The Invisible Hand of Short Selling: Does Short Selling Discipline Earnings Management?

Massimo Massa
INSEAD

Bohui Zhang
School of Banking and Finance, UNSW Business School, UNSW Australia

Hong Zhang
PBC School of Finance, Tsinghua University and INSEAD

We hypothesize that short selling has a disciplining role vis-à-vis firm managers that forces them to reduce earnings management. Using firm-level short-selling data for thirty-three countries collected over a sample period from 2002 to 2009, we document a significantly negative relationship between the threat of short selling and earnings management. Tests based on instrumental variable and exogenous regulatory experiments offer evidence of a causal link between short selling and earnings management. Our findings suggest that short selling functions as an external governance mechanism to discipline managers. (*JEL* G30, M41)

Short selling has traditionally been identified as a factor that contributes to market informational efficiency.[1] However, short selling has also been regarded

We thank two anonymous referees, Andrew Karolyi (the Editor), Reena Aggarwal, George Aragon, Douglas Breeden, Michael Brennan, Murillo Campello, Henry Cao, Gu Chao, Hui Chen, Bernard Dumas, Philip Dybvig, Alex Edmans, Vivian Fang, Nickolay Gantchev, Mariassunta Giannetti, Zhiguo He, Pierre Hillion, Soren Hivkiar, Albert (Pete) Kyle, Ting Li, Bryan Lim, Jun Liu, Mark Maffett, Ronald Masulis, David Ng, Lilian Ng, Marco Pagano, Stavros Panageas, Neil Pearson, Lasse Pedersen, Joel Peress, Yaxuan Qi, David Reeb, Amit Seru, Philip Strahan, Kumar Venkataraman, Jiang Wang, Yajun Wang, Wei Xiong, Feifei Zhu, and participants of numerous seminars for valuable comments. We also thank the 2013 China International Conference in Finance's program committee for awarding us the TCW Best Paper Award, the 2013 Asian Finance Association's program committee for awarding us the JUFE Best Paper Award, the 2013 Northern Finance Association Conference's program committee for awarding us the CFA Society Toronto Award. We are grateful to Russell Investments for generously providing data on the Russell index components. Bohui Zhang acknowledges research grants from the ARC discovery grant (DP 120104755) and ARC linkage grant (LP130101050) from Australian Research Council, and CIFR research grants (E026 and E028) from the Centre for International Finance and Regulation. Supplementary data can be found on *The Review of Financial Studies* web site. Send correspondence to Bohui Zhang, the School of Banking and Finance, UNSW Business School, UNSW Australia, Sydney, NSW 2052, Australia; telephone: 61-2-93855834. E-mail: bohui.zhang@unsw.edu.au or bohui.zhang@gmail.com.

[1] Please see Miller (1977), Diamond and Verrecchia (1987), Duffie, Garleanu, and Pedersen (2002), Bris, Goetzmann, and Zhu (2007), Boehmer, Jones, and Zhang (2008), Boehmer and Wu (2013), Saffi and Sigurdsson (2011), and Akbas et al. (2013).

doi:10.1093/rfs/hhu147 Advance Access publication December 23, 2014

as "dangerous" to the stability of financial markets and has even been banned in many countries during financial crises.[2] Notably, these two seemingly conflicting views are based on the same traditional wisdom that short selling affects *only* the way in which information is incorporated into market prices by making the market reaction either more effective or overly sensitive to existing information but does not affect the behavior of firm managers, who may shape, if not generate, information in the first place.

However, short selling may also directly influence the behavior of firm managers. To understand the intuition, consider a manager who can manipulate a firm's earnings to reap some private benefits but who faces reputational or pecuniary losses if the public uncovers this manipulation. The manager will be confronted with a trade-off between the potential benefits and losses. The presence of short sellers affects this trade-off. As short sellers increase price informativeness and attack the misconduct of firms (e.g., Hirshleifer, Teoh, and Yu 2011; Karpoff and Lou 2010), their presence, by increasing the probability and speed with which the market uncovers earnings management, reduces managers' incentives to manipulate earnings. We call this view the disciplining hypothesis.

On the other hand, the downward price pressure of short selling may increase the negative impact of failing to meet market expectations. Therefore, any additional downward price pressure arising from short selling may incentivize firms to manipulate earnings. In other words, the threat of potential bear raids may drive managers to manipulate earnings to avoid the attention of short sellers and thus the confounding impact associated with the downward price pressure of their trades. We call this view the price pressure hypothesis. These considerations, together with the aforementioned traditional wisdom implying that managers may simply ignore the existence of short sellers (which can thus be labeled the ignorance hypothesis), suggest that short selling may have conflicting effects in the real economy. Distinguishing among these competing hypotheses is critical to elucidate the real impact of short selling, which is the aim of this paper.

To detect the potential impact of short selling, we focus on the ex ante "short-selling potential" (SSP)—that is, the maximum potential impact that short sellers may have on firm behavior or stock prices[3]—as opposed to the ex post actions taken by short sellers in response to observed firm manipulation. The main proxy for SSP is the total supply of shares that are available to be lent for short sales (hereafter, *Lendable*). This variable is directly related to the theory on the ex ante impact of short selling. Diamond and Verrecchia (1987),

[2] The general public concern is the potential that short selling is inherently speculative and exerts downward price pressure that may destabilize the market. The Securities and Exchange Commission (SEC), for instance, believes that the adoption of a short sale-related circuit breaker is beneficial as it avoids the price impact of manipulative or abusive short selling (http://www.sec.gov/rules/final/2010/34-61595.pdf).

[3] Even with limits to arbitrage, small short sellers can also affect stock prices of hard-to-short companies by using media campaigns (Ljungqvist and Qian 2014).

for instance, demonstrate that short-sale constraints reduce informative trades and the speed of adjustment to private information. A limited supply of lendable shares imposes precisely this type of constraint (Saffi and Sigurdsson 2011). Thus, a high fraction of shares lendable to short sellers implies a high degree of SSP that may either discipline managers or exert price pressure. Moreover, more active shareholders are also less likely to lend shares to short sellers on a large scale (e.g., Prado, Saffi, and Sturgess 2013).[4] This unique property will also help us to identify the passive supplies of lendable shares as an instrument to control for the spurious impact of internal monitoring.

We focus on earnings management because it represents one of the "most tangible signs" of distorted information in global markets (e.g., Leuz, Nanda, and Wysocki 2003). Moreover, earnings management has important normative and policy implications in numerous countries that have fallen under regulatory scrutiny, following Regulation Fair Disclosure and the Sarbanes-Oxley Act in the United States (Dechow, Ge, and Schrand 2010). In line with the literature (e.g., Jones 1991; Dechow, Sloan, and Sweeney 1995; Dechow, Ge, and Schrand 2010; Hirshleifer, Teoh, and Yu 2011), we use discretionary accruals as the main proxy for earnings management. In this context, the disciplining hypothesis posits that SSP reduces discretionary accruals, while the price pressure hypothesis posits the opposite. No effect is expected under the ignorance hypothesis.

We test these hypotheses by using a worldwide sample of short selling covering 17,555 firms from thirty-three countries over the 2002–2009 period. We begin by documenting a strong negative correlation between the SSP of a stock and the extent of the firm's earnings management. This effect is both statistically significant and economically relevant. A one-standard-deviation increase in SSP is associated with 5.12% standard deviation less earnings management. This relationship is robust to the use of fixed effects and the adoption of a dynamic-panel generalized method of moments (GMM) estimator (Arellano and Bond 1991). These findings offer the first evidence supporting the disciplining hypothesis.

To address issues of potential endogeneity and spurious correlation, we adopt a twofold approach. First, we use an instrumental variable approach based on the ownership of exchange-traded funds (ETFs) that fully replicate benchmarks. On the one hand, fully replicating ETFs are passive investors. These funds typically do not monitor firms or blow the whistle on corporate fraud (Dyck, Morse, and Zingales 2010), as they thrive on a low-fee strategy, which makes

[4] Activist investors have less incentive to lend out their shares because the ownership and voting rights of lendable shares will be transferred because of the short sale—and the lack of voting rights is known to discourage the participation of active institutional investors (e.g., Li, Ortiz-Molina, and Zhao 2008). Indeed, lending may occur precisely to transfer voting rights rather than exercising voting rights (e.g., Christoffersen et al. 2007), and majority lenders do not seem to actively exercise the voting power of their lendable shares, evident by the fact that only less than 2% of shares on loan are called back on the proxy voting record date (Aggarwal, Saffi, and Sturgess forthcoming).

active monitoring unlikely, if not impossible. On the other hand, the same low-fee strategy also induces ETFs to supply lendable shares to the short-selling market, which enables them to further reduce fees. In this regard, the astonishing 40% annual growth rate of the ETF industry over the past decade, driven by investor demand for index investment, provides large exogenous variation in the amount of shares that are available for short selling. In line with our expectations, ETF ownership significantly explains the SSP variations in our sample. All of these features—the passive nature of ownership, the supply of lendable shares motivated by fees, and the time series variations in ETF ownership attributable to investor flows focusing on benchmarks—make ETF ownership an ideal instrument for the share of SSP unrelated to earnings management. To further control for unobservable firm characteristics, we use both firm-level and industry-wide ETF ownership in our tests.

We find that instrumented SSP also significantly reduces earnings management. Moreover, when we directly link ETF ownership to earnings management, we find that ETF ownership does not reduce earnings management when SSP is included in the full-sample regressions or when SSP is low or prohibited in subsample regressions. These results suggest that ETF ownership affects earnings management through its effect on short selling.

Second, we consider an event-based approach that explores two regulatory experiments: the SEC Regulation SHO in the United States and the gradual introduction of (regulated) short selling on the Hong Kong Stock Exchange. The U.S. experiment began in 2005 and lasted until 2007. The SEC established a pilot program that exempted one-third of the stocks on the Russell 3000 Index from price restrictions that were related to short selling. The choice of the stock was purely random across average daily trading volume levels within the NYSE, NASDAQ, and AMEX stock exchanges (e.g., Diether, Lee, and Werner 2009; Grullon, Michenaud, and Weston 2012). We find compelling evidence that lifting short-selling restrictions—that is, Regulation SHO—reduced earnings management by between 6.57% and 7.88%, on average, depending on the specifications.

In Hong Kong, short selling was prohibited until 1994, when the Hong Kong Stock Exchange introduced a pilot scheme allowing short selling for a list of seventeen stocks. Since then, the list of firms that are eligible for short selling has changed, creating both time-series and cross-sectional variations with respect to short-selling restrictions for firms listed in Hong Kong. Similarly to the case of Regulation SHO in the United States, we find that stocks for which short selling has been allowed experience dramatic reductions in earnings management.

For both experiments, we design "placebo" tests to further confirm that changes in earnings management are only related to the regulatory changes in short-selling restrictions. Overall, these tests support a causal interpretation of the relationship between SSP and reduced earnings management, which is consistent with the disciplining hypothesis as opposed to the alternative hypotheses.

In addition to the above main tests, we also implement a series of additional tests to further enrich our economic intuition. First, we show that, worldwide, regulations that restrict short selling (such as country-wide short-selling bans) are typically associated with greater earnings management. These results suggest that the importance of short-selling regulations in affecting earnings management incentives is not limited to a few selected markets. Second, in line with the observation that short-selling activity grew tremendously in our sample period with the emergence of hedge funds (e.g., Saffi and Sigurdsson 2011), we document that the disciplining impact of short selling on earnings management also increases over time. Third, we show that the disciplining effect is robust to the use of alternative SSP proxies and that it applies to a wide spectrum of earnings management measures, including not only additional discretionary accruals but also a list of target-beating, earnings persistence, and earnings misstatement measures. Finally, based on the framework of Morck, Yeung, and Yu (2000) and Jin and Myers (2006) in general and Bris, Goetzmann, and Zhu (2007) in particular, we not only confirm a positive relationship between short selling and stock price informativeness (Saffi and Sigurdsson 2011), but also find that this positive relationship is more pronounced when the potential impact of short selling on earnings management is high, suggesting that short selling may increase price efficiency by reducing the incentives for firms to manage their earnings.

Overall, these results offer evidence of a beneficial, rather than detrimental effect of the short-selling market on the corporate market. They are closely related to Hirshleifer, Teoh, and Yu (2011) and Karpoff and Lou (2010). These authors show that short sellers attack firms that manipulate earnings or exhibit misconduct; we show that the very possibility of such potential attacks reduces earnings management. This finding has important normative implications because it shows that short selling—which is generally considered a source of the problem of deceptive market information—does in fact contribute to solve such a problem. In a contemporaneous paper, Fang, Huang, and Karpoff (2014) confirm our conclusions focusing on the SHO experiment. Our paper differs by providing extensive evidence related to lendable shares and their passive suppliers. This approach not only allows us to measure the disciplining impact of short selling using a concrete proxy, but also enables us to identify a pivotal economic channel—that is, ETFs—that affects the prosperity and efficiency of the short selling market. The link to ETF investment enriches our knowledge on how passive and active investors intertwine in affecting the real economy. Moreover, our focus is broader, considering the impact of short selling on the global market, not merely the U.S. market. Jointly, therefore, our results provide both concrete channels and unique international experience that policy makers, especially those in emerging markets, can rely on to improve firm efficiency through both the adoption of better short-selling-related regulations and the development of passive (ETF-alike) and long-term investors.

