	(2.58)	(2.67)	(−6.29)	(−6.55)	(−4.38)	(2.00)	(0.64)
$OP_{it} \times MB_{it}$								0.001		
								(0.87)		
$OP_{it} \times MB_{it-1} \times AGE_{it-1}$								−0.0001*		
								(−2.40)		
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	37,994	37,994	37,994	37,994	37,994	37,994	37,994	37,994	37,994	37,994

particularly when they operate in low-quality information environments or have poor governance mechanisms.

G. Intensive Margins

Using options listing requirements as our instrument allows us to establish the causal effect of the existence of listed options on a wide array of corporate policies. Having established this causal effect, we now shift our focus to another aspect of the impact of options market activity, namely, the intensity of options trading. Specifically, the effects of options markets on corporate policies will depend on the extent to which those options are traded and thus on the amount of private information revealed through options trading. While we recognize that identifying exogenous variations in options trading is arguably more challenging, we also believe that obtaining consistent results from properly designed tests allows us to cautiously draw causal inferences about the intensive margin effects of options trading.

We use options dollar volume to measure options activity (e.g., Roll et al. (2009)). Specifically, for each stock-day, we first calculate the premium on each options contract linked to the stock by multiplying its daily trading volume and the midpoint of the closing bid and ask prices. We then aggregate the dollar volume of all options contracts on the same stock-day and compute the average daily aggregate dollar volume of the same stock in a given year. To reduce the effect of outliers, we use the natural logarithm of 1 plus the average daily dollar volume as our measure of options volume (OPVOL) for a firm-year observation. In addition to trading volume, we also investigate the open interest of listed options contracts as an alternative measure of market activity. Open interest captures the risk exposure of investors and complements trading volume in describing market liquidity. Specifically, we construct the variable OPOI using the open interest on options contracts linked to the same stock following the calculation of OPVOL.

To test the intensive margin effects of listed options, we replace the options trading indicator (OP) with the corresponding continuous measures of OPVOL and OPOI in equations (1) and (2) and estimate these models using FE regressions with firm and year FEs to control for the same firm characteristics as in our main analysis. Because this analysis focuses on options trading intensity rather than listing events, we include all firm-years in the joint sample of OptionMetrics and Compustat after excluding financial firms and utilities and firms with assets below 1 million USD. The resulting sample has 71,983 observations.

Table 9 consistently shows that options trading intensity has an incremental impact on corporate policies. In Panel A of Table 9, the within-firm effects indicate that among firms with listed options, higher options trading volumes are associated with higher equity and debt issues, lower financial leverage, a lower level of repurchases activities, and lower dividend payout. At the same time, higher options trading volumes are associated with larger investments, higher cash holdings, higher investment-q sensitivity, and more intense innovation in terms of both the number of patents and patent citations. Specifically, all the coefficients on OPVOL have the same sign as those of the options availability dummy, OP, in Tables 3 and 5 for the same corporate policy. These relations are all statistically significant at conventional levels. Using OPOI as the measure of options market activity in Panel B of Table 9, we find that the results are largely the same as those obtained using options trading volume. Therefore, the evidence is consistent with the notion that in addition to options availability, active options trading further mitigates information frictions and affects underlying firms' decisions regarding financing and investment. However, although these results are consistent and informative vis-à-vis our main evidence on extensive margins, the resulting causal inferences should be taken with some caution given the nature of the empirical specification.

IV. Robustness

We conduct a battery of robustness tests to examine the sensitivity of our results. We summarize our findings here and, for brevity, report the detailed results in the Supplementary Material. In the first test, we experiment with alternative measures of the corporate policies. Specifically, we use the market value of assets instead of the book value of assets as the scalar for financial policy variables and add R&D expenses to capital expenditures to measure firm investment. We also examine alternative patent measures, including a patent dummy as well as patent originality and generality. Table A1 in the Supplementary Material shows that all of the results are qualitatively the same as our main findings.

TABLE 9

Intensive Margin Effects of Option Trading

Panel A of Table 9 reports the results of firm fixed effect regressions of corporate policies on various measures of options market activity. OPVOL is the options trading volume, calculated as the natural logarithm of 1 plus the average daily premium of all options contracts on the same stock (in thousands of U.S. dollars). Panel B reports the IV estimation that OPVOL is instrumented by OPOI, where options open interest is calculated as the natural logarithm of 1 plus the average daily option open interest of all options contracts on the same stock (in thousands of U.S. dollars). We only report the results for the variables of interest while all regressions control for the same variables as in Tables 3 and 5. Standard errors are clustered by firm because options listing occurs at the firm level. Corresponding *t*-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

Panel A. Fixed Effects Results

	EQISSUE	DTISSUE	LEV	DIV	REPO	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5	6	7	8	9	10
$OPVOL_{it}$	0.007**	0.004**	−0.005**	−0.001**	−0.002**	0.003**	0.009**	0.003**	0.001	0.056**
	(8.68)	(7.36)	(−5.71)	(−3.42)	(−5.87)	(9.64)	(10.36)	(8.46)	(0.33)	(11.72)
$OPVOL_{it} \times MB_{it-1}$								0.000		
								(0.99)		
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	71,983	71,983	71,983	71,983	71,983	71,983	71,983	71,983	71,983	71,983

