The Causal Effects of Short-Selling Bans: Evidence from Eligibility Thresholds

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Abstract

We identify the causal effects of short-selling bans on stock prices and market quality using regression discontinuity (RD). We exploit three threshold-based rules that determine a stock's short-selling eligibility on the Hong Kong Stock Exchange. Short-selling bans do not affect stock prices or market quality despite affecting short-selling volume at all thresholds. Stock returns, volatility, bid-ask spreads, and crash risk are not statistically or economically different for banned vs. unrestricted stocks when appropriate counterfactual stocks are used to measure a ban's effects. Our findings suggest that short-selling bans are not as costly as previously argued, but are ineffective at reducing volatility or buttressing prices.

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1. Introduction

Regulation of short sales are among the oldest and most controversial regulations of financial markets. Short-selling bans are almost as old as the first publicly listed common stocks, as exemplified by the shorting ban of the Dutch East India Company in the early 17th century. The widespread adoption of short-selling bans during the recent financial crisis suggests that policy makers still view the prohibition of short sales as an important regulatory tool. Proponents of these restrictions argue that they stabilize markets by reducing volatility and preventing market declines. Opponents argue that bans adversely impact liquidity and price discovery. Despite the long history and frequent changes to short-selling regulations around the world, there is still little consensus about the effect of short-selling eligibility on stock markets.

The lack of a clear consensus is not surprising given the difficulties researchers face when studying the effects of short-selling regulations. Governments and regulatory agencies often design new regulations in response to changes in market conditions or the political environment. This biases estimates of the true effects of short-selling regulation. Therefore, it is not clear whether the estimated treatment effects documented in the existing literature are due to the implementation of the regulation or to differences between stocks, countries, or time periods.

In this paper, we identify causal effects by exploiting exogenous variation in short-selling bans using regression discontinuity (RD). Our analysis relies on three threshold-based rules that determine short-selling eligibility on the Hong Kong Stock Exchange (HKEX). Each quarter, firms are eligible to be shorted if they satisfy cutoffs related to public float, size, and turnover. For firms very close to a given threshold, falling to one side or the other of the cutoff is largely due to chance, providing plausibly exogenous variation in short-selling eligibility.

We find that short-selling bans bind. Short-eligible stocks experience discontinuously higher short-selling activity around each threshold, both economically and statistically. The discontinuities in short-selling activity are up to 20% (40%) of the mean (median) short-selling activity for all short-eligible firms in Hong Kong. Despite this, we find that these short-selling bans have no effect on stock prices or market quality. Stock returns, volatility, bid-ask spreads, and crash risk are not statistically or economically different for banned vs. unrestricted stocks. Our conclusions do not result from low-powered tests. The signs of the coefficient estimates differ across the thresholds for all outcome variables, suggesting inference would be unchanged even with lower standard errors. Additional tests using intraday data to measure stock market quality result in similar conclusions.

Our findings are consistent with rational expectations models that suggest short-selling bans have no effects on price levels (e.g., Diamond and Verrecchia (1987)). Our conclusions contradict Chang, Cheng and Yu (2007), whose finding that short-sale eligibility results in price declines supports the overpricing predictions of Miller (1977).[1] They use all stocks added to (but not deleted from) the short-sale eligibility lists in Hong Kong prior to adoption of threshold-based eligibility rules. The differences in our results stem from the fact that, away from the threshold, short-sale eligibility in Hong Kong may be endogenous to future returns. This is evident in large price run-ups for firms added to the short-sale eligibility list in Chang et al. (2007). Indeed, we reproduce their findings in our sample period as well. Our research design eliminates this bias by comparing market outcomes for similar firms that happen to be on either side of the short-sale eligibility thresholds.

Establishing causality between regulation and stock market outcomes is particularly important in the context of short-selling bans with which stock market regulators attempt to eliminate speculative and potentially destabilizing short sales. Boehmer, Jones and Zhang (2013) find that stock price levels are not affected by the 2008 SEC ban in U.S. stock markets, but that stock market quality is heavily affected for all but the smallest firms. Using a cross-country setting, Beber and Pagano (2013) find a negative effect of short-selling bans on market quality and price discovery, as well as an increase in stock market volatility in 30

[1] We discuss theoretical predictions in Section 2.

countries that imposed short sales restrictions during the financial crisis. Unlike Boehmer et al. (2013) who find strongest effects for large stocks, Beber and Pagano (2013) find that the strongest results are concentrated in smaller capitalization stocks. Bris, Goetzmann and Zhu (2007) analyze an international panel of short-selling regulations and find that countries without short-selling restrictions exhibit improved price efficiency, but more negative market return skewness. These studies document important regularities concerning short-selling eligibility, but they also acknowledge the empirical difficulty of disentangling the effects of short-selling eligibility from extreme stock market conditions and non-random selection of stocks, time periods, or countries for regulation.[2]

Our results regarding price levels are consistent with both Boehmer et al. (2013) and Beber and Pagano (2013) who find no effect of bans on stock returns. However, our conclusions differ markedly from both papers with regard to volatility and market quality. We find no effect, while they find that bans harm volatility, albeit for stocks of different sizes (large stocks in Boehmer et al. (2013); small stocks in Beber and Pagano (2013)). Our finding that short-sale bans do not diminish market quality suggests their use by regulators is less costly than previously argued, if ineffective at reducing volatility or buttressing price levels.[3]

Our setting has several distinct advantages relative to previous studies of short-selling bans/eligibility. First, RD is a quasi-experimental framework that helps alleviate concerns about endogeneity that other studies on short-selling bans face. Under the assumption of local continuity of potential outcomes, we can interpret our local average treatment effects causally (Roberts and Whited, 2013). Second, our estimation is based on a panel dataset exploiting 52 quarterly updates to the short-selling eligibility list. This panel covers more than a decade of data and spans periods of bull and bear markets, thus allowing us to test the effects of bans under varying economic conditions. Finally, we can estimate the average treatment effects of a short-selling ban/eligibility across three different thresholds, which

[2]We discuss the identification strategies employed in these papers in Section 4.

[3]The unexpected nature of the U.S. ban had unintended consequences related to regulatory uncertainty as documented in Battalio and Schultz (2011) for equity option markets.

demonstrates the robustness of our findings across different size groups.[4]

The rules we exploit allow for unbiased estimates of treatment effects on the sixth largest stock market in the world (HKEX). The HKEX has around 1,500 ordinary stocks, almost 700 of which were eligible for short selling as of 2013 year end. Short-selling activity is relatively important in Hong Kong, averaging 9% of traded volume during 2013 conditional on eligibility and non-zero shorting.[5] Previous studies of short selling in HKEX find strong effects of short sales on stock prices and other outcome variables (Chang et al. (2007), Massa, Qian, Xu and Zhang (2014a), and Massa, Zhang and Zhang (2014b)).[6]

Our work is related to a number of studies that investigate the causal effects of altering costs associated with short selling. Diether, Lee and Werner (2009), Alexander and Peterson (2008), and Grullon, Michenaud and Weston (2015) study the Regulation SHO randomized experiment conducted by the SEC from 2005 to 2007. The experiment repealed the uptick rule for a set of Pilot stocks. Kaplan, Moskowitz and Sensoy (2013) study the effects of a lending shock to randomly selected stocks of a large portfolio manager. Arnold, Butler, Crack and Zhang (2005) show that an increase in shorting costs due to a tax law change increased the information content of short interest. These studies focus on changes in short-selling restrictions that affect the costs of selling short while we focus on short-selling bans.[7]

The remainder of the paper is organized as follows. Section 2 briefly reviews the theoretical predictions of the effects of short-selling eligibility. Section 3 describes our empirical setting and data. Section 4 replicates previous findings concerning short sales in Hong Kong

[4]The mean market capitalization differs across thresholds. For example, the mean market capitalization for the market capitalization threshold sample is about HK$840 million while it is much higher (over HK$2,000 million) for the turnover velocity threshold sample.

[5]Our short volume data does not include short sales by market makers, which may understate the magnitude of observable shorting activity in Hong Kong relative to the U.S.

[6]On the other hand, Hong Kong has also been used to show what short selling cannot explain such as the weekend effect (Gao et al., forthcoming).

[7]Theory suggests that the effects of cost changes may differ from those of eligibility, even beyond differences in economic magnitudes (Diamond and Verrecchia, 1987). In their model, bans and restrictions affect the relative composition of informed and uninformed investors differently. Short-selling bans eliminate short sales by both informed and uninformed traders. In contrast, short-selling restrictions decrease short-selling by relatively uninformed investors more than short-selling by informed investors.

and discusses endogeneity and identification strategies in studies of short-sale bans/eligibility. Section 5 examines the causal effects of short-selling on stock market outcomes using regression discontinuity. Section 6 concludes.

2. Theoretical Predictions of Short-sale Eligibility

Theoretical predictions of the effect of short-selling bans on stock prices are mixed. Miller (1977) argues that if investors have differences of opinion about a stock's value, the price will reflect the optimistic investors' valuation in the absence of short-sale eligibility. Thus, short-sale bans help support prices by eliminating pessimistic investors' shorting activities. In the absence of short selling, security prices should be overvalued. Therefore, allowing short-selling should lead to stock price declines. The intuition of Miller is formalized in Chen, Hong and Stein (2002), which relates differences of opinion to breadth of ownership. In their model, securities can become overvalued even if only a subset of investors is constrained from short selling. Hong and Stein (2003) add another prediction of the short-sale constraints under differences of opinion. In their model, market crashes can result from short-sale constraints as negative information fails to be gradually incorporated into prices. Unrevealed bad news accumulates until previously optimistic investors abandon the market, leading to large negative price adjustments. The model predicts short-sale prohibitions to be associated with a greater prevalence of extreme downward price movements.

Diamond and Verrecchia (1987) model short-sale bans in a rational expectations framework. Because market makers account for the availability of short-selling in valuing a stock, the stock is not overvalued in their model. However, overvaluation due to short-sale prohibitions is possible even in rational models if investors value the option to sell to another trader with a different expectation of value as in Harrison and Kreps (1978). For instance, differences of opinion on asset valuation arise due to overconfidence in Scheinkman and Xiong (2003). In their dynamic setting, traders are willing to purchase a security for a price above their valuation in the hopes of selling to another buyer, generating a bubble under short-sale restrictions. The bubble is accompanied by increased volume and volatility. If short-sale restrictions prevent these bubbles, short-sale bans should reduce volatility and volume. Allen, Morris and Postlewaite (1993) show that overpricing can result from short-sale bans as a result of private information in a finite-horizon rational expectations equilibrium if agents do not know other agents' beliefs. This allows the backward induction argument eliminating overpricing to fail. Therefore, short-sale eligibility may push stock prices down and increase volume and volatility, even in rational expectations models of short-selling bans/eligibility.