We contribute to different strands of the literature. First, we provide the first analysis—to the best of our knowledge—of the real impact of the short-selling market on corporate behavior in general and earnings management in particular. While the standard short-selling literature links short sellers' activities to stock returns (Senchack and Starks 1993; Asquith and Meulbroek 1995; Aitken et al. 1998; Cohen, Diether, and Malloy 2007; Boehmer, Jones, and Zhang 2008; Boehmer and Wu 2013; Saffi and Sigurdsson 2011), we contribute by directly linking short sellers' activity—or more specifically, the threat of their activity—to managerial behavior.

Second, we contribute to the corporate governance literature, which has studied the trade-off between "voice and exit" (Maug 1998; Kahn and Winton 1998; Faure-Grimaud and Gromb 2004). This stream of literature has focused on "voice" as the primary disciplining device, though recent studies also show that "exit" is a governance mechanism in itself (e.g., Admati and Pfleiderer 2009; Edmans 2009; Edmans and Manso 2011; and Edmans, Fang, and Zur 2013). Unlike the previously discussed governance mechanisms, the disciplining force of the short-selling channel identified in our paper arises from the outside (i.e., from the external market) as opposed to the inside (i.e., from existing shareholders). Thus, the "invisible hand" of the market affects and disciplines managers.

Third, our results contribute to the literature on the determinants of earnings management, which has focused on firm operating and financial characteristics (see DeFond and Park 1997; Watts and Zimmerman 1986; Nissim and Penman 2001), auditing quality and financial reporting practices (DeAngelo 1981; Barth, Landsman, and Lang 2008), market pressure (Das and Zhang 2003; Morsfield and Tan 2006), as well as investor protection and regulations (Leuz, Nanda, and Wysocki 2003; Dechow, Ge, and Schrand 2010). Our evidence on the role of SSP provides another external channel to mitigate managers' incentives to manage accounting earnings.

1. Data, Variable Construction, and Preliminary Evidence

1.1 Data sample and sources

The sample of short selling covers the period between 2002 and 2009. We begin with all publicly listed companies worldwide for which we have accounting and stock market information from Datastream/WorldScope. This sample is then matched with short-selling information data from DataExplorers and data on institutional investors' stock holdings from FactSet/LionShares.

We obtain equity-lending data from DataExplorers, a research company that collects equity- and bond-lending data directly from the securities lending desks at the world's leading financial institutions. Information detailed at the stock level is available from May 2002 to December 2009. In particular, the dataset provides unique information on the value of shares that are on loan to short sellers and on the value of shares that are available to be lent to short sellers;

both sets of information are important for the analysis in this paper. More detailed descriptions of the data can be found in Saffi and Sigurdsson (2011) and Jain et al. (2013). DataExplorers provides monthly information in 2002 and 2003 (weekly information from July 2004 on and daily information after 2006). Because of DataExplorers' low coverage during the first two years, we also show the robustness of our findings by focusing on a shorter period from 2004 to 2009 or from 2006 to 2009 in Section 4 to address concerns regarding data quality in the early years of the period considered.

The data on institutional investor ownership are from the FactSet/LionShares database, which provides information on portfolio holdings for institutional investors worldwide. Ferreira and Matos (2008) and Aggarwal, Erel, et al. (2011) provide a more detailed description of this database. Because institutional ownership represented over 40% of total global stock market capitalization during our sample period, we control for institutional ownership in all our regressions. FactSet/LionShares also provides us with data on ETF ownership of stocks. The identity and replicating methods of ETFs (i.e., whether an ETF physically replicates its index), however, are provided by Morningstar. We match data from Morningstar with data from the FactSet/LionShares database and identify ETFs that fully replicate the indices they track using Morningstar and then use the latter database to aggregate ETF stock ownership.

We combine Datastream data with the short-selling and institutional holdings data by using SEDOL and ISIN codes for non-U.S. firms. We use CUSIP to merge short-selling data with U.S. security data from Datastream. The initial sample from the matched datasets of Datastream and DataExplorers covers 22,562 unique firms. After the match with FactSet/LionShares, the sample was reduced to 20,128 firms over the period considered. Countries like China, India, Malaysia, and Thailand, for instance, have been excluded due to the lack of short-selling information. We further require stocks to have nonmissing financial information on firm size, book-to-market ratio, financial leverage, annual stock return, and stock return volatility. These requirements reduce the number of stocks to 17,555 in thirty-three countries. Appendix B tabulates the number of stocks covered by each of these thirty-three countries over the sample period, from 3,637 non-U.S. firms and 1,193 U.S. firms in the year 2002 to 7,878 for non-U.S. firms and 4,031 for U.S. firms in December 2009. In the year 2008, for instance, we cover 13,082 stocks, a number comparable to the sample of 12,621 stocks in the same year in twenty-six countries in Saffi and Sigurdsson (2011). Regarding the coverage of market capitalization, the sample includes more than 90% of global stocks.

1.2 Main variables

Following the literature we use accruals as the main proxy for earnings management (e.g., Dechow, Ge, and Schrand 2010). Total accruals (*Accruals*) are calculated from balance sheet and income statement information. In particular, $Accruals = ((\Delta CA - \Delta Cash) - (\Delta CL - \Delta SD - \Delta TP) - DP)$, where

ΔCA is the change in current assets, $\Delta Cash$ is the change in cash and equivalents, ΔCL is the change in current liabilities, $\Delta \mathrm{SD}$ is the change in short-term debt included in the current liabilities, ΔTP is the change in income tax payable, and *DP* denotes depreciation and amortization expenses. All of the numbers are scaled by lagged total assets. Total accruals include discretionary and nondiscretionary components. Because nondiscretionary components depend on the economic performance of a firm—such as changes in revenues and the depreciation of fixed assets—the discretionary component can measure managerial discretion in reported earnings more precisely. Therefore, to measure the discretionary component of accruals, we rely on Dechow, Sloan, and Sweeney's (1995) modification of Jones's (1991) residual accruals ($Accrual_{MJones}$) as the main measure. $Accrual_{MJones}$ denotes the residuals obtained by regressing total accruals on fixed assets and revenue growth, excluding growth in credit sales, for each country and year.[5]

We proxy for our main measure of SSP by using *Lendable*—that is, the annual average fraction of shares of a firm that are available to be lent to short sellers. We rely on Saffi and Sigurdsson (2011) and compute the ratio between the values of shares that are supplied to the short-selling market (as reported by DataExplorers) and the market capitalization of the stock (as reported by Datastream), and we then define the time-series average of the monthly (weekly or daily) ratio as the annual *Lendable* ratio. We primarily consider the annual frequency because earnings management variables are defined annually. In addition to our main dependent and independent variables, we have also constructed alternative measures of earnings management and SSP. These variables will be detailed in subsequent sections when we conduct robustness checks.

Our control variables are the logarithm of firm size (*Size*), the logarithm of the book-to-market ratio (*BM*), financial leverage (*Leverage*), the logarithm of annual stock return (*Return*), stock return volatility (*STD*), American Depositary Receipts (*ADR*), MSCI country index membership (*MSCI*), the number of analysts following the firm (*Analyst*), closely held ownership (*CH*), institutional ownership (*IO*), and Amihud's (2002) illiquidity (*Illiquidity*).[6] Institutional ownership denotes the aggregate equity holdings by domestic and foreign institutional investors as a percentage of the total number of outstanding shares. Similarly, we also construct ETF ownership (*ETF*), which is defined as the percentage of the total number of outstanding shares that are invested by ETFs. Industry-level ETF ownership ($ETF_{Industry}$) is computed

5 Our results are also robust to regressions by industry and year, by country, or by year only. We do not run regressions by country, year, and industry, as many countries do not have sufficient observations to support regressions at the industry level.

6 Our results are also robust to alternative illiquidity measures, such as the proportion of zero daily returns in a year, the turnover ratio, proportional effective spread, and proportional relative spread. Here, we primarily rely on the Amihud measure because of its importance in the global market (e.g., Karolyi, Lee, and Van Dijk 2012). We tabulate the results for the alternative liquidity measures in Table IA11 in the Internet Appendix.

Table 1
Summary statistics

Panel A: Summary statistics

Variable	*N*	Mean	STD	10%	25%	Median	75%	90%
$Accrual_{MJones}$	61,624	0.002	0.081	−0.081	−0.034	0.003	0.037	0.080
Lendable	61,624	0.065	0.094	0.000	0.004	0.023	0.086	0.211
Size	61,624	12.939	1.821	10.725	11.673	12.803	14.108	15.369
BM	61,624	−0.622	0.879	−1.676	−1.133	−0.593	−0.067	0.419
Leverage	61,624	0.198	0.180	0.000	0.024	0.170	0.320	0.450
Return	61,624	0.023	0.649	−0.752	−0.241	0.094	0.373	0.678
STD	61,624	0.448	0.316	0.189	0.258	0.372	0.548	0.776
ADR	61,624	0.036	0.185	0.000	0.000	0.000	0.000	0.000
MSCI	61,624	0.663	0.473	0.000	0.000	1.000	1.000	1.000
Analyst	61,624	5.053	5.998	0.000	1.000	2.833	7.500	13.417
CH	61,624	0.311	0.241	0.002	0.109	0.285	0.491	0.653
IO	61,624	0.248	0.299	0.000	0.020	0.110	0.379	0.785
Illiquidity	61,624	−3.430	2.965	−7.353	−5.641	−3.371	−1.284	0.407

(*continued*)

as the equally weighted average of ETF ownership in any industry excluding the corresponding firm. A detailed definition of these variables is provided in Appendix A.

Table 1 presents the summary statistics for the main dependent, independent, and control variables. The summary statistics for all the other variables used in later sections are provided in Table IA2 in the Internet Appendix. From now on, we will define the tables contained in the Internet Appendix with the prefix "IA." Panel A reports the number of observations and the mean, median, standard deviation, and decile (90% and 10%) and quartile (75% and 25%) distributions of the variables. Panel B reports the correlation coefficients among the main variables. We calculate both the Spearman correlation coefficients and the Pearson correlation coefficients. The former are reported in the upper right part of the table, whereas the latter are reported in the bottom left part of the table.

We can see that our dependent and independent variables have reasonable variation. For example, the mean (6.5%) of *Lendable* is close to the mean (8.0%) of the lending supply variable in Saffi and Sigurdsson (2011) for firms with reasonable financial information. The remaining difference arises from two sources. First, we require firms to have valid annual earnings management variables to be included in our sample, while Saffi and Sigurdsson (2011) require weekly stock return information. Second, our final sample focuses on the testing period from 2002 to 2009, while their sample is from 2004 to 2008. Broadly speaking, the two sources explain approximately two-thirds and one-third of the residual difference, respectively. Table IA3 tabulates the average value of *Lendable* for each year and provides a more detailed comparison. Later sections will show that our results are robust to the inclusion/exclusion of the early years. Our results are also robust whether we include or exclude the firms for which no shares are available to be sold short (i.e., zero *Lendable*).

Table 1
Continued

Panel B: Correlation Coefficients (Spearman for the upper-right part, highlighted; Pearson for the bottom-left part)

Variable	$Accrual_{MJones}$	*Lendable*	*Size*	*BM*	*Leverage*	*Return*	*STD*	*ADR*	*MSCI*	*Analyst*	*CH*	*IO*	*Illiquidity*
$Accrual_{MJones}$	–	−0.050	0.024	0.002	0.015	−0.001	−0.042	−0.053	−0.011	−0.014	−0.029	0.051	−0.019
Lendable	−0.030	–	0.501	−0.170	0.013	−0.026	−0.135	0.074	0.344	0.577	−0.219	0.552	−0.538
Size	0.026	0.319	–	−0.319	0.157	0.143	−0.341	0.161	0.598	0.732	−0.066	0.383	−0.864
BM	0.015	−0.114	−0.294	–	0.085	−0.204	−0.080	−0.042	−0.103	−0.285	0.097	−0.215	0.258
Leverage	0.026	0.008	0.122	0.016	–	−0.001	−0.099	0.022	0.114	0.095	0.007	−0.017	−0.154
Return	0.023	−0.028	0.146	−0.199	−0.023	–	−0.002	0.005	0.077	−0.007	0.035	0.014	−0.034
STD	−0.040	−0.110	−0.309	−0.097	−0.055	0.068	–	−0.021	−0.168	−0.178	−0.039	−0.072	0.243
ADR	−0.046	0.053	0.193	−0.036	0.017	0.003	−0.019	–	0.062	0.143	−0.063	0.001	−0.116
MSCI	−0.004	0.246	0.559	−0.090	0.099	0.082	−0.166	0.062	–	0.430	−0.027	0.298	−0.625
Analyst	−0.021	0.348	0.728	−0.222	0.064	−0.013	−0.174	0.196	0.344	–	−0.112	0.443	−0.693
CH	−0.016	−0.255	−0.084	0.073	0.007	0.039	−0.038	−0.055	−0.049	−0.134	–	−0.191	0.209
IO	0.037	0.493	0.352	−0.207	−0.005	−0.007	−0.070	−0.048	0.281	0.355	−0.275	–	−0.435
Illiquidity	−0.022	−0.384	−0.868	0.235	−0.132	−0.038	0.249	−0.123	−0.609	−0.675	0.239	−0.469	–

This table presents the summary statistics and Spearman (Pearson) correlation coefficients of the main variables that are used in this study. The variables are modified Jones (1991) residual accruals ($Accrual_{MJones}$), lendable shares (*Lendable*), log of firm size (*Size*), log of book-to-market ratio (*BM*), financial leverage (*Leverage*), log of annual stock return (*Return*), stock return volatility (*STD*), American Depository Receipts (*ADR*), MSCI country index membership (*MSCI*), number of analysts following the firm (*Analyst*), closely held ownership (*CH*), institutional ownership (*IO*), and Amihud's (2002) illiquidity (*Illiquidity*). All of the variables are defined in Appendix A. Panel A reports the number of observations (N) and the mean, median, standard deviation (*STD*), and decile (10% and 90%) and quartile (25% and 75%) distributions of the variables. Panel B reports the correlation coefficients among the above variables, where the highlighted upper-right part (bottom-left part) of the table refers to the Spearman (Pearson) correlation matrix. The sample period is between 2002 and 2009.