Panel B. IV Estimation Results

	EQISSUE	DTISSUE	LEV	DIV	REPO	CAPX	CASH	CAPX	PAT	CITE
	1	2	3	4	5	6	7	8	9	10
$OPVOL_{it}$	0.007**	0.005**	−0.007**	−0.001**	−0.002**	0.003**	0.011**	0.003**	0.007+	0.066**
	(8.24)	(8.28)	(−6.24)	(−2.94)	(−5.61)	(8.58)	(10.58)	(8.39)	(1.74)	(11.86)
$OPVOL_{it} \times MB_{it-1}$								0.001		
								(1.12)		
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	71,983	71,983	71,983	71,983	71,983	71,983	71,983	71,983	71,983	71,983

Our second test employs several alternative estimation methods of the effects of options listings, including i) firm FE regressions in the full sample, ii) firm FE regressions using annual changes in all variables, iii) a sharp RD test assuming that all eligible firms are selected for options listing near the boundary, iv) a tighter bandwidth of 0.5 (vs. 0.6 in our main analysis) in the RD analysis, and v) a wider bandwidth of 0.7 in the RD analysis. The results in Table A2 in the Supplementary Material confirm that our conclusions hold in these tests.

In the third test, we address the concern that managers may deliberately affect their firms' options listing eligibility via the stock price criterion. For example, they could potentially use reverse splits to increase the stock price to have the exchange introduce options on their firm (and splits if they do not want options). For this purpose, we exclude 1,630 stock splits and reverse splits from the sample and replicate the analysis. Our results in Table A3 in the Supplementary Material indicate that stock splits have no impact on our findings.

Our primary identification method in the study relies on an RDD analysis based on an IV for the regulatory requirements for options listing. As an alternative, we also use PSM to further examine the possibility that our results are driven by omitted variables related to the eligibility standards rather than to options listing. Specifically, using a matched sample, we compare firm policies before and after the options introduction (first difference) with otherwise similar firms (second difference). Table A4 in the Supplementary Material shows that the DD results are largely the same as our main findings for all corporate policies and that the pretreatment trend between the treated and control groups is statistically nonsignificant, thus reinforcing the causal impact of options trading.

Overall, our extensive empirical investigation using an RDD approach based on an IV related to the exchange decision to list options, 2 natural experiments that verify that the IV results are indeed due to listing options and not to omitted variables correlated with the attributes, and a PSM approach all lead to the same conclusion: that the presence of listed equity options has a direct causal impact on a host of corporate policies, which is consistent with the effect of reduced information asymmetry.

V. Conclusions

We examine the causal effect of changes in information quality associated with options trading on a wide array of firms' financing and investment decisions. To draw causal inferences, we rely on an RDD based on an IV along with natural experiments and DD estimators to exploit regulatory options listing requirements that exogenously affect the likelihood of having listed equity options.

Taken together, our results provide a coherent picture of how improved information quality affects corporate policies. Firms experiencing positive information shocks brought about by options trading will more frequently access external capital, and their behavior is consistent with the notion that external equity becomes cheaper than debt for these firms. Furthermore, these firms actively manage their capital structure to achieve lower financial leverage and concomitantly conduct fewer equity buybacks and pay lower dividends. Meanwhile, investors react less intensely when these firms decide to raise or retire equity or to change their dividend payout. With regard to the asset side of the balance sheet, firms with listed options invest more, innovate more, and build larger cash reserves. Concurrently, we find evidence that corporate investments become more sensitive to growth opportunities, as measured by Tobin's Q. Consistent with private information driving the results, the effects of listed options are stronger in weak information environments characterized by low analyst coverage, a high probability of informed trading, low institutional ownership, small market capitalizations, and younger firms. In addition to options availability, we find similar information spillover effects from options trading intensity as intensive margin effects.

Our evidence—based on a large sample of events in readily observable public exchange markets, staggered over a long study period, and derived from a variety of identification methods—demonstrates that information shocks from innovations in the capital markets have a causal and significant impact on corporate behavior. Overall, we conclude that by rendering underlying equity prices more informative and facilitating monitoring by investors, options market activity feeds back into and enhances the efficiency of firms' decisions. Our evidence clearly shows that option markets are not a sideshow.

Our findings also highlight the significance of spillover effects when evaluating the real effects of derivatives markets and perhaps of financial innovations

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more generally. Even though firms rarely engage in trading on their own options, we show that they still benefit from their introduction. Our evidence points to the positive externalities gained by firms and their shareholders. Lerner and Tufano (2011) note that such externalities are critical to measuring the social welfare of financial innovations. To the extent that similar spillover effects are relevant to other financial innovations such as credit derivatives or new trading systems, we advocate including these effects in analyses of their merits.

Supplementary Material

To view supplementary material for this article, please visit http://doi.org/10.1017/S0022109023001229.

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