The Real Effects of Short Selling Restrictions

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This Version: April, 2017

Abstract:

Short selling enhances the price feedback mechanism for corporate investment decisions. Using the staggered introduction and repeal of short selling restrictions around the world, we show that the sensitivity of investment to Tobin's Q is 25% higher when short selling is permitted. Further, short selling strengthens the link between the completion of announced M&A deals and the price reaction on announcement day. The positive impact of short selling on investment-Q sensitivity is higher for firms with low analyst coverage and institutional ownership, mitigated for cross-listed firms, and subsequently results in improved firm performance. The effects do not appear to be driven by a higher *total* amount of information in prices, faster price discovery, improved firm disclosure, or differences in governance. Our results suggest that short selling incentivizes traders to acquire private information new to firm managers, thus enhancing managers' reliance on stock prices as an information signal for investment decisions.

Key Words: Short Selling, Price Feedback Mechanism, Managerial Learning, Real Efficiency

JEL Classification: G14, G15, G31, G39

1. Introduction

In this paper, we examine if and how short selling constraints affect the efficacy of the price feedback mechanism for corporate investment decisions. For a long time, short selling has been considered an important determinant of market and price efficiency by both regulators and finance scholars. For example, Diamond and Verrecchia (1987) show that short selling constraints can reduce informed trading and impact the incorporation of private information into prices. Regulators like the SEC on the other hand argue that "unconstrained short selling can generally undermine confidence in our markets".[1] Despite this longstanding interest and research effort, little is known about the real effects of short selling, particularly for economic decision making and allocative efficiency.

Financial markets can improve the efficiency of managerial decisions by aggregating the information of millions of investors (Hayek 1945; Bond, Edmans, and Goldstein 2012). Imagine, for example, that a firm's stock price falls dramatically after the firm announces plans to open a plant in another country. Management might then conclude that market participants have identified investment risks the managers missed, and subsequently abandon the plan. Importantly, the ability for stock prices to act as a guide for managerial actions depends on the amount of information in prices *not already available or known* to the manager (Edmans, Jayaraman, and Schneemeier 2016). Bond et al. (2012) refer to this concept as "revelatory price efficiency" (RPE). Surprisingly, previous empirical research on the implications of short selling for price efficiency has focused almost exclusively on aspects related to the *total* amount of information in prices (FPE)[2], such as market return co-movement (Bris, Goetzmann, and Zhu 2007), the speed at which prices incorporate public news (Saffi and Sigurdsson 2010), and the deviation of return processes from random walks (Boehmer and Wu 2013). In contrast, we focus on the real effects of short selling constraints regarding RPE and the price feedback effect for corporate investments in this paper.

In theory, there are two main channels through which short selling can affect the amount of information in stock prices previously unknown to managers (RPE), and in turn the price feedback mechanism. On the one hand, the inability to benefit from private information due to short selling constraints can deter market participants from gathering and impounding RPE-relevant

[1] See for example http://www.sec.gov/rules/final/2010/34-61594.

[2] Bond et al. (2012), as well as Bai, Philippon, and Savov (2016) and Edmans, Jayaraman, and Schneemeier (2016) call the extent to which prices reflect all available information "forecasting price efficiency" (FPE).

information into stock prices. In a model of endogenous information acquisition and investment, Nezafat and Wang (2013) show that traders optimally choose to acquire less RPE-relevant information under short selling constraints, compared to the equilibrium without such restrictions. In this case, we would expect short selling constraints to diminish outsiders' incentives to acquire firm-specific information and reduce the efficacy of the price feedback mechanism. At the same time, short selling constraints and restrictions can lead to overvaluation by making it more difficult for prices to reflect existing negative information, as suggested e.g. by Miller (1977), Scheinkman and Xiong (2003), and Hong and Stein (2003).[3] If inflated stock prices lower firms' costs of capital, short selling constraints can directly lead to higher investment, without a feedback and learning effect.[4]