Panel B illustrates that a negative correlation exists between discretionary accruals and SSP, suggesting a disciplining effect of short selling on earnings management. For example, the Pearson (Spearman) correlation coefficient between $Accrual_{MJones}$ and *Lendable* is −0.030 (−0.050), with a t-statistic of 8.23 (9.56), and its absolute magnitude is the fourth (second) largest among the Pearson (Spearman) correlation coefficients between other control variables and accruals. Although this result provides preliminary evidence of such a correlation, this correlation may be spurious. Thus, the next step of the analysis is to examine the relationship in a multivariate framework.

2. Short-Selling Potential and Earnings Management: Initial Evidence

We rely on the following regression as a baseline for our multivariate analyses:

$$Accrual_{MJonesi,t+1} = \alpha + \beta_1 \times Lendable_{i,t} + \beta_2 \times X_{i,t} + \varepsilon_{i,t}, \qquad (1)$$

where $Accrual_{MJonesi,t+1}$ refers to our main earnings management proxy for firm i in year $t+1$; $Lendable_{i,t}$ is the fraction of lendable shares for the same firm in the previous year t; and $X_{i,t}$ refers to a list of lagged control variables, including firm size, book-to-market ratio, financial leverage, annual stock return, stock return volatility, American depository receipts, MSCI country index membership, number of analysts following the firm, closely held ownership, institutional ownership, and Amihud's (2002) illiquidity. All the control variables are as of the previous year.

Table 2 reports the results of the regression with various econometric specifications. Model (1) presents our baseline specification, in which we include industry, country, and year fixed effects (ICY) and cluster standard errors at the firm level. This regression specification is the standard one in the literature when accruals are used as the dependent variable (e.g., Yu 2008; Francis and Wang 2008; Francis, Michas, and Seavey 2013).[7] The results show a strong negative correlation between SSP and earnings management. Specifically, a one-standard-deviation increase in SSP is associated with 5.12% standard deviation less earnings management.[8] This impact is both statistically significant and economically relevant. Models (2), (3), and (4) remove the year fixed effect, control for firm and year fixed effects, and control for firm fixed effects, respectively. Our main conclusions are robust across all the different specifications.

Next, Models (5) and (6) apply the dynamic-panel GMM estimator of Arellano and Bond (1991). This method exploits the lagged explanatory

[7] Compared with Saffi and Sigurdsson (2011), we further control for industry-level fixed effects to remove industry-specific factors that affect earnings management (e.g., Yu 2008; Dechow, Ge, and Schrand 2010). However, Saffi and Sigurdsson (2011) adopt firm-year double clustering following Petersen (2009) and Thompson (2010). Tables IA4 and IA13 will show that our results are robust to the use of double clustering.

[8] The economic impact is computed as the regression coefficient multiplied by the one-standard-deviation change in *Lendable*, which is scaled by the standard deviation of discretionary accruals in the sample.

Table 2
Short selling and earnings management

	Baseline Model		Firm Fixed Effects		Arellano-Bond	
	Model (1)	Model (2)	Model (3)	Model (4)	Model (5)	Model (6)
Lendable	**–0.044**	**–0.032**	**–0.017**	**–0.052**	**–0.041**	**–0.032**
	(–7.88)	**(–7.10)**	**(–2.33)**	**(–8.39)**	**(–7.38)**	**(–7.01)**
Lagged Accrual$_{MJones}$					−0.010	−0.009
					(−1.02)	(−1.00)
Size	0.004	0.005	0.022	0.021	0.003	0.004
	(7.10)	(8.09)	(12.44)	(12.14)	(5.98)	(6.99)
BM	0.004	0.005	0.013	0.013	0.004	0.005
	(6.96)	(7.60)	(8.15)	(8.63)	(7.02)	(7.74)
Leverage	0.010	0.010	0.070	0.071	0.011	0.011
	(4.51)	(4.52)	(10.04)	(10.22)	(4.65)	(4.71)
Return	0.003	0.001	−0.000	−0.000	0.003	0.002
	(3.63)	(1.99)	(−0.26)	(−0.49)	(3.90)	(2.42)
STD	−0.006	−0.004	−0.009	−0.007	−0.008	−0.006
	(−3.85)	(−2.81)	(−3.54)	(−2.90)	(−4.55)	(−3.59)
ADR	−0.013	−0.014	0.002	0.005	−0.013	−0.014
	(−5.82)	(−6.08)	(0.32)	(0.69)	(−5.96)	(−6.19)
MSCI	−0.006	−0.006			−0.006	−0.006
	(−5.57)	(−5.85)			(−5.62)	(−5.70)
Analyst	−0.001	−0.001	−0.001	−0.000	−0.001	−0.001
	(−10.29)	(−10.51)	(−3.63)	(−1.52)	(−10.33)	(−10.35)
CH	−0.004	−0.004	0.002	0.000	−0.004	−0.004
	(−2.19)	(−2.26)	(0.42)	(0.09)	(−2.10)	(−2.21)
IO	0.003	0.003	−0.005	−0.006	0.003	0.004
	(1.26)	(1.35)	(−0.83)	(−0.87)	(1.54)	(1.67)
Illiquidity	−0.001	−0.000	−0.003	−0.001	−0.001	−0.001
	(−2.48)	(−1.35)	(−4.16)	(−2.12)	(−3.76)	(−2.61)
Fixed Effects	ICY	IC	FY	F	ICY	IC
Obs	61,624	61,624	61,624	61,624	59,446	59,446
AdjRsq	2.9%	2.9%	19.4%	19.1%		

This table examines the baseline effect of short selling on earnings management. The main specification is based on a panel regression of a firm's modified Jones (1991) residual accruals ($Accrual_{MJones}$) on lendable shares (*Lendable*) and firm-level control variables (X), as well as unreported industry, country, and year fixed effects (ICY). The regression model is $Accrual_{MJones i,t+1} = \alpha + \beta_1 Lendable_{i,t} + \beta_2 X_{i,t} + \varepsilon_{i,t}$, where $X_{i,t}$ includes firm size (*Size*), book-to-market ratio (*BM*), financial leverage (*Leverage*), annual stock return (*Return*), stock return volatility (*STD*), American Depository Receipts (*ADR*), MSCI country index membership (*MSCI*), number of analysts following the firm (*Analyst*), closely held ownership (*CH*), institutional ownership (*IO*), and Amihud's (2002) illiquidity (*Illiquidity*). The construction of these variables is detailed in Appendix A. Models (1) and (2) control for industry, country, and year fixed effects (ICY) and industry and country fixed effects (IC), respectively. The t-statistics reported in parentheses are based on standard errors adjusted for heteroscedasticity and firm-level clustering. Models (3) and (4) control for firm and year (FY) and firm (F) fixed effects, with standard errors clustered at the year level. Models (5) and (6) apply the Arellano-Bond dynamic panel GMM estimation to the same relationship while controlling for industry, country, and year fixed effects (ICY) and industry and country fixed effects (IC), respectively. Obs denotes the number of firm-year observations, and AdjRsq is adjusted R^2. The sample period is from 2002 to 2009.

variables as instruments and is especially suitable for small time-series and large cross-sectional panels, providing unbiased and consistent estimates. The results show that the dynamic GMM estimator supports a negative correlation between SSP and earnings management.[9]

9 Based on the Arellano-Bond test for both AR(1) and AR (2) specifications and Hansen's test of overidentifying restrictions, our estimation satisfies the zero-autocorrelation request for the residuals.

Overall, our results provide consistent multivariate evidence that a higher level of SSP is associated with less earnings management in the future. Before we move on to develop a causal interpretation, it is also worthwhile to note that the parameters of the other variables are consistent with the existing literature on earnings management. For example, large-sized firms have aggressive accruals because of income-increasing accounting method choices (Watts and Zimmerman 1986). Being listed in the U.S. market (i.e., *ADR*) is negatively and significantly associated with a firm's accruals. This observation provides the first supporting evidence—in terms of earnings management—of the bonding hypothesis that cross-listings on U.S. stock exchanges strengthen corporate governance and protect outside investors (Doidge, Karolyi, and Stulz 2004).

3. Endogeneity Issues

The previous results, although favorable to the disciplining hypothesis, may be subject to the issue of endogeneity. We have already addressed the issue of spurious correlation through the omission of relevant firm-specific information with alternative fixed effect and dynamic panel data specifications. In this section, we address this issue using a twofold approach: we employ an instrumental variable specification and provide two regulatory experiments in which short selling is exogenously determined.

3.1 An instrumental variable approach

We begin with an instrumental variable specification. Relying on the findings of Hirshleifer, Teoh, and Yu (2011), we argue that the ownership of ETFs that fully replicate benchmarks can be used as an ideal instrument.[10] On the one hand, ETFs are among the main contributors to the short-selling market. Because the ETF industry thrives on its low-fee reputation, ETFs often lend out shares to the short-sellers to generate additional income that allows them to reduce fees. For instance, *iShares Russell 2000 Index Fund* (IWM), a $15 billion ETF with an expense ratio of 23 basis points (bps), generated 21 bps from security lending in a one-year period. Overall, *iShares* made $397 million in securities lending fees in 2011. On the other hand, ETFs are not typically concerned with enjoying active control over the managers of the firm because their goal is simply to replicate benchmarks. Moreover, precisely because ETFs replicate benchmarks rather than paying attention to the performance of individual stocks, the time-series variations in ETF ownership can only be attributed to investor flows related to benchmark characteristics, as opposed to stock-specific information.

10 Hirshleifer, Teoh, and Yu (2011) use institutional ownership as an instrument for the amount of lendable shares to proxy for the ease of short arbitrage. Our approach is in the same spirit, except we further require the instrument to be uncorrelated with internal governance. This consideration motivates us to use ETFs to identify the passive component of lendable supply as a general control to alleviate concerns related to internal governance.

These characteristics make the fraction of stock ownership held by ETFs a suitable instrument because it reasonably meets both the exclusion restriction (i.e., it is unrelated to earnings management except through the short-selling market) and the inclusion restriction (i.e., ETFs make shares available to short sellers). Moreover, the exogenous and high growth rate of the ETF industry over the past decade suggests that the instrument is likely to have power. In addition to ETF ownership at the stock level (*ETF*), we also utilize industry-level (excluding the specific firm) ETF ownership ($ETF_{Industry}$). The latter instrument has lower cross-sectional variation, but is less related to firm-specific characteristics.

Based on these two instruments, we perform a two-stage IV regression as follows. We regress SSP on either instrument in the first stage and then regress our earnings management measure ($Accrual_{MJones}$) on ETF ownership (*ETF*)–instrumented SSP in the second stage, together with firm-level control variables (X) and industry, country, and year fixed effects:

$$The\ 1st\ stage: Lendable_{i,t} = \alpha + \beta_1 ETF_{i,t} (or\ ETF_{Industryi,t}) + \beta_2 X_{i,t} + \varepsilon_{i,t},$$

$$The\ 2nd\ stage: Accrual_{MJonesi,t+1} = \alpha + \beta_1 \hat{Lendable}_{i,t} + \beta_2 X_{i,t} + \varepsilon_{i,t}. \qquad (2)$$

The results are tabulated in Table 3. In Panel A, Models (1) and (3) regress lendable shares on ETF ownership and industry-level ETF ownership, respectively. Models (2) and (4) regress earnings management on predicted lendable shares. If we focus on the first-stage regressions, we observe that SSP is strongly positively related to the fraction of ETF ownership at both the firm and the industry levels. A one-standard-deviation increase in firm-specific (industry-specific) ETF ownership is associated with 19.89% (38.53%) higher lendable shares. The t-statistics are always above 5, translating into an F-test of 29.78 in Model (1) and 28.13 in Model (3), both of which are well above the threshold of weak exogeneity provided by Staiger and Stock (1997).

The second-stage regressions show a strong negative correlation between instrumented SSP and earnings management. A one-standard-deviation increase in ETF-instrumented lendable shares is correlated with 9.61% lower discretionary accruals. Note that in all the regressions, we control for the level of a firm's institutional ownership, effectively controlling for any monitoring role played by institutional investors. Additional tests, reported in Table IA4, confirm that our results are robust to firm-year double clustering following Petersen (2009) and Thompson (2010), to later periods (2005–2009), to stocks that are members of the MSCI country index, and to the presence of country-level characteristics as additional controls. In unreported tests, we further orthogonalize ETF ownership with respect to a list of attention and liquidity variables, such as analyst following, news coverage, and Amihud's (2002) illiquidity measure, and the results remain identical.