Theoretical predictions of the effect of short-selling bans on market quality are also mixed. A short sale ban could decrease spreads if short sellers are informed about the fundamental value of a stock and trade competitively on negative information (the information effect). Diamond and Verrecchia (1987) predict that short-sale prohibitions affect the speed of information incorporation into prices. Short-sale bans reduce the speed of price discovery since some short-sellers are privately informed. The reduced information asymmetry leads to reduced bid-ask spreads under a short-sale ban. Diamond and Verrecchia (1987) is a sequential trade model in the spirit of Glosten and Milgrom (1985). As such, liquidity provision by short-sellers is outside the model. Boehmer et al. (2013) point out that a short-sale ban could increase spreads if short sellers compete to provide liquidity (the liquidity provision effect). Therefore, short-sale eligibility has a theoretically ambiguous impact on market quality. If the liquidity provision effect offsets the information effect, the result is no changes in market quality.

Our empirical tests focus on the pricing and market quality implications of these models. In particular, we test whether predictions of overpricing or increased volatility are borne out in the data. We also test whether bans affect other aspects of the trading environment such as volume and liquidity. We argue that identification using exogenous variation in short-sale eligibility is key to testing these theoretical predictions.

3. Empirical Setting and Data

3.1. Short Selling in Hong Kong

HKEX initially allowed short-selling in 1994 for 17 stocks designated Pilot Stocks. Over the following years, HKEX modified the eligible list 11 times to eventually include over 200 stocks by the end of 2000. These initial list changes were somewhat haphazard and dependent on market conditions; see Chang et al. (2007) for details.

By 2001, HKEX utilized several threshold-based rules, among other requirements, for inclusion to the Designated Securities List. Specifically, stocks were eligible to be shorted if they satisfied any of the following criteria:

1. all constituent stocks of indices which are the underlying indices of equity index products traded on the Exchange;
2. all constituent stocks of indices which are the underlying indices of equity index products traded on Hong Kong Futures Exchange Limited;
3. all underlying stocks of stock options traded on the Exchange;
4. all underlying stocks of Stock Futures Contracts (as defined in the rules, regulations and procedures of Hong Kong Futures Exchange Limited) traded on Hong Kong Futures Exchange Limited;
5. stocks that meet the minimum liquidity requirement for the issuance of basket derivative warrants (i.e. market capitalization of public float of not less than HK$1 billion, being maintained for the 60 days qualifying period);
6. stocks with market capitalization of not less than HK$1 billion and an annual turnover to market capitalization ratio of not less than 40%;
7. Tracker Fund of Hong Kong and other Exchange Traded Funds approved by the Board in consultation with the Commission; and
8. all securities traded under the Pilot Programme.

Effective July 3, 2012, HKEX altered the sixth eligibility requirement. In particular, they increased the market capitalization requirement to $3 billion from $1 billion and increased the turnover-to-market capitalization ratio (henceforth, turnover velocity) requirement to 50% from 40%.

In this study, we exploit the thresholds identified in the fifth and sixth eligibility requirements to identify exogenous variation in the probability of short sale eligibility.

3.2. Sample construction

The current list of short-sale eligible securities is available on the Hong Kong Stock Exchange website. To construct the historical list, we start with the designated security list as of November 5, 2014 and work backwards in time using additions and deletions. Firms being added or removed from the short sale eligibility list are identified in periodic HKEX press releases.[8] Press releases are available online for all quarterly evaluations since 2001. We hand collect these additions and deletions to create the history of the eligible security list.

Our daily data on HKEX stocks is from Bloomberg and includes information on prices, returns, market capitalization, shares outstanding, float outstanding, total volume, short volume, and bid/ask prices. We have quarterly list additions and deletions from HKEX's press releases beginning in 2001, so the sample runs from 2001 to 2014. Float data is only available from 2006 on; therefore, tests using the float-adjusted market capitalization use data from 2006 to 2014. We also use intraday data from Thomson Reuters Tick History in additional robustness tests that yield similar results.

In general, HKEX evaluates the Designated Securities List on a quarterly basis. However, the evaluation is not conducted at regular intervals, nor does HKEX disclose the date on which eligibility is determined (i.e., when market capitalization is evaluated as above or below the threshold).[9] To construct the thresholds, we evaluate whether a firm has met a

[8]One exception is for deletions due to acquisition or delisting. We manually correct for this by identifying any firms not on our historical list as short-sale eligible if they experience shorting volume at any point over the sample (prior to acquisition or delisting).

[9]In order to correct for the associated noise in the measurement date, we employ a fuzzy regression discontinuity approach described in Section 5.

given threshold using end-of-month data from two months prior to the month of the updated list's effective date. For example, we use data as of the last trading day of March 2005 (which we call the measurement date) for the May 17, 2005 effective date. We use the minimum float-adjusted market capitalization over the 60 trading days preceding the measurement date to evaluate if the firm met the basket derivative warrants threshold. For turnover velocity, we use the aggregate dollar volume traded over the 365 calendar days preceding the measurement date, divided by the market capitalization as of the measurement date.[10] Market capitalization is the closing market value as of the measurement date.

We evaluate all quarterly changes to the Designated Securities List over 2001 to 2014. To avoid confounding effects of other reasons for list inclusion, our analysis in Section 5 excludes all firms that are members of various indices or that are the underlying securities for options or futures.[11] We also exclude from our analyses any time windows in which a stock appears not to trade, as evidenced by a return standard deviation of exactly zero.[12]

3.3. Summary statistics

Table 1 reports summary statistics of stock-quarter observations for various subsamples of interest on the Hong Kong Stock Exchange. Panel A reports statistics for all stocks. For each firm-quarter, the float-adjusted market capitalization, turnover velocity, and market capitalization are month-end values, measured two months prior to a given quarterly effective date. These variables are the threshold values determining short-sale eligibility for a particular effective quarter and are measured as discussed above. Average daily returns, standard deviation of daily returns, and short volume as a fraction of total volume (RELSS) are measured over the 250 trading days preceding a quarterly effective date.

[10] To be included in our analysis, we require a firm trade on at least 200 trading days over the annual window.

[11] We exclude all firms that are member of the Hang Seng, Hang Seng Composite, Hang Seng LargeCap, Hang Seng MidCap, Hang Seng SmallCap, and the Hang Seng China Enterprise indices. We obtain historical index constituent lists from the Hong Kong Stock Exchange website.

[12] Our results are not sensitive to this screen, and we explicitly test whether non-trade days are a function of short-sale eligibility.

Panel B and C report statistics for firms that are currently ineligible or eligible for shorting, respectively. It is immediately clear that firms that are eligible for shorting appear quite different from those that are not eligible. Eligible stocks are dramatically larger firms, which is not surprising given that size is an eligibility requirement. Not surprisingly, smaller, short-ineligible stocks have higher return volatility than the larger, short-eligible securities. Finally, since firms can be removed from the short-sale eligibility list, some of the deleted securities may have experienced shorting activity over the preceding year. However, the vast majority of the short ineligible securities have experienced no shorting activity over the year preceding a given quarterly effective date. For short-sale eligible securities, the average (median) amount of shorting as a fraction of total volume is 4.6% (2.5%).

4. Endogeneity of Short-Sale Eligibility

4.1. Hong Kong Additions

The paper most closely related to ours is Chang et al. (2007). They find that stocks that are added to the Hong Kong short-sale eligibility list experience negative abnormal returns, increased volatility and prevalence of extreme negative returns, and less positive skewness in returns subsequent to being added to the eligibility list. Their sample period runs from 1994 to 2003; the threshold rules we analyze in Section 5 were in effect only at the tail end of this sample.

Our sample period of 2001 to 2014 contains a total of 1,528 addition events. In Table 2, we present return characteristics for the 90 (91) trading day window preceding (following) the event date. Subsequent to becoming eligible for shorting activity, firms being added to the list experience statistically and economically large negative abnormal returns. The cumulative abnormal return (in excess of the HK value-weighted market return) is almost -6%. The average daily abnormal return is -7 basis points per day. This confirms the tests in Chang et al. (2007), who find cumulative negative abnormal returns in a 60 day window of

-4%.[13] However, firms being added to the list also have significant abnormal returns in the 90 days *preceding* the addition effective date. The average daily return is 12 basis points; the cumulative return is a sizable 11%. These results highlight the fact that selecting all firms that have been added to the Designated List creates a sample selection bias that confounds inference. This makes it difficult to distinguish the effects of short-selling eligibility from those of recent abnormal positive performance followed by subsequent reversals.

Results from our sample period differ from those in Chang et al. (2007) with respect to return volatility. Using one-year windows, they find that return standard deviation is significantly higher subsequent to list inclusion.[14] In our sample, we find that daily return volatility is actually 49 basis points lower in the 91 days following short sale eligibility. One possible explanation for these findings is the Asian crisis. Out of the 519 addition events studied by Chang et al. (2007), 129 occur on May 1, 1997, so the one-year window following this addition date includes the height of the Asian financial crisis.

Finally, we confirm in our sample the results of Chang et al. (2007) for skewness and the prevalence of extreme negative returns (defined as the number of days with a return more than two standard deviations below the mean for a given period). Return skewness is less positive following the initiation of short-sale eligibility. This difference is statistically significant. Moreover, the prevalence of extreme negative returns increases from 1.6% to 1.8% following addition to the short-sale eligibility list.

With the exception of return volatility, the results of Chang et al. (2007) hold up remarkably well in our out-of-sample replication. However, the results also highlight the potential pitfalls in studying all firms added to the eligibility list. In general, additions to the list are not random. There is substantial (positive) pre-treatment in returns preceding a firm's addition to the short-sale list suggesting reversals, not short-selling eligibility, may be responsible for subsequent negative returns.

[13] This corresponds to the Market-Adjusted Model results in Panel B of their Table 2.

[14] See Table 6 of Chang et al. (2007).

4.2. Identification Strategies for Short-Sale Bans

Several recent studies use time-series and cross-sectional variation in regulatory environments to study the effects of short-selling bans on market quality and asset pricing. Boehmer et al. (2013) find that a number of measures of market quality (spreads, price impacts, and intraday volatility) are substantially worse for firms in which shorting was banned in the United States in late 2008. Similarly, using variation in short-sale restrictions across countries and time, Beber and Pagano (2013) find that short-selling bans hurt liquidity, slowed price discovery, and did not support prices. Bris et al. (2007) also use an international panel to identify the effects of short-sale restrictions. They find that countries without short-sale restrictions exhibit more efficient prices and more negative skewness in market returns, but restrictions are unrelated to the distribution of individual stock returns.