On the other hand, a large theoretical literature on the effects of short selling points to the potentially detrimental effect of strategic "bear raiders". As shown in Goldstein and Guembel (2008), the price feedback mechanism can break down if these strategic bear raiders short-sell and manipulate stock prices to distort firms' investment policies. By making otherwise value-creating investment projects appear unprofitable and thus preventing their implementation, bear raiders may have an incentive to short-sell stocks even when they are uninformed. As suggested by the SEC, this can reduce confidence in market prices. Hence, if managers are aware of these distortions, they may optimally decide not to rely on stock prices to guide their decision making in equilibrium. While reducing the incentives for traders to acquire information, under this scenario we would expect short selling constraints to ultimately enhance the price-feedback mechanism, as they prevent malicious manipulation.

When studying this question, we are faced with two major empirical challenges. First, there exists no obvious, reliable measure or proxy for revelatory price efficiency (RPE), making it difficult to directly examine the effect of short selling on the usefulness of stock prices as a guide for managerial decisions. In our baseline tests, we therefore follow in the footsteps of Chen, Goldstein, and Jiang (2007), Foucault and Frésard (2012), and Edmans et al. (2016), among others, and rely on the sensitivity of investment to Tobin's Q to study the price feedback mechanism.

[3] Several papers argue that the direct effect of short selling constraints on prices is small or negligible (e.g. Diether, Lee, and Werner 2009; Beber and Pagano 2013; and Kaplan, Moskowitz, and Sensoy 2013).
[4] This direct impact of short selling constraints on investment through firms' cost of capital is the focus of Grullon, Michenaud, and Weston (2015).

Specifically, we hypothesize that managers will rely more heavily on the signal contained in stock prices (Tobin's Q) for subsequent investment decisions, if prices contain more information previously unknown or unavailable to them (RPE).

We additionally consider one key investment decision to study the impact of short selling on the price feedback mechanism: mergers & acquisitions (M&A). Luo (2005) shows that the abnormal return of the acquiring firm around an M&A announcement is a strong driver of the likelihood that the transaction is ultimately completed or withdrawn, consistent with the idea of RPE. Hence, if a short selling ban reduces the incentives of traders to gather firm-specific information, or alternatively decreases the likelihood of uninformed bear raids and improves the reliability of stock price signals, we should observe a decrease (increase) in investment-Q sensitivity and the ability of M&A announcement returns to predict deal completion. Since information already possessed by the manager will not help inform managerial decisions, such an effect could not be explained solely by changes in the total amount of information in prices (FPE).

Second, stock prices, firm fundamentals, investments, and short selling by market participants are of course jointly determined. To alleviate potential endogeneity concerns, we follow Bris, Goetzmann, and Zhu (2007) and Saffi and Sigurdsson (2011), among others, and exploit country-level regulation changes and restrictions of short selling obtained from Jain, Jain, McInish, and McKenzie (2013) as a quasi-natural experiment. During the wave of market liberalizations around the world in the 1990s, several countries removed their bans on short selling or naked short selling. Further, following the Financial Crisis of 2007/08, several countries introduced (partial) bans on short selling and repealed these restrictions a few years later. This staggered introduction and repeal of institutional short selling constraints reduces the risk that any single event was correlated with other factors driving the investment-Q sensitivity.

We test our predictions about the price feedback effect of short selling for corporate investments using a large panel of approximately 220,000 firm-year observations from 45 countries over the period from 1994 to 2014. Our key finding is that the legalization of short selling increases investment-Q sensitivity by 27% to 35%, significant at the 1% level. This result is robust to the use of various estimation methodologies, the additional inclusion of country-by-industry-by-year fixed effects to control for possible between-country, between-industry, and across-year differences in investment opportunities, as well as the inclusion of firm fixed effects and alternative

subsample periods and investment proxies. Regarding the sensitivity of M&A deal completion to announcement-day returns, we similarly find a stronger link in countries and months with fewer restrictions to short selling.