As we have argued, the features of the ETF industry—for example, the passive nature of ownership, the supply of lendable shares motivated by fees, the

Table 3
Instrumental variable approach

A. ETF and industry-level ETF as instrumental variables

	Instrument=ETF		*Instrument=ETF$_{Industry}$*	
Dep. Variable=	*Lendable* (1st Stage)	*Accrual$_{MJones}$* (2nd Stage)	*Lendable* (1st Stage)	*Accrual$_{MJones}$* (2nd Stage)
	Model (1)	Model (2)	Model (3)	Model (4)
Instrument	**0.847**		**3.039**	
	(5.46)		**(5.30)**	
$\hat{Lendable}$		**–0.102**		**–0.041**
		(–4.69)		**(–1.86)**
Size	−0.004	0.004	−0.004	0.004
	(−9.08)	(6.51)	(−6.85)	(6.97)
BM	0.009	0.005	0.008	0.004
	(20.86)	(7.43)	(19.33)	(6.61)
Leverage	0.007	0.011	0.004	0.010
	(3.90)	(4.66)	(2.10)	(4.50)
Return	0.004	0.003	0.002	0.003
	(6.89)	(3.86)	(2.99)	(3.60)
STD	−0.006	−0.006	−0.005	−0.006
	(−5.88)	(−4.02)	(−4.92)	(−3.83)
ADR	0.004	−0.013	−0.002	−0.013
	(2.48)	(−5.67)	(−1.23)	(−5.84)
MSCI	0.015	−0.005	0.017	−0.006
	(15.71)	(−4.48)	(18.21)	(−5.37)
Analyst	0.001	−0.001	0.001	−0.001
	(11.55)	(−9.56)	(9.70)	(−10.05)
CH	−0.008	−0.004	−0.013	−0.004
	(−5.80)	(−2.44)	(−8.16)	(−2.17)
IO	0.115	0.011	0.132	0.002
	(20.03)	(2.93)	(40.23)	(0.61)
Illiquidity	−0.007	−0.001	−0.007	−0.001
	(−20.37)	(−3.36)	(−26.14)	(−2.17)
Fixed Effects	ICY	ICY	ICY	ICY
Obs	61,624	61,624	61,624	61,624
AdjRsq	65.6%	2.8%	67.4%	2.9%

(*continued*)

time-series variations in ownership attributable to benchmark-related investor flows—imply that ETFs do not directly affect managerial behavior. Although this implication is widely supported (e.g., Dyck, Morse, and Zingales [2010] show that ETFs do not blow the whistle on corporate fraud, although short sellers do), we provide additional evidence that the ETF ownership seems to affect earnings management only through its impact on short selling. Models (1) and (6) in Panel B report the regression of accruals on the two ETF instruments without SSP. If ETF ownership indirectly affects earnings management through SSP, we would expect ETF ownership to be significantly (and negatively) related to earnings management in the absence of SSP. However, if the inclusion of SSP removes the significance of the effect of ETF ownership on earnings management, it would provide further evidence supporting the

Table 3
(continued)

B. Tests on Exclusion Restrictions (Regress Accruals on ETFs)

			Accruals on ETFs when SSP is Low					Accruals on Industry ETF when SSP is Low		
			Legality=0	*SSban*=1	0 < *Lendable* < 0.5%			*Legality*=0	*SSban*=1	0 < *Lendable* < 0.5%
	Model (1)	Model (2)	Model (3)	Model (4)	Model (5)	Model (6)	Model (7)	Model (8)	Model (9)	Model (10)
ETF	**–0.086**	–0.016	**–0.371**	**–0.300**	**0.010**					
	(–3.57)	(–0.75)	**(–0.50)**	**(–1.73)**	**(0.21)**					
$ETF_{Industry}$						**–0.124**	0.011	**–0.857**	**0.885**	**0.266**
						(–2.01)	(0.14)	**(–0.23)**	**(0.31)**	**(1.29)**
Lendable		**–0.040**					**–0.044**			
		(–6.99)					**(–7.03)**			
Firm Controls and Constant	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Fixed Effects	ICY	ICY	ICY	ICY	ICY	ICY	ICY	ICY	ICY	ICY
Obs	61,624	61,624	1,025	3,159	16,578	61,624	61,624	1,025	3,159	16,578
AdjRsq	2.9%	2.9%	8.7%	5.0%	2.7%	2.8%	2.9%	8.7%	4.9%	2.8%

Panel A addresses the endogeneity problem by using ETF ownership (*ETF*) or industry-level ETF ownership ($ETF_{Industry}$) as an instrumental variable and presents a panel regression of a firm's earnings management measure ($Accrual_{MJones}$) on predicted lendable shares ($\hat{Lendable}$) and firm-level control variables (X) as well as unreported industry, country, and year fixed effects (ICY) on the variation of the following models:

$$\text{The 1 st stage: } Lendable_{i,t} = \alpha + \beta_1 ETF_{i,t}(ETF_{Industry\,i,t}) + \beta_2 X_{i,t} + \varepsilon_{i,t};$$

$$\text{The 2 nd stage: } Accrual_{MJones\,i,t+1} = \alpha + \beta_1 \hat{Lendable}_{i,t} + \beta_2 X_{i,t} + \varepsilon_{i,t},$$

where $Lendable_{i,t}$ refers to lendable shares and $X_{i,t}$ includes the list of standard control variables. Models (1) and (3) regress lendable shares on ETF ownership and industry-level ETF ownership, respectively. Models (2) and (4) regress modified Jones (1991) residual accruals on predicted lendable shares. Panel B provides the diagnostic analyses on the impact of *ETF* and $ETF_{Industry}$ on $Accrual_{MJones}$. Models (1) and (6) directly regress $Accrual_{MJones}$ on the two instruments. Models (2) and (7) also include lendable shares in the same regression. The remaining models regress $Accrual_{MJones}$ on the two instruments for subsamples of the stocks for which short selling is either prohibited owing to regulation (*Legality*=0 or *SSban=1*) or low—when very few shares could be lent out (0 < *Lendable* < 0.5%). The t-statistics reported in parentheses are based on standard errors adjusted for heteroscedasticity and firm-level clustering. Obs denotes the number of firm-year observations, and AdjRsq is adjusted R^2. The sample period is from 2002 to 2009.

exclusion restriction of the instruments.[11] Models (2) and (7) show that the inclusion of SSP does indeed remove the significance of the effect. While the two instruments are correlated with the dependent variable in general, they lose their explanatory power when the specific channel, SSP, is presented.[12]

To further illustrate this point, we reestimate our specifications within the subsamples of firms for which short selling is more constrained and report the results for Models (3)–(5) and (8)–(10). In Models (3)–(5), short selling is either prohibited owing to regulation ($Legality = 0$ or $SSban = 1$) or difficult to implement owing to the low supply of lendable shares ($0 < Lendable > 0.5\%$). We see again that firm-level ETF ownership loses its power in affecting earnings management, suggesting that short selling is the necessary channel for ETF ownership to affect earnings management. Similar subsample regressions for industry-level ETF ownership are estimated in Models (8)–(10), from which we obtain the same conclusion.[13] The key message is therefore that ETFs provide short sellers with the "ammunition" that they use to "discipline" managers. However, this provision of shares to short sellers alone is unlikely to directly affect earnings management.

In Table IA6, we also consider a further instrument: the degree of concentration of institutional ownership. Prado, Saffi, and Sturgess (2013) demonstrate that because high ownership concentration is related to investor activism, it typically introduces additional constraints into the lending market. In this regard, a lower degree of ownership concentration provides an instrument for the supply of lendable shares that is less related to investor activism. When this instrument is used alone or jointly with ETF ownership, our main findings remain unchanged. These instrumental variable regressions, therefore, provide the first evidence supporting the causal disciplining impact of SSP on earnings management.

3.2 An event-based approach: The U.S. SHO experiment

We now consider two regulatory "experiments" that have exogenously affected the ability to short sell. The advantage of this approach is that these policy events can create shocks and variations in short-selling costs that are orthogonal to

[11] Following Conley, Hansen, and Rossi (2012), when the dependent variable y, the independent variable x, and the instrument z are included in the same regression as $y = \beta \times x + \gamma \times z + \epsilon$, the exclusion restriction is equivalent to the condition of $\gamma = 0$. We also conduct overidentification tests based on the Hansen $J-$statistic in Table IA6.

[12] We also apply the plausible exogeneity test of Conley, Hansen, and Rossi (2012) to compute the confidence interval of β when the γ coefficient may take nonzero values (β and γ are specified in note 9). For instance, when firm-level ETF ownership is used as the instrument, the 95% confidence interval for the impact of SSP can be estimated as $[-.169, -.035]$ under the assumption that γ may take values from the support of $[-2\delta, 2\delta]$, where δ is the estimated standard deviation of γ. The results further confirm that the negative impact of SSP is statistically robust.

[13] ETF ownership, however, is not a necessary condition for SSP to affect earnings management. We show in Table IA5 that the disciplining effect of short selling is not attenuated when the level of ETF ownership is low. This result is logical because other (passive) institutional investors, such as pension funds and insurance companies, may also be willing to lend shares to short sellers.

firm-specific spurious correlation and endogeneity. We begin with the changes in short-sale price restrictions under Regulation SHO.

In this U.S. experiment, the SEC established a pilot program to exempt one-third of the stocks in the Russell 3000 Index from uptick rules and other price restrictions (e.g., Diether, Lee, and Werner 2009; Grullon, Michenaud, and Weston 2012).[14] The stocks were selected at random. As described in SEC Release No. 50104, the regulator "sorted the securities into three groups—Amex, NASDAQ, and NYSE—and ranked the securities in each group by average daily dollar volume over the year prior to the issuance of the order from highest to lowest for the period. In each group, we then selected every third stock from the remaining stocks.[15] Thus, the SEC essentially generated a randomized experiment that we can exploit to assess whether a relaxation in short-selling restrictions, which exogenously enhances SSP, translates into more effective disciplining. We therefore relate earnings management to an indicator of whether the restrictions have been lifted for the specific stock.

We begin by directly relating earnings management to the removal of the uptick restrictions. Specifically, we estimate the following annual panel regressions with firm and year fixed effects (FY):

$$Accrual_{MJonesi,t+1} = \alpha + \beta_1 USSHO_i \times Dummy(2005-2007) + \beta_2 USSHO_i \times Dummy(2008-2009) + \beta_3 X_{i,t} + \varepsilon_{i,t}, \quad (3)$$

where $Accrual_{MJonesi,t+1}$ is modified Jones's (1991) residual accruals, $USSHO_i$ refers to the dummy variable, which equals 1 if the stock is selected as a SHO pilot firm, and $X_{i,t}$ refers to a list of control variables. The variable $Dummy(2005-2007)$ takes a value of 1 for the period between 2005 and 2007, when pilot firms face fewer short-selling restrictions and thus higher SSP than the control group. The variable $Dummy(2008-2009)$ takes a value of 1 for the period from 2008 to 2009, in which the regulatory difference between pilot firms and the control group vanishes. The latter period is selected to provide placebo tests of our analyses: if the difference in earnings management between pilot and control firms is indeed driven by the difference in short-selling restrictions between the two groups of firms in the first period of 2005–2007, the difference in earnings management should vanish when the regulatory difference evaporates in the latter period.

We report the results in Table 4, Models (1)–(3). The testing period in Model (1) is from 2001 to 2007, in which the announcement year (2004) of Regulation SHO is removed from the sample. In Model (2), the sample period is expanded

[14] The regulation was announced in 2004 and implemented in 2005. Because firms may have begun reducing earnings management immediately after the announcement of the policy, the important change in management for our purposes is from 2003 to 2005, not from 2004 to 2005.

[15] The details are available at http://www.sec.gov/rules/other/34-50104.htm. The experiment was performed by the Office of Economic Analysis.