These papers document important empirical regularities related to short-selling regulation. They also discuss and use various strategies to deal with the endogenous adoption of short-selling bans. Boehmer et al. (2013) use a difference-in-difference approach to identify the effects of the short sale ban, matching banned securities to non-banned securities on listing exchange, the presence of listed options, market capitalization, and dollar volume. They also evaluate an industry-matched subsample, but the sample is quite small given that the vast majority of banned stocks were financial securities (and practically all financials were banned). The identifying assumption is thus that there are no omitted, unobserved differences between banned and unbanned securities that are also related to return or market outcomes (except through the effect of the ban). The plausibility of this assumption is difficult to judge as the exclusion restriction is fundamentally untestable. However, the ban was almost exclusively applied to financial sector securities (or companies with strong financial sector segments).[15] As a result, as discussed in their paper, there is a limited ability to match on industry between treatment and control firms. Many papers find strong relation-

[15]Additions to the short-sale list subsequent to the initial banned list were made at the discretion of exchanges. Listed firms could request addition (or deletion) from the list, resulting in yet another source of potential endogeneity.

ships between industry and stock market outcomes (e.g., Moskowitz and Grinblatt (1999), Hou and Robinson (2006), and Hou (2007)); it is therefore not clear that the short-selling ban is the only difference between the treatment and control samples. Put simply, it is hard to identify treatment effects of short-sale eligibility versus possible industry differences.

Beber and Pagano (2013) also acknowledge potential identification issues. They instrument for the likelihood of a ban in a first stage using two possible instruments: the lagged value of a country-level financial CDS spreads or the lagged value of the country's financial stress index proposed by Tytell, Elekdag, Danninger and Balakrishnan (2009). In order to satisfy the exclusion restriction, these lagged financial variables need to be uncorrelated with subsequent bid-ask spreads, except through their effect on the probability of a ban. Again, this assumption is difficult to test. However, other work suggests possible lead-lag relationships between aggregate price levels and market quality measures such as liquidity through channels other than short-selling. For instance, the findings of Chordia, Sarkar and Subrahmanyam (2005) suggest that aggregate Treasury spreads may be related to subsequent market quality through monetary policy or mutual fund flows. The instruments could therefore affect outcomes through these channels rather than short-selling regulation.

In their international study, Bris et al. (2007) rely on two identification strategies. First, they consider firms that are dual-listed due to American Depository Receipts (ADRs), comparing how the differences between dual-listed firms and non-dual-listed firms change based on whether the home country allows unrestricted shorting. Of course, if the motivation to dual-list is different in more developed countries (i.e., those that allow shorting) than in countries that restrict shorting, the estimated effect may be biased. In addition, recent work by Jain, Jain, McInish and McKenzie (2013) shows that home country short-selling restrictions are remarkably effective at curbing short-selling in ADRs. If true, this implies that the variation in ADR and non-ADR firm differences may be due to underlying country heterogeneity rather than short-selling restrictions. The second method is an event study using only five countries exhibiting time-series variation in short-sales regulation over their sample.

As discussed below, short-selling regulation is often endogenous to market conditions.

In general, regulators choose to regulate particular firms in particular time periods. For example, the Securities and Exchange Commission's Emergency Order (SEC Release No. 34-58592 (2008)) enacting the September/October 2008 short-selling ban opens with the following passages:

> The Commission is aware of the continued potential of sudden and excessive fluctuations of securities prices and disruption in the functioning of the securities markets that could threaten fair and orderly markets. ... Recent market conditions have made us concerned that short selling in the securities of a wider range of financial institutions may be causing sudden and excessive fluctuations of the prices of such securities in such a manner so as to threaten fair and orderly markets.
>
> Given the importance of confidence in our financial markets as a whole, we have become concerned about recent sudden declines in the prices of a wide range of securities. Such price declines can give rise to questions about the underlying financial condition of an issuer, which in turn can create a crisis of confidence, without a fundamental underlying basis. This crisis of confidence can impair the liquidity and ultimate viability of an issuer, with potentially broad market consequences.

The order goes on to list financial institutions whose stocks could no longer be sold short (with some exceptions for market makers), although substantial uncertainty remained over implementation of the order as detailed in Battalio and Schultz (2011). The language of the order highlights the fact that the SEC was particularly concerned about a select group of firms (i.e., financials) at that time.

In the next section, we examine whether results on asset prices and market quality from the many studies examining recent short-selling regulations hold in our setting where short-sale eligibility is plausibly exogenous.

5. Causal Effects of Short-Selling Eligibility

5.1. Methodology

To establish the causal effects of short selling, we examine outcomes for firms immediately within the vicinity of one of the three thresholds that makes the firm eligible for short selling. For example, consider the market capitalization threshold. When a firm surpasses the HK$1 billion threshold in market capitalization (prior to 2012), it is eligible to be shorted provided the firm also meets the turnover velocity threshold. Within the set of firms satisfying the turnover velocity threshold, firms that fall just short of the market capitalization threshold should be similar to those firms that just exceed that threshold. As such, we can estimate an unbiased treatment effect, τ, by comparing the outcomes of firms just above and just below the threshold. The only assumption we require is continuity in potential outcomes around the threshold. This assumption requires that there should be no discontinuity in outcomes if there were no difference in treatment. While this assumption is fundamentally untestable (we cannot observe the "treated" outcomes of untreated firms), the nature of the Hong Kong eligibility criteria suggests this assumption is satisfied provided that firms do not have precise control over the forcing variable.[16] In other words, within a small bandwidth around the threshold, short-sale eligibility should be quasi-random.

If the econometrician observes treatment and the underlying forcing variable (for example, market capitalization) perfectly (e.g., on the exact date on which eligibility is determined), then the treatment effect can be estimated using a standard "sharp" regression discontinuity:

$$\begin{aligned} y_{i,t} = \theta_0 + \tau \mathbb{1}(X_{i,t} - c_t > 0) + \sum_{j=1}^{N} \beta_j (X_{i,t} - c_t)^j \\ + \sum_{j=1}^{N} \theta_j \mathbb{1}(X_{i,t} - c_t > 0)(X_{i,t} - c_t)^j + \varepsilon_{i,t} \end{aligned} \tag{1}$$

[16]Manipulation by firms seems unlikely. Two of the thresholds are based on rolling averages, which would be hard to manipulate. Moreover, any manipulation would likely result in discontinuities in population density around the thresholds. We find no evidence of such discontinuities under the McCrary (2008) density test.

where $y_{i,t}$ is an outcome variable for firm i at time t, and $\mathbb{1}(X_{i,t} - c_t > 0)$ is an indicator function equal to one if the value of the forcing variable, $X_{i,t}$ (for example, market cap), exceeds the threshold value for inclusion, c_t. The outcome variable is allowed to have a flexible relationship relative to the forcing variable to either side of the threshold. To achieve this, one includes as control variables the centered distance from the threshold, $(X_{i,t} - c_t > 0)$, and the interaction between the distance and the indicator function, $\mathbb{1}(X_{i,t} - c_t > 0)$, to allow the relationship between $y_{i,t}$ and the distance to the threshold to have different slopes on either side of the cutoff. Higher-order polynomials of both the distance and the interaction can be included as well to control for non-linear effects. In the paper, we report results for N=3.[17] Under the "sharp" RD specification, τ measures the discontinuity at the threshold, which is the treatment effect of short-sale eligibility.

We do not observe the underlying forcing variables perfectly because the exact date on which the Hong Kong Stock Exchange determines eligibility is not known. To account for this, we proceed via a "fuzzy" regression discontinuity following Lee and Lemieux (2010).[18] Estimation is by two-stage least squares where the first stage estimates the probability of treatment as a function of the threshold and the second stage estimates the treatment effect of short-sale eligibility on outcomes:

$$\begin{aligned} D_{i,t} = \omega_0 + \phi \mathbb{1}(X_{i,t} - c_t > 0) + \sum_{j=1}^{N} \gamma_j (X_{i,t} - c_t)^j \\ + \sum_{j=1}^{N} \omega_j \mathbb{1}(X_{i,t} - c_t > 0)(X_{i,t} - c_t)^j + \eta_{i,t} \end{aligned} \tag{2}$$

$$y_{i,t} = \theta_0 + \tau \widehat{D}_{i,t} + \sum_{j=1}^{N} \beta_j (X_{i,t} - c_t)^j + \sum_{j=1}^{N} \theta_j \widehat{D}_{i,t} (X_{i,t} - c_t)^j + \varepsilon_{i,t} \tag{3}$$

where $D_{i,t}$ is an indicator equal to one if firm i is included on the short-sale eligibility list at

[17] Results for other values are available from the authors.
[18] The conclusions of the paper are unchanged if we instead use the "sharp" specification throughout.

time t and zero otherwise, and $\widehat{D}_{i,t}$ is the probability of treatment estimated in the first-stage equation (2). Intuitively, the fuzzy RD recognizes that observed short-selling eligibility is not perfectly predicted by the forcing variable (i.e., $\phi < 1$), but that the probability of short-sale eligibility jumps at the threshold (i.e., $\phi > 0$). Thus, one can use predicted eligibility $\mathbb{1}(X_{i,t} - c_t > 0)$ as an instrument for short-sale eligibility and estimate the treatment effect τ in the second-stage equation (3). In effect, the "fuzzy" estimate scales up the discontinuity in the outcome variable by the observed discontinuity ϕ in actual treatment.[19]

We estimate equations (2) and (3) for the three thresholds described in Section 3.1, excluding firms that may be on the list due to index inclusion or options/futures listing. For each threshold, we only estimate the results for firms where a given threshold is the only one that will affect short-sale eligibility. Specifically, for the float-adjusted market capitalization threshold, we estimate equations (2) and (3) for firms that do not satisfy both the market capitalization and turnover velocity thresholds.[20] For the turnover velocity threshold, the sample contains firms that are not included in the Designated List under the float-adjusted market capitalization rule and where the market capitalization threshold is satisfied. Similarly, for the market capitalization threshold, we estimate equations (2) and (3) for firms that are not included in the Designated List under the float-adjusted market capitalization rule and where the turnover velocity threshold is satisfied.

We consider fixed bandwidths around the centered threshold variables. These are plus/minus HK\$1 billion for both the float-adjusted market capitalization and market capitalization samples and plus/minus 50% for the turnover velocity sample.[21] The various filters and bandwidths result in different sample sizes for each threshold analysis. The float-adjusted market capitalization threshold sample contains approximately 14,000 firm-quarter observations, the turnover velocity threshold sample contains about 3,000 firm-quarter observations,

[19] If the forcing variable perfectly predicts treatment without error (i.e. $\phi = 1$), then the "sharp" estimate is equal to the "fuzzy" estimate.

[20] The float-adjusted market capitalization threshold sample runs from 2006 to 2014 because float data from Bloomberg begins in 2006 for Hong Kong firms.

[21] Our results are robust to other bandwidth choices.

and the market-capitalization threshold sample contains approximately 11,000 firm-quarter observations.

Note that certain eligibility criteria, such as index inclusion or option trading, occur disproportionately in larger stocks. Additionally, since two thresholds are based on size, there is correlation between the sample filters that require the other thresholds to not be satisfied. As a result, using a fixed bandwidth results in more observations away from the threshold on the left (ineligible) compared to the right for our two market capitalization-based thresholds. This is not a problem in our estimation because we identify the treatment effect locally around the cutoff by including the distance from the threshold and the interaction with the indicator dummy in our regression specification. An alternative approach would be to focus on a very narrow bandwidth and calculate simple mean differences. When we do this, we find similar results with relatively balanced sample sizes on both sides of the threshold. Moreover, our results are similar under the turnover threshold, which is not subject to this issue.

5.2. Short-Sale Eligibility

We first present evidence that the threshold-based rules create discontinuities in inclusion in the short sale eligibility list. Figure 1 plots the probability of list inclusion relative to the centered forcing variables.[22] The forcing variables are minimum float-adjusted market capitalization over the 60 days preceding the measurement date, turnover velocity over the year preceding the measurement date, and market capitalization as of the measurement date, shown in Panels (a)-(c), respectively. We also plot the predicted list inclusion value estimated from equation (2) along with 90% confidence bands. For all three thresholds, there is a clear discontinuity in short-sale eligibility at the threshold.

The corresponding estimates from equation (2) are tabulated in Table 3. As shown in the

[22]In the plots, we bin firms that are the same distance away from each threshold and take an average. For the turnover velocity threshold, the data are binned to the nearest half percentage difference from the threshold value. For the market value thresholds, the bins are to the the nearest quarter (half) percentage difference to the right (left) of the threshold.

plots, list inclusion is significantly associated with predicted eligibility due to each forcing variable. The increased probability ranges from 35% for the turnover velocity threshold to almost 50% for the float-adjusted market capitalization threshold.

5.3. Short-Sale Activity

A natural question is whether eligibility is in fact associated with short-selling activity. To assess this, we estimate equation (3) with shorting activity as the outcome variable. We measure shorting activity by short volume as a fraction of total volume over the 30 trading days following each quarterly effective date, denoted RELSS. The results are tabulated in Table 4 and plotted in Figure 2.[23] The results provide clear evidence that shorting activity is higher for firms to the right of each of the three thresholds.

The float-adjusted market capitalization threshold is associated with an 87 basis point increase in shorting as a fraction of total volume. The effect is even larger for the turnover velocity threshold; RELSS increases by 103 basis points around this threshold. The increase in shorting is smaller for the market capitalization threshold, at 24 basis points. At first glance, these discontinuities in actual shorting activity seem relatively modest. However, Table 1 (Panel C) shows that for all short-eligible securities, the average RELSS in Hong Kong over our sample period is 4.6%; the median RELSS is 2.5%. This means that discontinuities of approximately 1% (i.e., that of the float-adjusted market capitalization and turnover samples) are about 20% of the mean and 40% of the median RELSS for all short-eligible firms in Hong Kong, which are quite substantial relative increases.

It is useful to compare the magnitudes of our differences to those used in other studies of short-sale eligibility.[24] In their study of the 2008 U.S. financial short-sale ban, Boehmer et al. (2013) find that RELSS falls about 3% for below-median firms (on a base of 10-20%

[23] In the Internet Appendix, we tabulate these estimates for other windows corresponding to the analysis in the remainder of the paper (Table IA.1). We find similar results using other measures of short volume such as raw short volume or the logarithm of one plus short volume over the same window.

[24] We do not discuss Beber and Pagano (2013) here since they are unable to measure actual changes in short-selling activity due to the international nature of their study.

pre-ban RELSS) while the largest size quartile experiences large reductions in RELSS of 20% (relative to a pre-ban base of about 30%).[25] As such, while the absolute changes are smaller in our sample, the relative changes are not as far from the relative changes studied by Boehmer et al. (2013). Note also that short-selling due to market makers appears in the RELSS for U.S. firms (as evidenced by non-zero RELSS during the ban), but such activity is not included in short volume in Hong Kong. Therefore, the smaller average RELSS for short-eligible securities in Hong Kong versus the U.S. is partially due to measurement differences.

Our actual short-selling effects are also in line with the magnitudes found in Hong Kong by Chang et al. (2007). While they find large asset pricing effects for additions to the short-eligible list, the average (median) RELSS in their sample is a modest 0.175% (0.00%).[26] Our estimated discontinuities, which are at least as large as their average estimate, thus seem sufficient to draw inference regarding the effects of short-sale eligibility.

5.4. Returns

In this section, we examine the effects of short selling on asset prices, revisiting a number of the tests from Chang et al. (2007) in our regression discontinuity setting. We evaluate several aspects of stock returns for various windows after quarterly effective dates. Table 5 presents estimates of equation (3) for average returns, cumulative returns, return volatility, return skewness, and the prevalence of extreme downside returns (defined as the number of days with a return more than two standard deviations below the mean for a given period). Results for the three thresholds are presented in Panels A-C of Table 5.[27] The results provide no support for the conclusions that short selling causes downward pressure on prices, increased volatility, more negative skewness, or increased prevalence of extreme negative returns. In fact, the only estimate that is statistically significant is the prevalence of extreme

[25]See Figure 1 and Table 3 of Boehmer et al. (2013).

[26]See Table IV of Chang et al. (2007).

[27]We present results for raw returns. Results are unchanged if returns are measured in excess of the market return or as abnormal returns from a market model. The former are tabulated in Internet Appendix Table IA.2.

negative returns over the 30 days following effective dates for the market capitalization threshold, and the direction of the coefficient indicates that extreme returns are *less* likely for securities that are eligible to be shorted. However, this result is not robust to different windows and thresholds.

The lack of causal effects of short-selling eligibility around the thresholds on asset prices is shown graphically in Figure 3. The figure shows the return variables of interest as a function of the centered forcing variables. It is clear from the plots that there is no discontinuity at the threshold for the return variables.

The results contradict the findings of Chang et al. (2007) that there are strong return effects due to short-sale eligibility in Hong Kong. The authors further evaluate the Miller (1977) hypothesis and argue that overvaluation is more prevalent for firms subject to differences of opinions. In untabulated results, we find that there are no pricing effects around the three thresholds for subsets of stocks *ex-ante* more likely to be overvalued or subject to disagreement as measured by dispersion of analyst forecasts, higher turnover, higher market-to-book ratio, or higher past volatility of returns.

The results highlight the endogeneity of returns when using additions to the short sale eligibility list to draw inference about the effects of short-selling eligibility. In Section 4, we demonstrated that additions to the list have substantial return pre-treatment effects. We now present evidence that our regression discontinuity setting is free of this endogeneity problem. In Table 6, we examine whether there is any pre-treatment in the various return measures associated with each of our threshold samples by estimating equation (3) for various windows *preceding* the effective date. With only a single exception, the results indicate that there are no differences for the windows preceding the quarterly effective dates, confirming the validity of our setting. The one exception is a lower long-term prevalence of extreme negative returns for future short-eligible securities for the market capitalization threshold, but this effect is only marginally significant. More importantly, we see no similar effects across either of the other two thresholds examined.

How does the Hong Kong evidence compare to evidence from short-selling bans during the crisis? In terms of the level of prices, our results are generally consistent with those of Boehmer et al. (2013) and Beber and Pagano (2013) that short-selling bans did not affect the level of prices. The former focuses on pricing analysis of firms added to the US banned list after the initial announcement to avoid confounding effects of the contemporaneous TARP announcement. For these subsequently banned stocks, Boehmer et al. (2013) find no evidence of a boost in prices associated with banned short-selling. Similarly, Beber and Pagano (2013) find no evidence of changes in returns in their international panel, except in the U.S., perhaps due to the TARP announcements. Our result that exogenous short-sale eligibility is not related to subsequent returns is thus consistent with the evidence from the financial crisis bans literature.

The evidence is less consistent when turning to an examination of volatility of prices. In our setting, we find no evidence that volatility is affected by short-sale eligibility. On the other hand, both Boehmer et al. (2013) and Beber and Pagano (2013) find substantial increases in volatility for banned securities. Of course, short-sale bans are often implemented by regulators in times of extreme volatility and for the most affected firms. Our result, using a different identification strategy, suggests that volatility may be unaffected by short-sale eligibility alone.

5.5. Market Quality

We also examine how short sale eligibility may affect other aspects of the trading environment. In particular, we consider various proxies for liquidity: bid-ask spread, turnover, the Amihud measure, and the fraction of days with zero returns. The results are tabulated in Table 7. We find no significant differences in liquidity related to short-sale eligibility.

The finding that short-sale eligibility does not affect market quality stands in stark contrast to the recent literature examining the effect of short sale bans on market quality during the financial crisis. Both Boehmer et al. (2013) and Beber and Pagano (2013) find that short sale bans are associated with reduced liquidity using U.S. and international regulations, respectively. However, the two papers differ on where these effects are concentrated. Boehmer et al. (2013) find that the deterioration in market quality is present in larger stocks only; the smallest quartile of U.S. stocks affected by the ban do not experience a decline in liquidity. On the other hand, Beber and Pagano (2013) find that the detrimental liquidity effects are most prevalent for stocks with small market capitalizations and no listed options.

Our results suggest that these findings may be specific to times of financial stress. However, when we examine a similar time period for our setting (untabulated), we find results consistent with our overall sample. These differences could be explained by differences in the effect of the financial crisis in Hong Kong. The firms in the threshold samples we study are also smaller than the largest firms in Boehmer et al. (2013); they are much closer in size to the firms in the first two quartiles of the U.S. short-sale ban. It is possible that market quality of smaller firms is not sensitive to short-sale eligibility, as Boehmer et al. (2013) find. Alternatively, as discussed in Section 4, results of previous studies may stem from the endogenous adoption of short sale bans, despite the best efforts of these authors to overcome the associated empirical challenges.

5.6. Robustness Tests

5.6.1. Announcement vs. Effective Dates

It is possible that effects relative to short-sale constraints arise not due to actual short-selling but due to the threat of short selling. Indeed, evidence supporting this hypothesis is found in Grullon et al. (2015). In the Internet Appendix, we investigate whether this is the case in our sample by examining our outcome variables in periods starting at the eligibility announcement date rather than the effective date. Consistent with the effective date analysis, we find no evidence of asset pricing or market quality effects associated with short-sale eligibility even when measured from the announcement date.

5.6.2. Intraday Analysis

Because the treatment horizon is quarterly, we use daily data in our primary analysis. Many short-selling studies in the U.S. utilize intraday data to calculate market quality measures. To assess whether this difference is a concern, we use intraday trade and quote data from Thomson Reuter Tick History. From these data, we calculate additional measures of market quality. Specifically, we analyze percentage quoted spread, effective spread, realized spread, and price impact. We also examine quoted bid and ask depth and the relative bid depth, defined as the difference between quoted bid and ask depth, divided by the average depth at the bid and ask.