Table 4
The U.S. Regulation SHO Experiment and Earnings Management

Dep. Variable=	$Accrual_{MJones}$				$\Delta Accrual_{MJones}$ *(After-Before)*	
	2001–2007 Ex. 2004	2001–2009 Ex. 2004	2001–2009 Ex. 2004–2007	After= Before=	2005–2007 2001–2003	2008–2009 2001–2003
	Model (1)	Model (2)	Model (3)		Model (4)	Model (5)
US SHO × Dummy (2005–2007)	**–0.006**	**–0.006**		*US SHO*	**–0.005**	**–0.003**
	(–2.20)	**(–2.16)**			**(–2.05)**	**(–0.88)**
US SHO × Dummy (2008–2009)		**–0.004**	**–0.003**			
		(–0.99)	**(–0.79)**			
Size	0.013	0.011	0.006	Δ*Size*	0.000	0.001
	(3.27)	(3.72)	(1.69)		(0.06)	(0.20)
BM	0.008	0.006	0.001	Δ*BM*	0.001	−0.004
	(2.54)	(2.26)	(0.44)		(0.32)	(−1.31)
Leverage	0.055	0.043	0.027	Δ*Leverage*	0.008	0.001
	(4.84)	(4.58)	(2.29)		(0.67)	(0.10)
Return	0.006	0.004	0.001	Δ*Return*	0.018	0.003
	(2.12)	(1.70)	(0.40)		(3.79)	(0.56)
STD	−0.018	−0.018	−0.020	Δ*STD*	−0.026	−0.026
	(−3.52)	(−4.57)	(−4.18)		(−3.67)	(−3.65)
Analyst	−0.002	−0.002	−0.002	Δ*Analyst*	−0.001	−0.001
	(−5.16)	(−5.87)	(−4.71)		(−1.81)	(−2.17)
CH	−0.001	−0.003	−0.007	Δ*CH*	0.009	0.000
	(−0.08)	(−0.45)	(−0.86)		(0.86)	(0.03)
IO	−0.017	−0.017	−0.010	Δ*IO*	−0.013	−0.013
	(−1.86)	(−2.27)	(−0.89)		(−1.27)	(−1.13)
Illiquidity	−0.006	−0.006	−0.005	Δ*Illiquidity*	−0.007	−0.005
	(−3.63)	(−4.16)	(−2.92)		(−3.68)	(−2.21)
Fixed Effects	FY	FY	FY	Fixed Effects	I	I
Obs	12,597	16,347	10,020	Obs	2,218	2,105
AdjRsq	15.4%	13.3%	13.8%	AdjRsq	8.1%	6.6%

This table examines Regulation SHO in the United States, in which the SEC randomly selected a sample of pilot firms announced in 2004 and formally removed their uptick restrictions from 2005 to 2007. Models (1)–(3) estimate the following annual panel regressions with firm and year fixed effects (FY):

$$Accrual_{MJonesi,t+1} = \alpha + \beta_1 USSHO_i \times Dummy(2005-2007) + \beta_2 USSHO_i \times Dummy(2008-2009) + \beta_3 X_{i,t} + \varepsilon_{i,t},$$

where $Accrual_{MJonesi,t+1}$ is modified Jones (1991) residual accruals, $USSHO_i$ refers to a dummy variable that equals 1 if the stock is selected as a SHO pilot firm, and $X_{i,t}$ refers to a list of control variables. The testing period in Model (1) is from 2001 to 2007, in which the announcement year (2004) of Regulation SHO is removed from the sample. In Model (2), the sample period is from 2001 to 2009, excluding 2004, and in Model (3), the sample period is from 2001 to 2009, excluding 2004–2007. Models (4) and (5) estimate the following cross-sectional regression:

$$\Delta Accrual_{MJonesi,t+1} = \alpha + \beta_1 USSHO_i + \beta_2 \Delta X_{i,t} + \varepsilon_{i,t}.$$

where $\Delta Accruals_{MJonesi,t+1}$ refers to the difference between the three-year average value of $Accrual_{MJones}$ after year 2004 (from 2005 to 2007 in Model (4) and from 2008 to 2009 in Model (5)) and that before 2004 (from 2001 to 2003) and $\Delta X_{i,t}$ refers to changes in the average value of the control variables over the same periods. Control variables are detailed in Appendix A. The *t*-statistics reported in parentheses are based on standard errors adjusted for heteroscedasticity and firm-level clustering. Obs denotes the number of firm-year observations, and AdjRsq is adjusted R^2.

to 2009 to allow for the placebo test. In Model (1), the sample period is from 2001 to 2009, excluding the event period of 2004–2007. Models (1) and (2) clearly show that lifting the short-selling restrictions is associated with a lower level of earnings management by the group of pilot firms relative to the control group. Exemption from the restrictions is associated with a 7.88% lower level of earnings management. It is interesting to note that in Models (2) and (3), the effect disappears after 2007, confirming that the aforementioned earnings management difference between pilot firms and control firms is specifically related to the relatively lower level of short-selling restrictions faced by the pilot firms.

In Models (4) and (5), we focus on a specification based on changes and estimate the following cross-sectional regression:

$$\Delta Accrual_{MJonesi,t+1} = \alpha + \beta_1 USSHO_i + \beta_2 \Delta X_{i,t} + \varepsilon_{i,t}, \tag{4}$$

where $\Delta Accruals_{MJonesi,t+1}$ refers to the difference between the three-year average value of $Accrual_{MJones}$ after year 2004 (from 2005 to 2007 in Model (4) and from 2008 to 2009 in Model (5)) and that before 2004 (from 2001 to 2003) and $\Delta X_{i,t}$ refers to changes in the average value of the control variables over the same periods.

We find a significant impact of the relative difference in short-selling restrictions on earnings management for the 2005 to 2007 period, but not afterward. The cross-sectional regression indicates that pilot firms subject to less restrictive regulation reduce their earnings management (relative to the control firms) by as much as 6.57%. When the regulations on the pilot and control firms converge, the difference between the two groups dissipates. This test and the previous tests (although different in nature) generate the same conclusion—that is, lifting short-selling restrictions reduces earnings management. The random nature of the experiment makes it impossible for any spurious cross-sectional correlation to dominate the negative correlation.

3.3 An event-based approach: The Hong Kong experiment

We now consider the introduction of regulated short selling into the Hong Kong Stock Exchange. The Hong Kong Stock Exchange provides a different experiment in which short selling was gradually introduced in the market (see Chang, Cheng, and Yu 2007). The most interesting feature of this experiment is that the list of firms eligible for short-selling changes over time, which creates both time-series and cross-sectional variation in terms of short-selling restrictions for firms listed in Hong Kong. Stocks were added at the discretion of the regulator as a function of "changing market conditions," initially at irregular frequency and subsequently, after February 12, 2001, on a quarterly basis according to a set of criteria primarily based on market capitalization, turnover, index membership, and derivative contracts written on shares. Although these selection conditions make the experiment less clean than the SHO experiment, the selection remains unlikely to create spurious correlation because we explicitly control for the relevant variables. Moreover, the use of firm fixed effects helps us reduce the effects of any other omitted firm-specific characteristics that may have led to the introduction of short selling.

Table 5
Hong Kong short-selling list and earnings management

Dep. Variable=	$Accrual_{MJones}$				$\Delta Accrual_{MJones}$	
	Model (1)	Model (2)	*HK SS=0* Model (3)		Model (4)	Model (5)
HK SS	**–0.022**	**–0.024**		Δ*HK SS*	**–0.023**	
	(–2.18)	**(–2.25)**			**(–1.98)**	
$HK\ SS_{Post}$		**–0.009**	**0.002**	$\Delta HK\ SS_{Exclusion}$		**0.036**
		(–0.77)	**(0.12)**			**(2.25)**
Size	0.045	0.045	0.049	Δ*Size*	0.086	0.086
	(6.05)	(6.05)	(5.48)		(8.41)	(8.43)
BM	0.013	0.014	0.015	Δ*BM*	0.048	0.047
	(2.00)	(2.01)	(1.84)		(4.17)	(4.16)
Leverage	0.136	0.136	0.146	Δ*Leverage*	0.329	0.328
	(3.65)	(3.66)	(3.18)		(5.62)	(5.62)
Return	−0.001	−0.001	−0.003	Δ*Return*	−0.001	−0.000
	(−0.18)	(−0.20)	(−0.41)		(−0.10)	(−0.04)
STD	0.003	0.003	0.007	Δ*STD*	0.011	0.011
	(0.37)	(0.37)	(0.66)		(0.92)	(0.93)
ADR	−0.148	−0.148	−0.151	Δ*ADR*	−0.127	−0.125
	(−5.10)	(−5.14)	(−3.23)		(−1.28)	(−1.25)
Analyst	−0.000	−0.000	−0.001	Δ*Analyst*	−0.001	−0.001
	(−0.06)	(−0.05)	(−1.03)		(−0.89)	(−1.07)
CH	0.018	0.018	0.033	Δ*CH*	0.020	0.020
	(0.78)	(0.78)	(1.14)		(0.65)	(0.65)
Illiquidity	0.002	0.002	0.001	Δ*Illiquidity*	0.001	0.001
	(0.92)	(0.91)	(0.42)		(0.37)	(0.34)
Fixed Effects	FY	FY	FY	Fixed Effects	CY	CY
Obs	4,411	4,411	3,483	Obs	3,528	3,528
AdjRsq	10.8%	10.7%	10.3%	AdjRsq	6.2%	6.2%

This table explores the unique regulatory setting in the Hong Kong market in which regulators changed the list of stocks eligible for short selling on a quarterly frequency from 1994 to 2005. Models (1)–(3) estimate the following panel regression with firm and year fixed effects (FY) and clustered standard errors at the firm and industry levels:

$$Accrual_{MJonesi,t+1} = \alpha + \beta_1 HKSS_{i,t} + \beta_2 HKSS_{Posti,t} + \beta_3 X_{i,t} + \varepsilon_{i,t}.$$

where $Accrual_{MJonesi,t+1}$ is modified Jones (1991) residual accruals, $HKSS_{i,t}$ is a dummy variable that equals 1 if a stock is eligible to short selling in year t, and $HKSS_{Posti,t}$ is a dummy variable that equals 1 if a stock is eligible for short selling in year *t–1* but becomes ineligible for short selling beginning in year t. Models (4) and (5) estimate the following panel regressions:

$$\Delta Accrual_{MJonesi,t+1} = \alpha + \beta_1 \Delta HKSS_{i,t} (\Delta HKSS_{Exclusioni,t}) + \beta_2 \Delta X_{i,t} + \varepsilon_{i,t}.$$

where $\Delta HKSS_{i,t}$ refers to net inclusion and equals 1 (–1) if a firm is included in (excluded from) the eligible list and $\Delta HKSS_{Exclusioni,t}$ is a dummy variable for exclusion. The control variables are detailed in Appendix A. The t-statistics reported in parentheses are based on standard errors adjusted for heteroscedasticity and firm-level clustering. Obs denotes the number of firm-year observations, and AdjRsq is adjusted R^2.

We therefore regress $Accrual_{MJonesi,t+1}$ on a dummy variable, $HKSS_{i,t}$, which equals 1 if stock i is eligible for short selling in year $t$in a panel regression with firm and year fixed effects and standard errors clustered at the firm level. We report the results in the first three models of Table 5. The results indicate that short-selling eligibility reduces earnings management by 15.53%, which is even stronger in terms of economic magnitude, compared with the SHO experiment. In a placebo test, we further define $HKSS_{Posti,t}$ as a dummy variable that equals 1 if a stock is eligible for short selling in year $t-1$ but becomes ineligible for short selling beginning in year t. We find that, once short selling becomes unavailable again—even when it was feasible one year beforehand—the firm no longer exhibits lower earnings management. This placebo test further confirms that short selling—not persistent firm characteristics—reduces earnings management.

Next, in Models (4) and (5), we consider a difference-in-differences specification in which we regress accrual changes on two variables—namely, $\Delta HKSS_{i,t}$, which refers to net inclusion and equals 1 (-1) if a firm in included (excluded) in the eligible list, and $\Delta HKSS_{Exclusion\ i,t}$, which is a dummy variable for exclusion. We find that the inclusion (exclusion) of a firm in (from) the eligible list reduces (increases) management by 16.23% (25.41%).[16] These results are consistent with the results from the U.S. regulation.

We further verify that our results are robust when we exclude the observations of eligible firms for the year prior to their inclusion to the short-selling list and when we apply the tests to the propensity score-matched sample (Table IA7). The first test aims to reduce the potential contamination when firms can, to some extent, anticipate the inclusion of the firm in the eligible list (this contamination works against us only in our main tests), and the second test confirms that disciplining effect can be observed even among firms with close characteristics. Thus, although different in nature, the Hong Kong experiment leads to a similar conclusion to that of the U.S. experiment: short selling is important in curbing the incentives for earnings management. By contrast, we do not find evidence supporting alternative hypotheses that short selling does not affect or may further distort firm incentives.

4. Extensions and Additional Tests

4.1 The disciplining implications of global short-selling regulations

The previous section shows that regulatory restrictions on short selling in the United States and Hong Kong affect earnings management. Given the importance of regulation in the global financial market, we first explore whether the impact of short-selling regulation on earnings management can also be observed in other markets.

We therefore construct a list of dummy variables that describe the ease of short selling at the country level. We assign a value of 1 if short selling is legal (*Legality*), if short selling is feasible (*Feasibility*), if put option trading is allowed (*Put Option*), and if short selling is feasible or if put option trading is allowed (*F or P*). The difference between legality and feasibility is that the latter requires not only short selling to be legal but also the existence of institutional infrastructures that support short selling, including a low cost of short selling and the availability of market makers who are willing to trade

[16] We divide the coefficient of $HK\ SS$ ($\Delta HK\ SS$) by the standard deviation of discretionary accruals in Hong Kong to obtain the magnitude.

on short positions. Whenever we use these variables, we also further control for a set of commonly used, country-specific variables in addition to firm characteristics. Additional details on the country regulation variables and the additional control variables are provided in Table IA1.

To save space, we also tabulate the regressions in Table IA8 and report only the main results here. Our tests reveal a strong negative correlation between market-level SSP and earnings management. For example, in countries in which short selling is banned (unfeasible), earnings management appears to be 11% (14%) higher than in countries in which short selling is legal. Thus, the impact of short-selling regulations on earnings management applies to the global market.

Next, we further conduct the market-wide SSP tests on the sample of ADR firms. This sample is particularly interesting because of the nature of the ADR market. All ADR firms are exposed to the U.S. regulatory environment, which is known to promote firm value and corporate governance (e.g., Doidge, Karolyi, and Stulz 2004, 2007; Karolyi 2004, 2006). If U.S. regulations are perfectly enforced, the link between a firm's earnings management incentives and its home-country short-selling regulation (as we have just observed) should be completely suppressed. With the growing importance of ADRs in the global market, it is important to determine whether concurrent U.S. regulations suffice to achieve this goal: different answers to this question may lead to drastically different policy implications.

We therefore reestimate the panel regressions within the sample of ADR stocks. The results, also tabulated in Table IA8, show that the negative correlation between market-level SSP and earnings management is not reduced by the ability of the firms to bond to the U.S. regulatory environment through ADR listing. In other words, the disciplining impact of short selling in the global market is a strong and distinct effect, enough to survive the additional regulatory requirements that the U.S. market may impose or any additional governance improvements that ADR firms might experience. This surprising result adds to our understanding of the complexity of financial regulation in the global market.

Apart from country-level bans, other forms of short-selling restrictions may appear even in markets allowing for short selling. According to Jain et al. (2013), short-selling restrictions in general include specific trading mechanisms, pre-borrowing requirements (e.g., naked short selling), and bans on shorting selected stocks. Due to the lack of data on naked short selling as well as the specialty of bans on selected stocks (e.g., financial stocks), we follow Cumming, Johan, and Li (2011) and consider one specific and one general form of trading rules. Also, we extend our cross-country analysis to broader regulations that aim to improve the quality of information supplied to the market—that is, disclosure requirements and investor protections. We use the disclosure requirement index of La Porta, Lopez-de-Silanes, and Shleifer (2006) to proxy for accounting regulations, and the anti-director index of Pagano and Volpin (2005) to proxy for investor protections.

These results are reported in Table IA9. More specifically, we divide the whole sample into two subsamples according to these disclosure and trading rules, and conduct our main tests in these subsamples. Our main result is that markets with weaker accounting regulation and investor protections are typically associated with higher impact of SSP, suggesting that more regulated information disclosure/investor protection and SSP substitute each other in disciplining firms. Furthermore, low trading restrictions may enhance both the disciplining impact and the price pressure of SSP. Our empirical results document that the disciplining impact dominates. These cross-country tests, therefore, complement the stock-level tests and reinforce our major conclusions.

4.2 Time-series and subsample analyses

In this section, we examine the main results that we presented in Table 2 by considering various subsamples for robustness. We begin with the important observation made by Saffi and Sigurdsson (2011) that short-selling activity grew tremendously from 2004 to 2008 (their testing period) because of the boom of hedge funds that short sold stocks at a large scale. We would therefore expect that the disciplining impact of short selling on earnings management also increases over time. Saffi and Sigurdsson (2011) also illustrate that the short-selling market experienced drastic changes during the global financial crisis. Finally, DataExplorers provides only monthly information for 2002 and 2003 (weekly information from July 2004 on and daily information after 2006), and within this period the coverage is low, which may give rise to data quality concerns in the early years of the period considered. These considerations lead us to explore the impact of SSP over time.

The results are presented in Panel A of Table 6. In Model (1), we consider the sample period beginning in 2005 (inclusive) ("≥2005"). The results indicate that the exclusion of earlier years does not change the disciplining impact of SSP. This result is expected, as the inclusion of earlier years in our main test would work against our hypothesis only if the sample were less reliable. In addition, we also find that, as reported in Table IA10, the disciplining role in general applies to both the crisis and the noncrisis periods.

In Model (2), we include the interaction between lendable shares and a time-sequence variable T, which equals the year minus 2001, in the baseline regression. We find that this interaction negatively and significantly correlates with earnings management, suggesting that the impact of SSP on earnings management increases over time. A robustness check is also provided in Table IA10, in which we decompose the impact of lendable shares in various subperiods, and find that the impact of *Lendable* is significant in all subperiods and that its magnitude increases over time.

Next, we consider different regions/subsamples and report the results in Models (3) and (4). The samples "U.S." and "N.U.S." refer to firms from the U.S. and non-U.S. countries, respectively. We again find that the link

Table 6
Subsample analyses

A. Subsample analyses on different time periods and regions

	>=2005	Full Sample	U.S.	N.U.S.
	Model (1)	Model (2)	Model (3)	Model (4)
Lendable	**–0.045**	**–0.003**	**–0.064**	**–0.044**
	(–6.37)	**(–0.14)**	**(–7.39)**	**(–3.52)**
Lendable × *T*		**–0.006**		
		(–2.17)		
Firm Controls and Constant	Yes	Yes	Yes	Yes
Fixed Effects	ICY	ICY	ICY	ICY
Obs	44,171	61,624	21,825	39,799
AdjRsq	2.7%	2.9%	3.3%	2.8%

B. Subsample analyses on different degrees of earnings management

	$Accrual_{MJones} \geq 0$		$Accrual_{MJones} < 0$	
	Small Firms Model (1)	Large Firms Model (2)	Small Firms Model (3)	Large Firms Model (4)
Lendable	**–0.097**	**–0.025**	**–0.038**	**0.004**
	(–4.27)	**(–3.66)**	**(–1.55)**	**(0.55)**
Firm Controls and Constant	Yes	Yes	Yes	Yes
Fixed Effects	ICY	ICY	ICY	ICY
Obs	6,022	26,446	6,224	22,932
AdjRsq	6.9%	8.7%	13.3%	10.0%

This table examines the impact of short selling on earnings management in several important subsamples. Panel A explores the impact of short selling in subperiods, when it is interacted with various time dummies, and different regions. In the first two columns, the variable "≥2005" refers the sample period from 2005 to 2009, and T equals the year minus 2001. In Models (3) and (4), "U.S." and "N.U.S." refer to the subsamples of U.S. and non-U.S. firms, respectively. Panel B explores the source of the effects of short selling on earnings management by dividing the full sample into the accrual and firm size subsamples. $Accrual_{MJones} \geq 0$ refers to firms with positive $Accrual_{MJones}$, whereas $Accrual_{MJones} < 0$ refers to firms with negative $Accrual_{MJones}$. Small firms are firms with market capitalization below the median value of market capitalization for each country and year, whereas large firms are firms with market capitalization greater than the median value of market capitalization for each country and year. The t-statistics reported in parentheses are based on standard errors adjusted for heteroscedasticity and firm-level clustering. Obs denotes the number of firm-year observations, and AdjRsq is adjusted R^2. The sample period is from 2002 to 2009.

between SSP and earnings management holds across different subsamples. It is interesting to note that the disciplining impact is relatively stronger in the United States, suggesting that the real impact of the short-selling market in the United States is more prominent. This result can likely be explained by its well-developed institutional infrastructures to support short selling in the United States. Nonetheless, the disciplining impact of short selling applies to the global market and is not limited to the United States alone.[17]

Finally, we investigate whether the impact of short selling concentrates in firms with aggressive (i.e., inflated) earnings, as opposed to conservative (i.e., deflated) earnings, and whether short selling disciplines small stocks

[17] We also examine the impact of short selling on earnings management in subsamples of countries sorted by the feasibility of short selling and by controlling for firm-level investments. The results are tabulated in Table IA10.

for which investors have less information to a greater extent. Therefore, in Panel B, we divide the full sample into the accrual and firm size subsamples. $Accrual_{MJones} \geq 0$ refers to firms with positive $Accrual_{MJones}$ as a proxy for aggressive earnings, whereas $Accrual_{MJones} < 0$ refers to firms with negative $Accrual_{MJones}$ as a proxy for conservative earnings. *Small firms* and *large firms* are firms with market capitalization below and above the median value of market capitalization for each country and year, respectively. We find that short selling affects only firms with aggressive earnings, not those with conservative earnings; this result is exactly what should be expected from the disciplining effect. Also, the disciplining effect applies to both large and small firms, but its impact on the latter is greater in magnitude. This observation is consistent with the notion that the disciplining impact of short selling, which arises from the additional informativeness introduced by the short-selling market, should be more pronounced for the sector of the market that has less public information.

4.3 Robustness checks

Next, we consider a series of robustness checks to determine whether the disciplining effect is robust to the use of alternative SSP proxies and different earnings management measures, including not only various types of discretionary accruals but also a list of target-beating, earnings persistence, and earnings misstatement measures. We further examine whether the disciplining effect is robust to alternative discipline channels and clustering specifications.

Panel A of Table 7 considers alternative SSP measures. These variables are more endogenous than *Lendable* in describing the ex ante impact of short selling; hence, we only consider them as a robustness check. The first variable, *On Loan*, is the annual average fraction of the shares of a firm that are lent out (or short interest). A high level of realized historical short interest confirms a high level of short seller attention, which should discipline the earnings management incentives of firms in the future.

Next, a high short-selling fee implies lower SSP, which should be associated with a less effective disciplining impact. We can assess this potential association by using three variables: *Fee* is the annual value-weighted average loan fee expressed as a percentage; *STDFee* is the standard deviation of monthly value-weighted average loan fee expressed as a percentage; and *Specialness* is a dummy variable that proxies for very high short-selling costs and equals 1 if the average loan fee is above 1%.

The problem with relying on the short-selling fee is that it can be substantially affected by the demand side of the short-selling market (Cohen, Diether, and Malloy 2007; Kolasinski, Reed and Ringgenberg 2013)—and hence is less exogenous. To overcome this issue, we also define a new variable: *Constraint*, which is a dummy variable that equals 1 if the stock's average number of shares on loan is in the bottom quartile and its average loan fee is in the top quartile (quartiles are sorted by country and year). That is, when *Constraint* takes a value of 1, the stock will be very difficult to short sell because of its

Table 7
Robustness checks

A. Alternative SSP measures

SSP=	*On Loan*	*Fee*	*STD_{Fee}*	*Specialness*	*Constraint*
	Model (1)	Model (2)	Model (3)	Model (4)	Model (5)
SSP	**–0.052**	**0.001**	**0.003**	**0.004**	**0.004**
	(–4.12)	**(3.74)**	**(4.93)**	**(4.05)**	**(2.35)**
Firm Controls and Constant	Yes	Yes	Yes	Yes	Yes
Fixed Effects	ICY	ICY	ICY	ICY	ICY
Obs	61,623	61,575	59,011	61,575	61,574
AdjRsq	2.9%	2.9%	2.9%	2.9%	2.9%

B. Alternative earnings management measures on target beating and earnings persistence

	Target beating			**Earnings persistence**	
Dep. Variable=	*SPAF*	*SPDE*	*SPE*	*Earnings*	*$Accrual_{MJones}$*
	Model (1)	Model (2)	Model (3)	Model (4)	Model (5)
Lendable	**–0.767**	**–1.111**	**–0.914**	**0.033**	**–0.019**
	(–3.63)	**(–3.99)**	**(–2.06)**	**(4.50)**	**(–3.73)**
Earnings				**0.805**	
				(12.68)	
Lendable × Earnings				**–0.301**	
				(–4.66)	
$Accrual_{MJones}$					**0.109**
					(0.96)
Lendable × $Accrual_{MJones}$					**–0.453**
					(–4.15)
Firm Controls and Constant	Yes	Yes	Yes	Yes	Yes
Earnings ($Accrual_{MJones}$) × Control	No	No	No	Yes	Yes
Fixed Effects	ICY	ICY	ICY	ICY	ICY
Obs	46,381	35,986	19,091	58,302	55,816
AdjRsq (PseRsq)	3.4%	4.6%	10.2%	68.1%	5.5%

(*continued*)

limited supply and high cost—and we expect earnings management to increase as a consequence. *Constraint* is related to the variable for supply shocks in Cohen, Diether, and Malloy (2007). The difference between these variables is that supply shocks are defined by using changes in short interest and fees, while *Constraint* is related to the levels of short interest and fees. We focus on *Constraint* because the first-order impact of disciplining should be related to the level of short selling, though unreported tests using supply shocks lead to a similar conclusion.

For all these variables, we find a strong negative correlation between SSP and earnings management. A one-standard-deviation increase in SSP is associated with 2.32% (2.11%, 2.64%, 2.42%, and 1.19%) less earnings management in the case of *On Loan* (*Fee*, *STD_{Fee}*, *Specialness*, and *Constraint*, respectively). Thus, fewer restrictions in the short-selling market are associated with lower earnings management incentives.