We report regression discontinuity results for the three threshold for these variables in the Internet Appendix. Table IA.8 tabulates regression coefficients and Figure IA.4 plots the intraday market quality measures around the thresholds. As in the daily sample, the intraday analysis reveals no systematic patterns in intraday liquidity measures across threshold samples. Our inference that short-sale eligibility has no causal effect on market quality holds when we study intraday data as well.

6. Conclusion

Despite the extensive research investigating the economic effects of short selling, it is still unclear whether and how short-selling eligibility affects stock market outcomes. We use a unique setting provided by regulation of short-selling eligibility on the Hong Kong Stock Exchange to examine the causal effects of short-selling bans on pricing and market quality.

We exploit quarterly evaluations of three threshold-based rules that determine a stock's eligibility for short selling. Using a regression discontinuity design, we find that bans on short selling do not affect asset pricing or market quality despite having a discontinuous impact on short-selling volume. In particular, we find that stock returns, volatility, bid-ask spreads, and crash risk are not statistically or economically different for banned versus unrestricted stocks. Further, we find no pricing effects for subsets of firms more likely to be overvalued or subject to disagreement.

Our paper contributes to the literature on short selling by providing new evidence on the consequences of short-selling regulation on asset pricing and market quality. Our pricing results are inconsistent with models that predict overvaluation (e.g., Miller (1977)). Theoretical predictions are ambiguous with respect to market quality and volatility. Around eligibility thresholds, we find that market quality and volatility are unaffected by short sale eligibility, in contrast to previous literature. We argue that this is likely the result of the endogenous nature of regulation. Our findings suggest that the costs of imposing short-selling ban regulations may not be as high as previously argued in the literature. Overall, the paper highlights the usefulness of novel empirical strategies to identify causal effects of short-sale regulation.

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Table 1: Summary Stats

This table presents summary statistics for stock-quarter observations on the Hong Kong Stock Exchange. The data is subset into three sub-samples. Panel A reports statistics for all stocks. Panel B and C divide the firms from Panel A into firms that are currently ineligible or eligible for shorting, respectively. For each firm-quarter, the float-adjusted market capitalization, turnover velocity, and market capitalization are month-end values, measured two months prior to a given quarterly effective date. The float-adjusted market capitalization is the minimum value measured over 60-days preceding the measurement date. Turnover velocity is the aggregate turnover over the year preceding the measurement date, divided by market capitalization at the measurement date. Market capitalization is measured at a point in time. Average daily returns, standard deviation of daily returns, and short volume as a fraction of total volume (RELSS) are measured over the 250 trading days preceding a quarterly effective date.

	Float-adj. Mkt Cap	Turnover Velocity	Market Cap	Average Return	Return SD	RELSS
Panel A: All stocks						
Mean	13,940	1.52	17,463	0.0009	0.0395	0.0149
SD	98,785	6.54	116,505	0.0031	0.0224	0.0373
P25	122	0.23	326	-0.0008	0.0247	0.0000
P50	421	0.47	1,563	0.0006	0.0342	0.0000
P75	2,031	1.00	3,674	0.0023	0.0483	0.0023
N	46,804	42,197	64,808	64,145	64,143	63,455
Panel B: Short ineligible stocks						
Mean	976	2.14	1,667	0.0009	0.0436	0.0004
SD	10,752	8.58	10,020	0.0034	0.0241	0.0041
P25	76	0.21	206	-0.0010	0.0275	0.0000
P50	183	0.48	574	0.0005	0.0387	0.0000
P75	423	1.28	2,071	0.0024	0.0541	0.0000
N	30,468	23,980	44,620	43,975	43,585	43,323
Panel C: Short eligible stocks						
Mean	38,118	0.70	52,374	0.0009	0.0305	0.0463
SD	163,848	1.08	203,920	0.0026	0.0145	0.0540
P25	1,385	0.26	3,320	-0.0005	0.0216	0.0018
P50	3,692	0.46	7,227	0.0007	0.0277	0.0249
P75	14,376	0.80	22,974	0.0021	0.0360	0.0748
N	16,336	18,217	20,188	20,170	20,165	20,132

Table 2: Analysis of All Hong Kong List Additions

This table presents market-adjusted return characteristics for the 90 (91) trading day window preceding (following) all additions to the Hong Kong short-sale eligibility list following Chang et al. (2007). Mean values of average returns, cumulative returns, return standard deviation, skewness, and extreme negative returns are presented. Extreme negative returns (Extreme Values) is defined as the number of days with a return more than two standard deviations below the mean for a given period. Standard errors are clustered by event date with t-statistics reported in parenthesis and significance represented according to: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

	Pre-Period	Post-Period	Difference
Window:	$[-90, -1]$	[0,90]	
Average Return	0.0012***	-0.0007**	-0.0019***
	(3.92)	(-2.04)	(-3.61)
Cumulative Return	0.1066***	-0.0584**	-0.1650***
	(3.93)	(-2.07)	(-3.69)
Std. Deviation	0.0353***	0.0304***	-0.0049***
	(20.18)	(24.41)	(-5.37)
Skewness	0.9137***	0.6382***	-0.2754***
	(13.88)	(11.86)	(-4.40)
Extreme Values	0.0155***	0.0181***	0.0026***
	(23.33)	(25.78)	(3.82)
Observations	1528	1528	1528

Table 3: First Stage Fuzzy Regression Discontinuity: Short Volume

This table presents the first stage (Equation (2)) of a fuzzy regression discontinuity estimate of the effect of short-sale eligibility on short volume. The second stage estimates are presented in Table 4. We report the adjusted R-squared of the first stage regression and the F-statistics on the excluded instruments. SS Eligible is an indicator equal to one if the firm is included on the short-sale eligibility list and zero otherwise. Predicted Eligibility is an indicator function equal to one if the value of the forcing variable (market capitalization, turnover velocity, and float-adjusted market capitalization) exceeds the threshold value for inclusion and zero otherwise. Distance is the centered distance from the threshold. For each firm-quarter, the float-adjusted market capitalization, turnover velocity, and market capitalization are month-end values, measured two months prior to a given quarterly effective date. The float-adjusted market capitalization is the minimum value measured over 60 days preceding the measurement date. Turnover velocity is the aggregate turnover over the year preceding the measurement date, divided by market capitalization at the measurement date. Market capitalization is measured as of the measurement date. We report the fraction of firm-quarter eligibility correctly predicted by the forcing variable. Standard errors are clustered by firm with t-statistics reported in parenthesis and significance represented according to: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

	SS Eligible	SS Eligible	SS Eligible
Predicted Eligibility	0.47***	0.35***	0.44***
	(7.63)	(4.92)	(10.31)
Distance	0.40**	3.49***	0.34***
	(2.24)	(4.66)	(3.96)
Distance^2	0.39	12.42***	0.53***
	(1.24)	(3.42)	(3.30)
Distance^3	0.13	13.26**	0.25***
	(0.79)	(2.51)	(2.74)
Predicted Eligibility*Distance	0.28	-3.95***	0.07
	(0.46)	(-3.38)	(0.18)
Predicted Eligibility*Distance^2	-1.61	-10.61*	-0.13
	(-1.14)	(-1.73)	(-0.13)
Predicted Eligibility*Distance^3	0.51	-14.23*	-0.92
	(0.49)	(-1.84)	(-1.35)
Constant	0.14***	0.33***	0.07***
	(4.63)	(7.16)	(4.96)
Observations	14349	2986	10826
Fraction correctly predicted	0.96	0.79	0.93
F-Stat	44.14	29.63	128.19
Threshold	Float-adj Mkt Cap	Turnover	Market Cap

Table 4: Second Stage Fuzzy Regression Discontinuity: Short Volume

This table presents the second stage (Equation (3)) of a fuzzy regression discontinuity estimate of the effect of short-sale eligibility on short volume. The first stage estimates are presented in Table 3. The second stage presents the estimate of the effect of short-sale eligibility on short volume as a fraction of total volume (RELSS) over the 30 trading days following a quarterly effective date. SS Eligible is the fitted value of short-sale eligibility as a function of the predicted eligibility at the threshold estimated in Equation (2). Distance is the centered distance from the threshold. For each firm-quarter, the float-adjusted market capitalization, turnover velocity, and market capitalization are month-end values, measured two months prior to a given quarterly effective date. The float-adjusted market capitalization is the minimum value measured over 60 days preceding the measurement date. Turnover velocity is the aggregate turnover over the year preceding the measurement date, divided by market capitalization at the measurement date. Market capitalization is measured as of the measurement date. Standard errors are clustered by firm with t-statistics reported in parenthesis and significance represented according to: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

	(1) RELSS	(2) RELSS	(3) RELSS
SS Eligible	0.0087**	0.0102***	0.0025**
	(2.00)	(3.10)	(2.36)
Eligibility*Distance	-0.0948	-0.0087	-0.0019
	(-0.78)	(-0.28)	(-0.09)
Eligibility*Distance^2	0.1348	0.1442	0.0117
	(0.71)	(1.06)	(0.29)
Eligibility*Distance^3	-0.1993	-0.1583	-0.0147
	(-0.79)	(-0.77)	(-0.33)
Distance	0.0291	-0.0159	0.0008
	(0.96)	(-1.12)	(0.32)
Distance^2	0.0529	-0.0262	0.0021
	(0.91)	(-0.33)	(0.28)
Distance^3	0.0265	0.0492	0.0016
	(0.88)	(0.34)	(0.27)
Constant	0.0017	-0.0020*	0.0001
	(1.46)	(-1.71)	(0.54)
Observations	14349	2986	10826
RMSE	0.0194	0.0083	0.0034
Threshold	Float-adj Mkt Cap	Turnover	Market Cap

Table 5: Second Stage Fuzzy Regression Discontinuity: Returns

This table presents the second stage (Equation (3)) of a fuzzy regression discontinuity estimate of the effect of short-sale eligibility on return characteristics for thresholds associated with float-adjusted market capitalization (Panel A), turnover velocity (Panel B), and market capitalization (Panel C). First stage estimates and second stage control variable coefficients are suppressed for space. The table presents estimates for return characteristics calculated over the windows given at the top of each sub-panel. Estimates of the short-sale eligibility discontinuity on average returns, cumulative returns, return standard deviation, skewness, and extreme negative values are presented. SS Eligible is the fitted value of short-sale eligibility as a function of the predicted eligibility at the threshold estimated in Equation (2). Extreme negative returns (Extreme Values) is defined as the number of days with a return more than two standard deviations below the mean for a given period. For each firm-quarter, the market capitalization, turnover velocity, and float-adjusted market capitalization are month-end values, measured two months prior to a given quarterly effective date. Market capitalization is measured as of this measurement date. Turnover velocity is the aggregate turnover over the year preceding the measurement date, divided by market capitalization at the measurement date. The float-adjusted market capitalization is the minimum value measured over 60 days preceding the measurement date. Standard errors are clustered by firm with t-statistics reported in parenthesis and significance represented according to: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Panel A: Float-adjusted Market Capitalization Threshold

	Avg. Ret	Cum. Ret	Ret S.D.	Skewness	Extreme Values
Window: [0,30]					
SS Eligible	0.0007	0.0159	0.0015	0.2606	-0.0034
	(0.42)	(0.35)	(0.11)	(0.70)	(-0.73)
Observations	14351	14351	14351	14351	14351
RMSE	0.0091	0.2719	0.0426	1.3145	0.0220
Window: [0,60]					
SS Eligible	0.0017	0.0930	0.0006	0.3424	0.0018
	(1.13)	(1.09)	(0.04)	(0.94)	(0.49)
Observations	14374	14374	14374	14374	14374
RMSE	0.0066	0.4086	0.0452	1.3405	0.0190
Window: [0,90]					
SS Eligible	0.0017	0.1374	0.0012	0.3301	-0.0027
	(0.94)	(0.97)	(0.08)	(0.64)	(-0.72)
Observations	14381	14381	14381	14381	14381
RMSE	0.0069	0.5987	0.0464	1.8161	0.0160
Window: [0,250]					
SS Eligible	0.0003	0.0328	0.0028	0.5210	-0.0001
	(0.19)	(0.09)	(0.20)	(0.80)	(-0.03)
Observations	14401	14401	14401	14401	14401
RMSE	0.0055	1.5443	0.0422	2.3561	0.0177

Panel B: Turnover Threshold

Window: [0,30]	Avg. Ret	Cum. Ret	Ret S.D.	Skewness	Extreme Values
SS Eligible	-0.0005	-0.0265	0.0034	-0.0454	0.0020
	(-0.26)	(-0.48)	(0.65)	(-0.16)	(0.28)
Observations	2986	2986	2986	2986	2986
RMSE	0.0059	0.1870	0.0156	1.2350	0.0305
Window: [0,60]					
SS Eligible	0.0001	-0.0102	0.0006	-0.0984	0.0013
	(0.06)	(-0.11)	(0.13)	(-0.31)	(0.27)
Observations	2989	2989	2989	2989	2989
RMSE	0.0047	0.2982	0.0163	1.0955	0.0192
Window: [0,90]					
SS Eligible	-0.0011	-0.1137	-0.0004	-0.3083	0.0038
	(-0.86)	(-0.89)	(-0.08)	(-0.88)	(0.94)
Observations	2989	2989	2989	2989	2989
RMSE	0.0040	0.4230	0.0182	1.1766	0.0165
Window: [0,250]					
SS Eligible	0.0003	-0.0368	-0.0004	0.2726	-0.0018
	(0.35)	(-0.17)	(-0.10)	(0.64)	(-0.63)
Observations	2992	2992	2992	2992	2992
RMSE	0.0026	0.8016	0.0136	1.5087	0.0116

Panel C: Market Value Threshold

	Avg. Ret	Cum. Ret	Ret S.D.	Skewness	Extreme Values
Window: [0,30]					
SS Eligible	-0.0003	0.0004	0.0019	0.0517	-0.0079*
	(-0.13)	(0.01)	(0.17)	(0.22)	(-1.78)
Observations	10828	10828	10828	10828	10828
RMSE	0.0123	0.3804	0.0559	1.1885	0.0219
Window: [0,60]					
SS Eligible	0.0009	0.0887	0.0016	-0.1074	-0.0027
	(0.35)	(0.65)	(0.15)	(-0.27)	(-0.71)
Observations	10855	10855	10855	10855	10855
RMSE	0.0124	0.6583	0.0565	2.0361	0.0170
Window: [0,90]					
SS Eligible	0.0012	0.1395	0.0011	-0.0589	-0.0015
	(0.50)	(0.68)	(0.10)	(-0.11)	(-0.33)
Observations	10869	10869	10869	10869	10869
RMSE	0.0110	0.9231	0.0587	2.5532	0.0219
Window: [0,250]					
SS Eligible	0.0001	0.0021	0.0010	-0.3637	-0.0001
	(0.09)	(0.01)	(0.09)	(-0.86)	(-0.04)
Observations	10895	10895	10895	10895	10895
RMSE	0.0051	1.5441	0.0510	2.4627	0.0130

Table 6: Pre-Treatment Effects: Returns

This table presents the second stage (Equation (3)) of a fuzzy regression discontinuity estimate of the effect of short sale eligibility on return characteristics in the period prior to the effective date of the short-sale list for thresholds associated with float-adjusted market capitalization (Panel A), turnover velocity (Panel B), and market capitalization (Panel C). First stage estimates and second stage control variable coefficients are suppressed for space. The table presents estimates for return characteristics calculated over the windows given at the top of each sub-panel. Estimates of the short-sale eligibility discontinuity on average returns, cumulative returns, return standard deviation, skewness, and extreme negative values are presented. SS Eligible is the fitted value of short-sale eligibility as a function of the predicted eligibility at the threshold estimated in Equation (2). For each firm-quarter, the market capitalization, turnover velocity, and float-adjusted market capitalization are month-end values, measured two months prior to a given quarterly effective date. Market capitalization is measured as of this measurement date. Turnover velocity is the aggregate turnover over the year preceding the measurement date, divided by market capitalization at the measurement date. The float-adjusted market capitalization is the minimum value measured over 60 days preceding the measurement date. Standard errors are clustered by firm with t-statistics reported in parenthesis and significance represented according to: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Panel A: Float-adjusted Market Capitalization Threshold

	Avg. Ret	Cum. Ret	Ret S.D.	Skewness	Extreme Values
Window: $[-30, -1]$					
SS Eligible	-0.0003	-0.0162	0.0038	-0.1696	0.0027
	(-0.12)	(-0.27)	(0.24)	(-0.50)	(0.52)
Observations	14355	14355	14355	14336	14355
RMSE	0.0111	0.3095	0.0539	1.4700	0.0241
Window: $[-60, -1]$					
SS Eligible	0.0007	0.0232	0.0046	-0.2105	-0.0008
	(0.44)	(0.28)	(0.29)	(-0.63)	(-0.20)
Observations	14378	14378	14378	14374	14378
RMSE	0.0073	0.4246	0.0509	1.5438	0.0175
Window: $[-90, -1]$					
SS Eligible	0.0009	0.0925	0.0042	-0.2043	0.0002
	(0.76)	(0.88)	(0.24)	(-0.68)	(0.06)
Observations	14383	14383	14383	14382	14383
RMSE	0.0056	0.5361	0.0544	1.5360	0.0140
Window: $[-250, -1]$					
SS Eligible	-0.0001	-0.0450	0.0031	0.0409	-0.0003
	(-0.11)	(-0.18)	(0.19)	(0.12)	(-0.11)
Observations	14403	14403	14403	14403	14403
RMSE	0.0033	1.0868	0.0505	1.9404	0.0119

Panel B: Turnover Threshold

Window: $[-30, -1]$	Avg. Ret	Cum. Ret	Ret S.D.	Skewness	Extreme Values
SS Eligible	-0.0017	-0.0483	0.0007	0.0721	-0.0126
	(-0.76)	(-0.69)	(0.13)	(0.25)	(-1.50)
Observations	2988	2988	2988	2985	2988
RMSE	0.0081	0.2522	0.0177	1.2339	0.0427
Window: $[-60, -1]$					
SS Eligible	-0.0003	0.0284	0.0020	-0.0874	-0.0024
	(-0.19)	(0.23)	(0.41)	(-0.29)	(-0.49)
Observations	2991	2991	2991	2991	2991
RMSE	0.0060	0.3839	0.0163	1.1355	0.0192
Window: $[-90, -1]$					
SS Eligible	-0.0014	-0.0891	0.0028	-0.2739	0.0002
	(-0.90)	(-0.50)	(0.57)	(-0.74)	(0.04)
Observations	2991	2991	2991	2991	2991
RMSE	0.0052	0.5887	0.0159	1.6345	0.0173
Window: $[-250, -1]$					
SS Eligible	-0.0008	-0.4030	-0.0011	0.0584	-0.0024
	(-0.82)	(-0.99)	(-0.24)	(0.14)	(-0.88)
Observations	2994	2994	2994	2994	2994
RMSE	0.0036	1.3874	0.0166	1.4317	0.0101

Panel C: Market Value Threshold

	Avg. Ret	Cum. Ret	Ret S.D.	Skewness	Extreme Values
Window: $[-30, -1]$					
SS Eligible	-0.0014	-0.0134	-0.0048	-0.1511	0.0019
	(-0.55)	(-0.19)	(-0.74)	(-0.72)	(0.35)
Observations	10830	10830	10830	10812	10830
RMSE	0.0129	0.3712	0.0366	1.1459	0.0254
Window: $[-60, -1]$					
SS Eligible	-0.0002	0.0106	-0.0049	-0.0657	-0.0008
	(-0.13)	(0.12)	(-0.61)	(-0.24)	(-0.21)
Observations	10857	10857	10857	10849	10857
RMSE	0.0076	0.4434	0.0431	1.4691	0.0189
Window: $[-90, -1]$					
SS Eligible	-0.0003	-0.0482	-0.0028	-0.3558	0.0006
	(-0.33)	(-0.46)	(-0.30)	(-1.38)	(0.20)
Observations	10869	10869	10869	10866	10869
RMSE	0.0054	0.5466	0.0481	1.5801	0.0148
Window: $[-250, -1]$					
SS Eligible	-0.0004	-0.0716	-0.0016	-0.1689	-0.0006
	(-0.35)	(-0.25)	(-0.19)	(-0.36)	(-0.21)
Observations	10895	10895	10895	10894	10895
RMSE	0.0061	1.2308	0.0438	2.5275	0.0148