Table 7
(continued)

C. Alternative earnings management measures on other accruals and earnings misstatements

	Other accrual measures			Misstatement and scandals	
Dep. Variable=	$Accrual_{Jones}$ Model (1)	$Accrual_{KLW}$ Model (2)	$Accrual_{DD}$ Model (3)	*Prob(Misstatement)* Model (4)	*Prob(Scandals)* Model (5)
Lendable	**–0.038**	**–0.035**	**–0.019**	**–0.818**	**–1.209**
	(–6.95)	**(–4.15)**	**(–3.83)**	**(–2.23)**	**(–2.58)**
Firm Controls and Constant	Yes	Yes	Yes	Yes	Yes
Fixed Effects	ICY	ICY	ICY	ICY	ICY
Obs	61,562	61,015	57,603	30,047	30,047
AdjRsq (PseRsq)	3.1%	0.5%	1.2%	5.8%	17.6%

D. Alternative discipline channels

	Model (1)	Model (2)	Model (3)	Model (4)	Model (5)
Lendable	**–0.027**	**–0.028**	**–0.032**	**–0.033**	**–0.022**
	(–3.06)	**(–3.04)**	**(–3.43)**	**(–3.62)**	**(–2.32)**
ISS	–0.025	–0.025	–0.025	–0.025	–0.023
	(–2.92)	(–2.95)	(–2.97)	(–2.94)	(–2.52)
IAS		–0.000	–0.001	–0.001	0.001
		(–0.00)	(–0.15)	(–0.23)	(0.29)
BigN			–0.008	–0.008	–0.005
			(–3.97)	(–4.13)	(–2.22)
NewsCoverage				–0.002	–0.003
				(–3.16)	(–3.75)
Disp					–0.308
					(–3.31)
Firm Controls and Constant	Yes	Yes	Yes	Yes	Yes
Fixed Effects	ICY	ICY	ICY	ICY	ICY
Obs	16,184	16,149	16,065	16,065	13,396
AdjRsq	4.7%	4.7%	4.8%	4.8%	6.1%

This table presents a series of robustness checks based on alternative short selling and earnings management measures, alternative controls for corporate governance, and alternative clustering specifications of the main regression model. Panel A presents the results for regressions using alternative short-selling potential (SSP) measures that include shares on loan (*On Loan*), loan fee (*Fee*), loan fee volatility (STD_{Fee}), specialness (*Specialness*), and short-selling constraint (*Constraint*). Panel B presents the results for regressions using alternative earnings management measures, including three target-beating measures—namely, small positive forecasting profits (*SPAF*), small positive past-earnings profits (*SPDE*), and small positive profits (*SPE*), in Models (1)–(3) and two earnings persistence specifications in Models (4) and (5). In particular, Models (4) and (5) estimate the following regression model:

$$Earnings_{i,t+1}(Accrual_{MJonesi,t+1}) = \alpha\beta_1 Lendable_{i,t} + \beta_2 Lendable_{i,t} \times Earnings_{i,t}(Accrual_{MJonesi,t})$$

$$+\beta_3 Earnings_{i,t}(Accrual_{MJonesi,t}) + \beta_4 X_{i,t} + \beta_5 X_{i,t} \times Earnings_{i,t}(Accrual_{MJonesi,t}) + \varepsilon_{i,t},$$

where $Earnings_{i,t+1}$ is operating income scaled by lagged total assets and $Accrual_{MJonesi,t+1}$ is modified Jones (1991) residual accruals. Panel C provides alternative discretionary accrual measures in Models (1)–(3), including Jones (1991) residual accruals ($Accrual_{Jones}$), KLW's (2005) residual accruals ($Accrual_{KLW}$), and DD's (2002) residual accruals ($Accrual_{DD}$). Models (4) and (5) proxy for manipulation practices by the likelihood that earnings misstatements or scandals occur. Panel D adds alternative discipline channels as controls. These alternative discipline channels include the ISS corporate governance index (*ISS*), big N auditor (*BigN*), international accounting standard (*IAS*), news coverage (*NewsCoverage*), and analyst dispersion (*Disp*). All these tests control for industry, country, and year fixed effects (ICY). The t-statistics reported in parentheses are based on standard errors adjusted for heteroscedasticity and firm-level clustering. Obs denotes the number of firm-year observations, and AdjRsq is adjusted R^2. The sample period is from 2002 to 2009.

Next, we consider a set of additional types of earnings management that are widely used in the literature to proxy for managerial distortion. The first type concerns "target-beating measures" (e.g., Burgstahler and Dichev 1997; Degeorge, Patel, and Zeckhauser 1999), which capture the incentives for managers to avoid reporting small losses relative to their heuristic target of zero. Such incentives lead to a well-known "kink" in the distribution of reported earnings near zero—that is, a statistically small number of firms with small losses and a statistically large number of firms with small profits (e.g., Burgstahler and Dichev 1997). This type of earnings management is especially powerful to complement our existing tests of the price pressure hypothesis, as the concerns that the downward price pressure of short selling may amplify the negative impact of not meeting analyst or market expectations could incentivize firms to engage in more target-beating actions. Therefore, we directly examine whether short selling can still discipline this type of earnings management incentive.

We use three proxies to capture such distortion. The first proxy is *target beating on small positive forecasting profits* (*SPAF*), based on Degeorge, Patel, and Zeckhauser (1999). This variable is a dummy that equals 1 if the difference between reported earnings per share and forecasted earnings per share scaled by stock price is between 0% and 1%. The variable captures managers' incentives to meet or beat analyst forecasts by a small margin. The second proxy is *target beating on small positive past-earnings profits* (*SPDE*) based on Burgstahler and Dichev (1997). This variable is a dummy that equals 1 if the change in net income scaled by lagged total assets is between 0% and 1%. The third proxy is *target beating on small positive profits* (*SPE*). Based on Burgstahler and Dichev (1997), this variable is a dummy that equals 1 if net income scaled by lagged total assets is between 0% and 1%. The last two variables proxy for managers' incentives to meet or beat market expectations by a small margin, where market expectations are measured by the previous year's earnings or a general request for firms to not report losses.

Models (1)–(3) in Panel B test the impact of our main proxy for SSP on these alternative earnings management measures. We find that the presence of SSP reduces the incentives to beat analyst or market expectations across all three measures of target-beating behavior. Thus, SSP not only disciplines the incentives to inflate earnings but also exerts a similar impact on the incentives to meet or beat market expectations. In other words, the disciplining effect dominates the potential concerns of downward price pressure in the real corporate world.

The second alternative earnings management proxy is earnings persistence. As Dechow, Ge, and Schrand (2010) have shown, pretending to be capable of generating "sustainable" earnings is another motivation for a firm to engage in earnings management (in addition to the desire to inflate earnings captured by our accruals variable) because superior business fundamentals may lead to sustainable earnings. By contrast, in the absence of earnings management, earnings will be less stable for all firms except for perhaps the very best group of firms in the economy. Although short selling should not affect firms with truly superior fundamentals, it reduces the incentives for bad firms to mimic good firms by manipulating earnings sustainability. We therefore expect that SSP will reduce earnings persistence.

Models (4) and (5) in Panel B test this effect by regressing earnings (operating income scaled by lagged total assets) or modified Jones's (1991) residual accruals on the interaction between SSP and the lagged dependent variable. The interaction terms are significantly negative for both variables. Therefore, lendable shares reduce both earnings and accrual persistence.

Panel C examines three alternative accrual measures—namely, Jones's (1991) residual accruals, Kothari, Leone, and Wasley's (2005) residual accruals, and Dechow and Dichev's (2002) residual accruals—as well as the probability that firms are involved in earnings misstatements or corporate scandals. Kothari, Leone, and Wasley's (2005) model further controls for firm fundamentals by matching a firm with another from the same country, industry, and year with the closest ROA; and Dechow and Dichev's (2002) further controls for operating performance by regressing results on past, current, and future cash flows. SSP disciplines all these alternative accrual and misbehavior variables.[18] These results, together with the test on market-beating expectations, demonstrate that short selling disciplines managerial incentives to manipulate accruals and apply other forms of earnings management.

We also consider the effect of alternative discipline channels based on the quality of a firm's corporate governance and accounting standards, including the quality of the firm's auditors, the quality of the firm's accounting standards, the quality of the firm's corporate governance (as defined by the ISS index), the transparency of the firm (dispersion of analysts), and press coverage by news agencies. We use the following variables: the ISS corporate governance index (*ISS*), big N auditor (*BigN*), international accounting standard (*IAS*), news coverage (*NewsCoverage*), and analyst dispersion (*Disp*). A higher value for any of these variables typically indicates better governance, except for *Disp*, for which a lower value helps mitigate bad managerial incentives.

All of these variables provide alternative means of disciplining managers or improving the ability of the market to obtain information about them. For example, the quality of governance has been used by Doidge, Karolyi, and Stulz (2007) and represents the standard governance metric based on the by-laws and statutes of the firm. Additionally, transparency—through improved accounting standards, better auditors, or a lower dispersion of their forecasts—improves the awareness of uninformed shareholders. In Panel D, we separately control for these variables because the addition of these alternative controls drastically reduces the size of the sample. The results are qualitatively and quantitatively similar to the main results.

An alternative interpretation of Panel D is that these alternative disciplining variables could be spuriously related to short selling. For instance, large firms may have both more lendable shares and news coverage. Controlling for these alternative variables helps to reduce the impact of spurious correlation.

[18] Our results are also robust to other accrual variables, such as Francis, LaFond, et al. (2005) and Allent, Larson, and Sloan's (2013) residual accruals. To save space, these results are tabulated in Table IA12.

4.4 Earnings management and stock price synchronicity

Finally, we link to Saffi and Sigurdsson (2011) and explore the extent to which short selling can increase the informativeness of the stock price specifically through its impact on earnings management. We follow Morck, Yeung, and Yu (2000) and Jin and Myers (2006) in spirit, and Bris, Goetzmann, and Zhu (2007) in particular, to construct a proxy for firm-specific information based on the idiosyncratic risk of the stock. The measure, downside-minus-upside R^2 (R^2_{DMU}) , is constructed as the difference between a firm's downside R^2 and its upside R^2, where downside (upside) R^2 is estimated by regressing weekly individual stock on weekly positive (negative) local and U.S. market returns.

As indicated by Bris, Goetzmann, and Zhu (2007), short-selling restrictions will reduce firm-specific information and thus price efficiency, especially during the downside of the market (compared with the upside of the market), leading to a higher value of R^2_{DMU}. As a robustness check, we also construct a second proxy, downside-minus-upside nonsynchronicity ($Nonsyn_{DMU}$), where downside (upside) nonsynchronicity is the logarithm of (1–downside (upside) R^2) divided by downside (upside) R^2. Since nonsynchronicity is high when firm-specific information is abundant, a higher value of $Nonsyn_{DMU}$ implies a higher degree of stock price informativeness.

In Models (1) and (3) of Table 8, proxies for stock price informativeness are regressed on *Lendable*, firm-level control variables, and the unreported industry, country, and year fixed effects. The results confirm the finding of Saffi and Sigurdsson (2011) that short selling is in general associated with more stock price informativeness.

In Models (2) and (4), we further interact *Lendable* with its potential degree of disciplining impact based on the finding (reported in Panel C of Table 6) that this impact is the highest among small-cap stocks that have positive $Accrual_{MJones}$. More specifically, we define a dummy variable, SSP_{Impact}, which takes a value of 1 for small-cap stocks with positive $Accrual_{MJones}$. We find that the interaction between *Lendable* and SSP_{Impact} greatly enhances the positive relationship between SSP and stock price informativeness, confirming that greater stock price efficiency can be achieved when the disciplining effect of SSP on earnings management is more pronounced. This analysis completes our analyses regarding the disciplining role of short selling in reducing managers' incentives to engage in earnings management.

These results are important. Thus far, we have shown that SSP reduces earnings management. Table 8 further suggests that short selling increases the informativeness of the stock price by reducing earnings management. This finding is consistent with existing evidence (e.g., Saffi and Sigurdsson 2011) indicating that short selling improves price efficiency. However, the channel is different: price efficiency is enhanced not by improved market conditions but by lower earnings management by firms.

Table 8
Stock price informativeness and earnings management

Price Informativeness=	R^2_{DMU}		$Nonsyn_{DMU}$	
	Model (1)	Model (2)	Model (3)	Model (4)
Lendable	**–0.055**	**–0.053**	**0.383**	**0.371**
	(–13.67)	**(–13.08)**	**(9.37)**	**(9.42)**
$Lendable \times SSP_{Impact}$		**–0.078**		**0.455**
		(–5.23)		**(1.76)**
SSP_{Impact}		0.002		−0.007
		(1.56)		(−0.47)
Size	−0.002	−0.001	0.014	0.014
	(−2.93)	(−2.87)	(2.79)	(2.83)
BM	−0.001	−0.001	0.006	0.006
	(−1.49)	(−1.33)	(1.45)	(1.34)
Leverage	0.003	0.003	−0.024	−0.024
	(1.62)	(1.59)	(−1.37)	(−1.36)
Return	0.005	0.005	−0.055	−0.054
	(7.58)	(7.41)	(−8.02)	(−7.91)
STD	−0.005	−0.005	−0.009	−0.009
	(−4.97)	(−4.95)	(−0.73)	(−0.73)
ADR	−0.003	−0.003	0.029	0.028
	(−1.55)	(−1.52)	(1.77)	(1.74)
MSCI	−0.001	−0.001	0.036	0.036
	(−0.91)	(−0.84)	(3.76)	(3.72)
Analyst	0.000	0.000	−0.001	−0.001
	(0.05)	(0.00)	(−1.24)	(−1.24)
CH	−0.007	−0.006	0.074	0.074
	(−4.28)	(−4.16)	(4.81)	(4.75)
IO	0.013	0.013	−0.126	−0.125
	(11.36)	(11.23)	(−11.09)	(−11.04)
Illiquidity	0.001	0.001	−0.012	−0.013
	(3.08)	(3.13)	(−3.73)	(−3.76)
Fixed Effects	ICY	ICY	ICY	ICY
Obs	59,953	59,953	59,952	59,952
AdjRsq	7.6%	7.6%	5.4%	5.4%

This table presents the results of a panel regression of a firm's stock price informativeness on lendable shares, its potential impact on earnings management, the interaction between lendable shares and its potential impact on earnings management, and firm-level control variables (X), as well as unreported industry, country, and year fixed effects (ICY) for the full sample and different subsamples. The regression model is

$$PriceInformativenss_{i,t} = \alpha + \beta_1 Lendable_{i,t} + \beta_2 Lendable_{i,t} \times SSP_{Impact\,i,t} + \beta_3 SSP_{Impact\,i,t} + \beta_4 X_{i,t} + \varepsilon_{i,t},$$

where $Price\ Informativenss_{i,t}$ refers to two proxies of stock price informativeness, downside-minus-upside R^2 (R^2_{DMU}) in Models (1) and (2), and downside-minus-upside nonsynchronicity ($Nonsyn_{DMU}$) in Models (3) and (4). A higher degree of price informativeness is associated with lower values of R^2_{DMU} and higher values of $Nonsyn_{DMU}$. $SSP_{Impact\,i,t}$ is a dummy variable that takes a value of 1 when the stock is a small firm with $Accrual_{MJones\,i,t}$ greater than zero, and zero otherwise. $X_{i,t}$ includes the same list of firm control variables as before. The construction of these variables is detailed in Appendix A. The t-statistics reported in parentheses are based on standard errors adjusted for heteroscedasticity and firm-level clustering. Obs denotes the number of firm-year observations, and AdjRsq is adjusted R^2. The sample period is from 2002 to 2009.