Table 7: Second Stage Fuzzy Regression Discontinuity: Market Quality

This table presents the second stage (Equation (3)) of a fuzzy regression discontinuity estimate of the effect of short sale eligibility on measures of market quality for thresholds associated with float-adjusted market capitalization (Panel A), turnover velocity (Panel B), and market capitalization (Panel C). Estimates of the short-sale eligibility discontinuity on bid-ask spread, turnover, the Amihud measure, and the fraction of days with zero returns (Zeros) are presented. The bid-ask spread is expressed as a fraction of the price. Turnover is expressed as a percent. Amihud is the price impact measure developed by Amihud (2002), expressed as the absolute return per $1 million dollars volume. First stage estimates and second stage control variable coefficients are suppressed for space. SS Eligible is the fitted value of short-sale eligibility as a function of the predicted eligibility at the threshold estimated in Equation (2). For each firm-quarter, the market capitalization, turnover velocity, and float-adjusted market capitalization are month-end values, measured two months prior to a given quarterly effective date. Market capitalization is measured as of this measurement date. Turnover velocity is the aggregate turnover over the year preceding the measurement date, divided by market capitalization at the measurement date. The float-adjusted market capitalization is the minimum value measured over 60 days preceding the measurement date. Standard errors are clustered by firm with t-statistics reported in parenthesis and significance represented according to: $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Panel A: Float-adjusted Market Capitalization Threshold

	Bid-Ask Spread	Turnover	Amihud	Zeros
Window: [0,30]				
SS Eligible	0.0043	0.0264	0.4735	0.0052
	(0.32)	(0.16)	(0.26)	(0.11)
Observations	14335	14355	14349	14351
RMSE	0.0804	0.6852	6.4849	0.1945
Window: [0,60]				
SS Eligible	0.0014	0.0500	0.4477	0.0147
	(0.17)	(0.28)	(0.24)	(0.38)
Observations	14357	14378	14371	14374
RMSE	0.0667	0.6781	6.8054	0.1576
Window: [0,90]				
SS Eligible	-0.0008	0.0526	0.4270	0.0137
	(-0.16)	(0.27)	(0.23)	(0.38)
Observations	14365	14383	14380	14381
RMSE	0.0623	0.6878	6.8978	0.1487
Window: [0,250]				
SS Eligible	-0.0018	0.0395	0.4417	0.0134
	(-0.20)	(0.18)	(0.24)	(0.43)
Observations	14387	14403	14401	14401
RMSE	0.0615	0.7345	6.8607	0.1352

Panel B: Turnover Threshold

	Bid-Ask Spread	Turnover	Amihud	Zeros
Window: [0,30]				
SS Eligible	0.0004	-0.0236	0.0100	-0.0361
	(0.07)	(-0.30)	(0.04)	(-0.93)
Observations	2984	2988	2986	2986
RMSE	0.0302	0.2510	1.6138	0.1468
Window: [0,60]				
SS Eligible	0.0002	0.0038	-0.7109	-0.0366
	(0.03)	(0.05)	(-1.26)	(-1.07)
Observations	2987	2991	2989	2989
RMSE	0.0286	0.2725	1.7586	0.1399
Window: [0,90]				
SS Eligible	0.0011	-0.0061	-0.7757	-0.0310
	(0.18)	(-0.08)	(-1.04)	(-0.95)
Observations	2987	2991	2989	2989
RMSE	0.0292	0.2525	2.1723	0.1384
Window: [0,250]				
SS Eligible	-0.0006	-0.0241	-0.4432	0.0066
	(-0.08)	(-0.47)	(-0.37)	(0.21)
Observations	2990	2994	2992	2992
RMSE	0.0310	0.1894	4.2999	0.1288

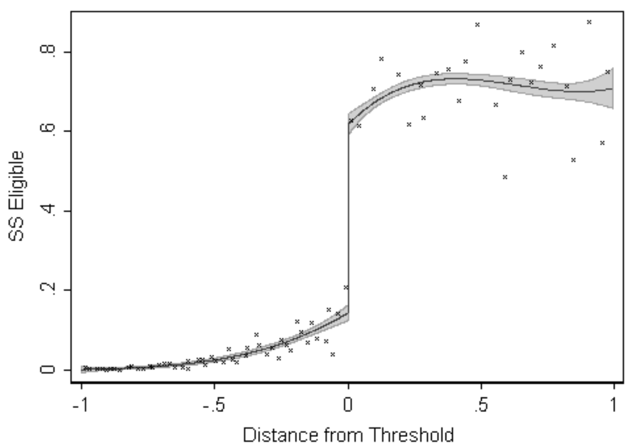
Panel C: Market Value Threshold

	Bid-Ask Spread	Turnover	Amihud	Zeros
Window: [0,30]				
SS Eligible	-0.0211	-0.0189	-0.4704	-0.0544*
	(-1.39)	(-0.05)	(-0.66)	(-1.93)
Observations	10826	10830	10826	10828
RMSE	0.0990	1.8808	4.4851	0.1625
Window: [0,60]				
SS Eligible	-0.0152	0.0008	-0.6702	-0.0459
	(-1.42)	(0.00)	(-1.19)	(-1.40)
Observations	10853	10857	10853	10855
RMSE	0.0825	1.8408	4.0194	0.1816
Window: [0,90]				
SS Eligible	-0.0190*	0.0302	-0.4826	-0.0619*
	(-1.91)	(0.08)	(-0.85)	(-1.65)
Observations	10867	10869	10868	10869
RMSE	0.0802	1.8571	3.9067	0.1906
Window: [0,250]				
SS Eligible	-0.0238	-0.0382	-0.2802	-0.0714*
	(-1.56)	(-0.11)	(-0.65)	(-1.80)
Observations	10894	10895	10894	10895
RMSE	0.0935	1.7957	3.6192	0.1881

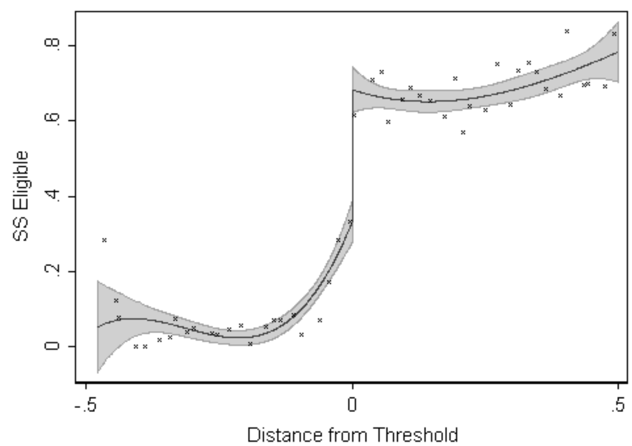
Figure 1: Short-sale eligibility thresholds

This figure plots the short-sale eligibility of firms relative to the three thresholds: float-adjusted market capitalization, turnover velocity, and market capitalization. For the float-adjusted market capitalization threshold, the sample runs from 2006 to 2014. For the market capitalization and turnover velocity thresholds, the sample runs from 2001 to 2014.

(a) Float-adjusted market capitalization



(b) Turnover velocity



(c) Market capitalization

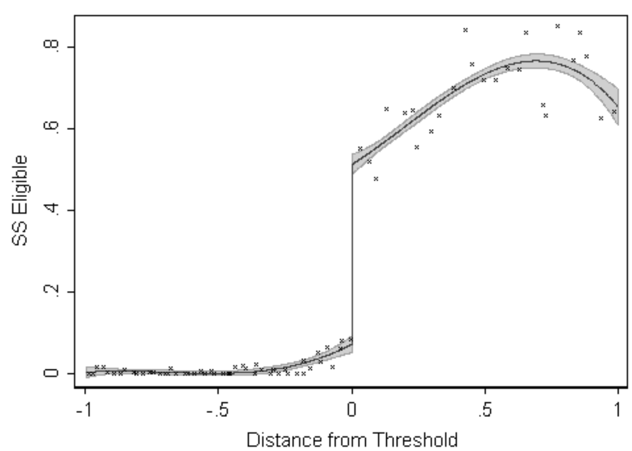
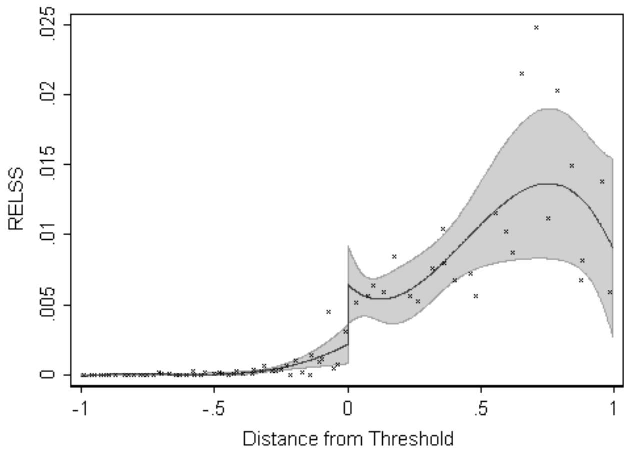


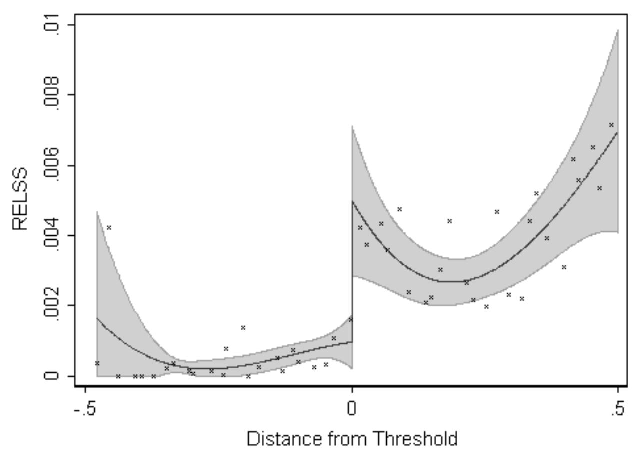
Figure 2: Short-sale volume

This figure plots short-sale volume as a fraction of total volume (RELSS) relative to the three thresholds: float-adjusted market capitalization, turnover velocity, and market capitalization. For the float-adjusted market capitalization threshold, the sample runs from 2006 to 2014. For the market capitalization and turnover velocity thresholds, the sample runs from 2001 to 2014.