5. Conclusion

In this paper, we study whether the potential for short selling has a disciplining impact on earnings management incentives. We argue that short selling affects the behavior and incentives of managers because its presence can accelerate the pace at which information is incorporated into the market and thus allows the market to uncover potential earnings management with a higher probability and at a higher speed. Thus, we expect SSP—the maximum potential impact that short selling may have on firm behavior—to significantly reduce firms' incentives to engage in earnings management. Alternatively, firms may simply ignore the short-selling market or manipulate earnings to a greater extent when they are concerned about the downward price pressure that may be associated with potential short selling.

We test these hypotheses by using data on worldwide short selling detailed at the stock level for the period from 2002 to 2009. Our results show a strong negative correlation between SSP and earnings management that is statistically significant and economically relevant. Endogeneity tests based on instrumental variables (ETF ownership) and two experiments (the SHO experiment in the United States and the introduction of short selling into the Hong Kong stock market) inform a causal interpretation of this negative relationship that SSP reduces earnings management. We show that the disciplining effect of short selling applies to various types of earnings management, and our results are robust to the use of alternative proxies for SSP. Moreover, alternative disciplining channels do not absorb the power of short selling.

These results confirm the disciplining hypothesis and offer evidence of the beneficial effects of the short-selling market on the corporate market. In this regard, short selling not only contributes to the efficiency of the information environment of the stock market but also may improve the contracting institutions of the real economy.

References

Admati, A. R., and P. Pfleiderer. 2009. The Wall Street walk and shareholder activism: Exit as a form of voice. *Review of Financial Studies* 22:2645–85.

Aggarwal, R., I. Erel, M. Ferreira, and P. Matos. 2011. Does governance travel around the world? Evidence from institutional investors. *Journal of Financial Economics* 100:154–81.

Aggarwal, R., P. A. C. Saffi, and J. Sturgess. Forthcoming. The role of institutional investors in voting: Evidence from changes in lendable. *Journal of Finance*.

Akbas, F., E. Boehmer, B. Erturk, and S. Sorescu. 2013. Short interest, returns, and fundamentals. Working Paper.

Allent, E., C. Larson, and R. Sloan. 2013. Accrual reversals, earnings and stock returns. *Journal of Accounting Economics* 56:113–29.

Amihud, Y. 2002. Illiquidity and stock returns: Cross-section and time-series effects. *Journal of Financial Markets* 5:31–56.

Aitken, M., A. Frino, M. McCorry, and P. Swan. 1998. Short sales are almost instantaneously bad news: Evidence from the Australian Stock Exchange. *Journal of Finance* 53:2205–23.

Arellano, M., and S. Bond. 1991. Some tests of specification for panel data: Monte Carlo evidence and an application to employment equations. *Review of Economic Studies* 58:277–97.

Asquith, P., and L. Meulbroek. 1995. An empirical investigation of short interest. Working Paper, Massachusetts Institute of Technology.

Barth, M., Landsman, W., and Lang, M. 2008. International accounting standards and accounting quality. *Journal of Accounting Research* 46:467–98.

Boehmer, E., C. Jones, and X. Zhang. 2008. Which shorts are informed? *Journal of Finance* 63:491–527.

Boehmer, E., and J. Wu. 2013. Short selling and the price discovery process. *Review of Financial Studies* 26:287–322.

Bris, A., W. Goetzmann, and N. Zhu. 2007. Efficiency and the bear: Short sales and markets around the world. *Journal of Finance* 62:1029–79.

Burgstahler, D., and I. Dichev. 1997. Earnings management to avoid earnings decreases and losses. *Journal of Accounting and Economics* 24:99–126.

Chang, E. C., J. W. Cheng, and Y. Yu. 2007. Short-sales constraints and price discovery: Evidence from the Hong Kong market. *Journal of Finance* 62:2097–2121.

Christoffersen, S. K., C. Geczy, D. K. Musto, and A. V. Reed. 2007. Vote trading and information aggregation. *Journal of Finance* 62:2897–2929.

Cohen, L., K. Diether, and C. Malloy. 2007. Supply and demand shifts in the shorting market. *Journal of Finance* 62:2061–96.

Conley, T. G., C. B. Hansen, and P. E. Rossi. 2012. Plausibly exogenous. *Review of Economics and Statistics* 94:260–72.

Cumming, D. J., S. A. Johan, and D. Li. 2011. Exchange trading rules and stock market liquidity. *Journal of Financial Economics* 99:651–71.

Das, S., and H., Zhang. 2003. Rounding-up in reported EPS, behavioral thresholds, and earnings management. *Journal of Accounting and Economics* 35:31–50.

DeAngelo, L. 1981. Auditor independence, “low balling,” and disclosure regulation. *Journal of Accounting and Economics* 3:113–27.

Dechow, P., and I. Dichev. 2002. The quality of accruals and earnings: The role of accrual estimation errors. *Accounting Review* 77:35–59.

Dechow, P., W. Ge, and C. Schrand. 2010. Understanding earnings quality: A review of the proxies, their determinants and their consequences. *Journal of Accounting and Economics* 50:344–401.

Dechow, P., R. Sloan, and A. Sweeney. 1995. Detecting earnings management. *Accounting Review* 70:193–225.

DeFond, M., and C. Park. 1997. Smoothing income in anticipation of future earnings. *Journal of Accounting and Economics* 23:115–39.

Degeorge, F., J. Patel, and R. Zeckhauser. 1999. Earnings management to exceed thresholds. *Journal of Business* 72:1–33.

Diamond, D. W., and R. E. Verrecchia. 1987. Constraints on short-selling and asset price adjustment to private information. *Journal of Financial Economics* 18:277–311.

Diether, K. L, K. H. Lee, and I. M. Werner. 2009. It is SHO time! Short-sale price tests and market quality. *Journal of Finance* 64:37–63.

Doidge, C., G. A. Karolyi, and R. M. Stulz. 2004. Why are foreign firms listed in the U.S. worth more? *Journal of Financial Economics* 71:205–38.

Doidge, C., G. A. Karolyi, and R. M. Stulz. 2007. Why do countries matter so much for corporate governance? *Journal of Financial Economics* 86:1–39.

Duffie, D., N. Garleanu, and L. H. Pedersen. 2002. Securities lending, shorting, and pricing. *Journal of Financial Economics* 66:307–39.

Dyck, A., A. Morse, and L. Zingales. 2010. Who blows the whistle on corporate fraud? *Journal of Finance* 65:2133–2255.

Edmans, A. 2009. Blockholder trading, market efficiency, and managerial myopia. *Journal of Finance* 64:2481–2513.

Edmans, A., V. W. Fang, and E. Zur. 2013. The effect of liquidity on governance. *Review of Financial Studies* 26:1443–82.

Edmans, A., and G. Manso. 2011. Governance through trading and intervention: A theory of multiple blockholders. *Review of Financial Studies* 24:2395–2428.

Fang, V., A. Huang, and J. Karpoff. 2014. Short selling and earnings management: A controlled experiment. Working Paper

Faure-Grimaud, A., and D. Gromb. 2004. Public trading and private incentives. *Review of Financial Studies* 17:985–1014.

Ferreira, M., and P. Matos. 2008. The colors of investors' money: The role of institutional investors around the world? *Journal of Financial Economics* 88:499–533.

Francis, J., R. LaFond, P. Olsson, and K. Schipper. 2005. The market pricing of accruals quality. *Journal of Accounting and Economics* 39:295–327.

Francis, J., P. Michas, and S. Seavey. 2013. Does market concentration harm the quality of audited earnings?: Evidence from audit markets in 42 countries. *Contemporary Accounting Research* 30:325–55.

Francis, J., and D. Wang. 2008. The joint effect of investor protection and Big 4 audits on earnings quality around the world. *Contemporary Accounting Research* 25:157–91.

Grullon, G., S. Michenaud, and J. P. Weston. 2012. The real effects of short-selling constraints. Working Paper.

Hirshleifer, D., S. H. Teoh, and J. Yu. 2011. Short arbitrage, return asymmetry and the accrual anomaly. *Review of Financial Studies* 24:2429–61.

Jain, A., P. Jain, T. McInish, and M. McKenzie. 2013. Worldwide reach of short selling regulations. *Journal of Financial Economics* 109:177–97.

Jin, L., and S. Myers. 2006. R-squared around the world: New theory and new tests. *Journal of Financial Economics* 79:257–92.

Jones, J. 1991. Earnings management during import relief investigations. *Journal of Accounting Research* 29:193–228.

Kahn, C., and A. Winton. 1998. Ownership structure, speculation, and shareholder intervention. *Journal of Finance* 53:99–129.

Karolyi, G. A. 2004. The role of American depositary receipts in the development of emerging equity markets. *Review of Economics and Statistics* 86:670–90.

Karolyi, G. A. 2006. The world of cross-listings and cross-listings of the world: Challenging conventional wisdom. *Review of Finance* 10:99–152.

Karolyi, G. A., K. Lee, and M. A. Van Dijk. 2012. Understanding commonality in liquidity around the world. *Journal of Financial Economics* 105:82–112.

Karpoff, J. M., and X. Lou. 2010. Short sellers and financial misconduct. *Journal of Finance* 65:1879–1913.

Kothari, S., A. Leone, and C. Wasley. 2005. Performance matched discretionary accrual measures. *Journal of Accounting and Economics* 39:163–97.

Kolasinski, A. C., A. V. Reed, and M. C. Ringgenberg. 2013. A multiple lender approach to understanding supply and search in the equity lending market. *Journal of Finance* 68:559–95.

La Porta, R., F. Lopez-de-Silanes, and A. Shleifer. 2006. What works in securities laws? *Journal of Finance* 61:1–32.

Leuz, C., D. Nanda, and P. D. Wysocki. 2003. Earnings management and investor protection: An international comparison. *Journal of Financial Economics* 69:505–27.

Li, K., H. Ortiz-Molina, and X. Zhao. 2008. Do voting rights affect institutional investment decisions? Evidence from dual-class firms. *Financial Management* 37:713–45.

Ljungqvist, A., and W. Qian. 2014. How constraining are limits to arbitrage? Working Paper.

Maug, E. 1998. Large shareholders as monitors: Is there a tradeoff between liquidity and control? *Journal of Finance* 53:65–98.

Miller, E. M. 1977. Risk, uncertainty, and divergence of opinion. *Journal of Finance* 32:1151–68.

Morck, R., B. Yeung, and W. Yu. 2000. The information content of stock markets: Why do emerging markets have synchronous stock price movements? *Journal of Financial Economics* 59:215–60.

Morsfield, S., and C. Tan. 2006. Do venture capitalists influence the decision to manage earnings in initial public offerings? *Accounting Review* 81:1119–50.

Nissim, D., and S. Penman. 2001. Ratio analysis and equity valuation: From research to practice. *Review of Accounting Studies* 6:109–54.

Pagano, M., and P. F. Volpin. 2005. The political economy of corporate governance. *American Economic Review* 95:1005–30.

Petersen, M. A. 2009. Estimating standard errors in finance panel data sets: Comparing approaches. *Review of Financial Studies* 22:435–80.

Prado, M. P., P. A. C. Saffi, and J. Sturgess. 2013. Ownership structure, limits to arbitrage and stock returns: Evidence from equity lending markets. Working Paper.

Saffi, P., and K. Sigurdsson. 2011. Price efficiency and short selling. *Review of Financial Studies* 24:821–52.

Senchack, A. J., and L. Starks. 1993. Short-sale restrictions and market reaction to short-interest announcements. *Journal of Financial and Quantitative Analysis* 28:177–94.

Staiger, D., and J. H. Stock. 1997. Instrumental variables regression with weak instruments. *Econometrica* 65:557–86.

Thompson, S. 2010. Simple formulas for standard errors that cluster by both firm and time. *Journal of Financial Economics* 99:1–10.

Watts, R., and J. Zimmerman. 1986. Positive accounting theory. Englewood Cliffs, NJ: Prentice-Hall.

Yu, F. 2008. Analyst coverage and earnings management. *Journal of Financial Economics* 88:245–71.