(a) Float-adjusted market capitalization



(b) Turnover velocity



(c) Market capitalization

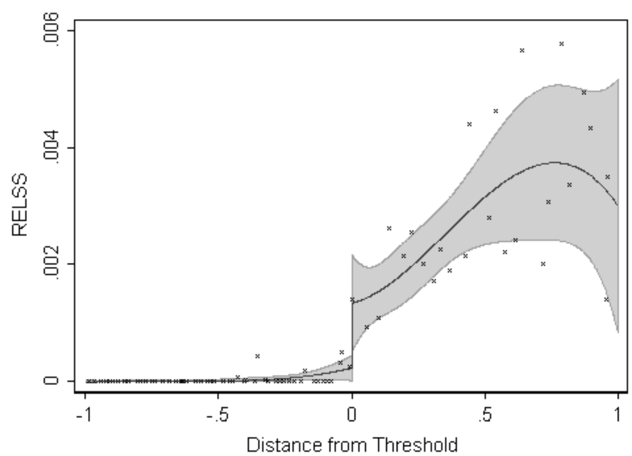
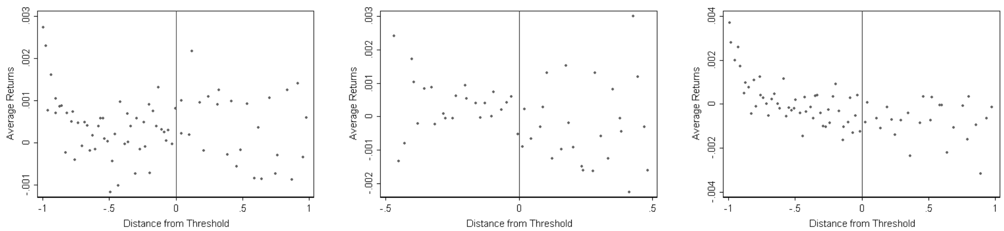


Figure 3: Returns

This figure plots return statistics relative to the three thresholds: float-adjusted market capitalization, turnover velocity, and market capitalization. For the float-adjusted market capitalization threshold, the sample runs from 2006 to 2014. For the market capitalization and turnover velocity thresholds, the sample runs from 2001 to 2014.

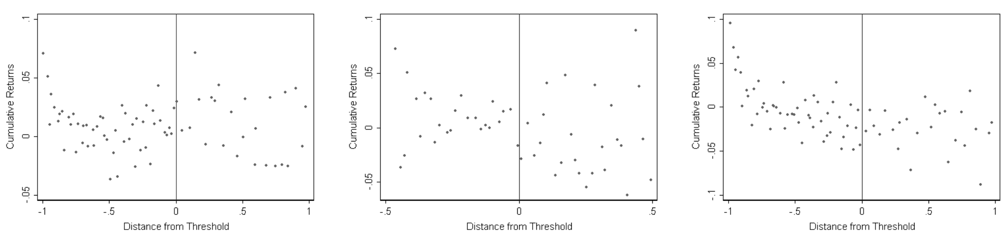
Panel A: Average Return $[0, 30]$

Float-adj. Mkt Cap | Turnover | Market Cap



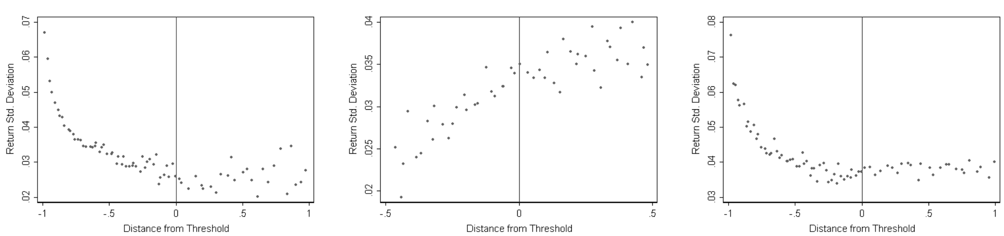
Panel B: Cumulative Return $[0, 30]$

Float-adj. Mkt Cap | Turnover | Market Cap



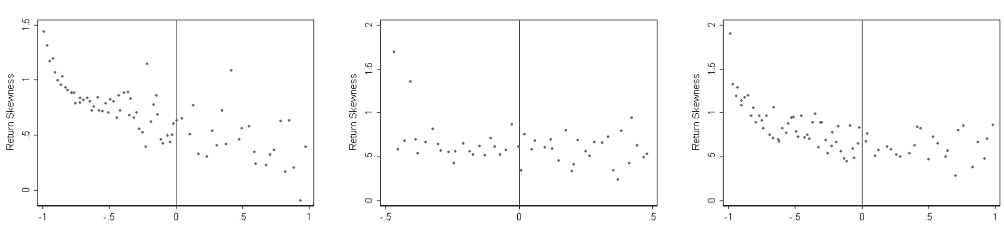
Panel C: Return Volatility $[0, 90]$

Float-adj. Mkt Cap | Turnover | Market Cap



Panel D: Return Skewness $[0, 90]$

Float-adj. Mkt Cap | Turnover | Market Cap



Panel E: Extreme Returns $[0, 90]$

Float-adj. Mkt Cap | Turnover | Market Cap

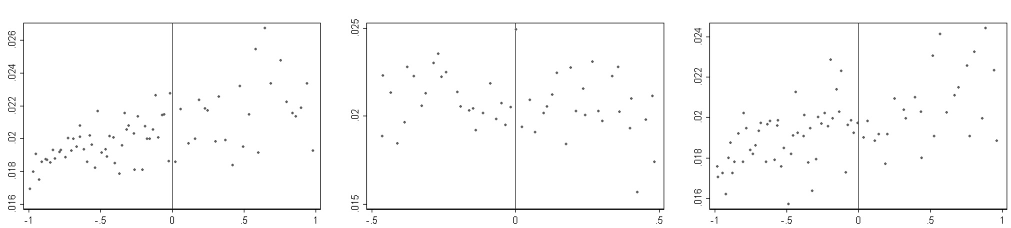
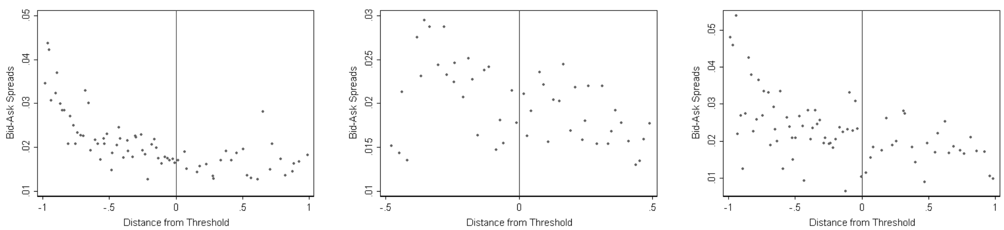


Figure 4: Market Quality

This figure plots market quality statistics relative to the three thresholds: float-adjusted market capitalization, turnover velocity, and market capitalization. For the float-adjusted market capitalization threshold, the sample runs from 2006 to 2014. For the market capitalization and turnover velocity thresholds, the sample runs from 2001 to 2014.

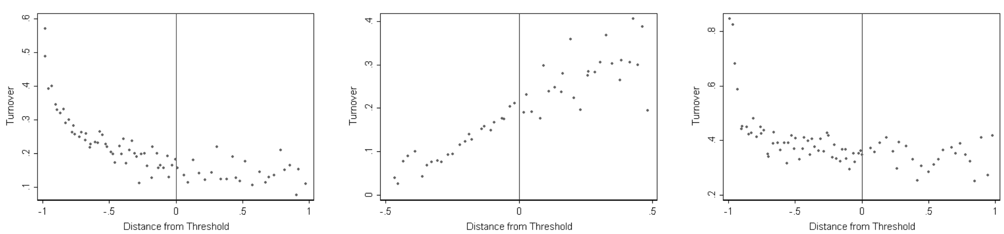
Panel A: Average Bid-Ask Spreads $[0, 30]$

Float-adj. Mkt Cap | Turnover | Market Cap



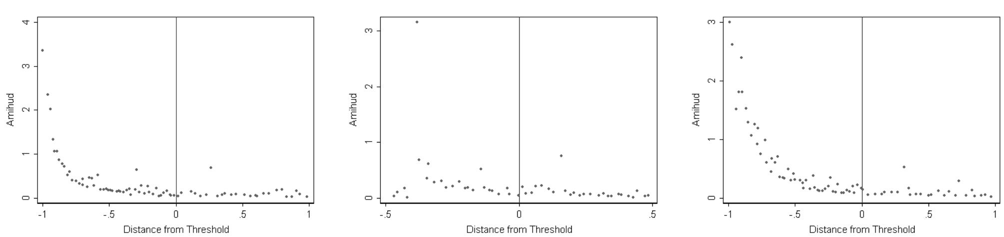
Panel B: Average Turnover $[0, 30]$

Float-adj. Mkt Cap | Turnover | Market Cap



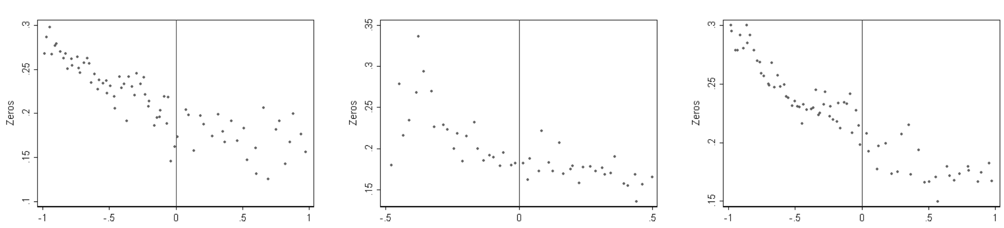
Panel C: Amihud $[0, 30]$

Float-adj. Mkt Cap | Turnover | Market Cap



Panel D: Zero Return Days $[0, 30]$

Float-adj. Mkt Cap | Turnover | Market Cap



Data Appendix: Variable Definitions and Source Data

Variable Name	Definition	Data Source
SS	Indicator variable equal to 1 if firm is eligible to short sold and 0 if not.	Hand-collected
RELSS	Short-selling volume over a given interval divided by total volume over the period	Bloomberg
Float-adjusted market cap	The minimum float-adjusted market capitalization over the 60 days preceding a given measurement date. Float-adjusted market capitalization is caclulated as price times publicly-available shares. This variable is available from 2006 onwards.	Bloomberg
Turnover velocity	Total dollar volume over the year preceding the measurement date divided by market capitalization at the measurement date	Bloomberg
Market Capitalization	The firm's equity value measured as price times total shares outstanding	Bloomberg
Average Return	The arithmetic average daily return over an indicated interval	Bloomberg
Return SD	The standard deviation of daily returns over an indicated interval	Bloomberg
Cumulative Return	The cumulative return over an indicated interval	Bloomberg
Skewness	The skewness of daily returns over an indicated interval	Bloomberg
Extreme Values	The fraction of days with returns two standard deviations below the average return for a firm. The standard deviation and average return are measured in the given interval.	Bloomberg
Predicted Eligibility	Equals 1 if a given forcing variable (e.g., turnover velocity) indicates the firm should be short eligible. Equals 0 otherwise.	Bloomberg
Distance	The difference between a forcing variable (e.g., turnover velocity) and the eligibility threshold.	Bloomberg
Bid-ask spread	The difference between the ask and bid price, divided by the midpoint price	Bloomberg
Turnover	Shares traded over a period divided by shares outstanding. Expressed as a percent.	Bloomberg
Amihud	The average daily value of the price impact measure of Amihud (2002), calculated as absolute value of return divided by millions of dollars trading volume.	Bloomberg
Zeros	Fraction of days in an interval with a zero-return.	Bloomberg