These results, supporting the price-feedback channel, are above and beyond the direct impact of short selling constraints on investment, for example through the cost of capital, as we control for this direct mechanism in each of our regressions. In fact, consistent with Grullon et al. (2015), we confirm that the reduction of short selling constraints indeed leads to a decrease in investment in our international sample.

We find qualitatively similar results for the regulation of naked short sales. However, consistent with the notion that more severe restrictions to short selling have a larger impact on the incentives of market participants to acquire information, the economic magnitude of our findings for naked short sales bans is smaller. In line with Grullon et al. (2015), we find no significant differences regarding the enforcement of an uptick rule in countries which allow short selling. Taken together, our results suggest that the reduction of short sale constraints increases the incentives for market participants to acquire and impound firm-specific information into stock prices, hence improving RPE and the price feedback mechanism.

Massa, Zhang, and Zhang (2015) and Fang, Huang and Karpoff (2016) show that the reduction of short selling constraints reduces earnings management and improves disclosure quality. Boehmer and Wu (2013) further provide evidence that short selling increases the efficiency at which prices incorporate public information. Hence, an important potential concern is that the positive effect of short selling on investment-price sensitivity is not due to the feedback and learning effect channel, but instead due to a higher total amount of information in prices (FPE). We address this concern in multiple ways. First, we include various proxies of FPE-type price efficiency previously used in the literature (firm-specific return variation (Chen et al. 2007), proportion of zero return days and FHT trading costs (Edmans et al. 2016), speed of information diffusion (Hou and Moskowitz 2005), and variance ratios (Boehmer and Wu 2013)), as well as their interactions with Tobin's Q and institutional short selling constraints and show that our results remain robust.

Second, we directly test the learning channel hypothesis by introducing an additional layer of interactions with several proxies designed to capture the ex-ante incentives of market

participants to become informed, following the approach used in Foucault and Frésard (2012). Our results show that the positive impact of short selling on investment-Q sensitivity is stronger if a firm's stock returns have lower co-movement with market returns, as well as lower analyst coverage and institutional ownership.

Third, we test how cross-listings impact the effect of short selling on the price feedback mechanism. As shown by Foucault and Frésard (2012), firms with cross-listings in the U.S. exhibit a stronger price feedback effect, since their stocks generally include more information relevant for investment decisions. In this sense managers of cross-listed firms in countries with short selling restrictions have an additional "source of information" abroad and do not need to rely on their local (constrained) stock price signals. Indeed, in a triple-difference setting, we find that the positive effect of short selling legalization on investment-Q sensitivity is almost completely mitigated for firms with cross-listings in the U.S.

Fourth, if introducing short selling increases the amount of information in stock prices previously unknown or unavailable to the decision maker (RPE), managers should be more likely to select value-enhancing projects as they can obtain a more precise signal from the market. This increased RPE should in turn translate into better operating performance down the road. Following the methodology of Durnev (2010), this is exactly what we find focusing on firms' return on assets and sales growth in the years following short sale regulation changes.

We further explore the idea that short selling incentivizes market participants to gather firm-specific information by examining cross-sectional differences in capital market development, governance, and the quality of the information environment across countries. Beasley (1996), Leuz, Nanda, and Wysocki (2003) and Dyck and Zingales (2004) suggest that less developed capital markets, weaker investor protection and lower levels governance standards are associated with a higher probability of financial fraud, lower disclosure quality, and higher private benefits of control. Since short sellers are often able to detect financial misconduct (Karpoff and Lou 2010), this environment may on the one hand attract more traders gathering firm-specific information, hence improving the amount of new information in prices. On the other hand, the threat of short selling might discipline managers to improve firm disclosure quality (Massa et al. 2015), or make them more responsive to the information already contained in prices. An increase in investment-Q sensitivity after a reduction in short selling constraints might therefore simply be a result of

prices reflecting firm fundamentals more accurately, or due to less extraction of private benefits by managers. If this was the main driver, we would expect greater increases in investment-Q sensitivity after the removal of a short selling bans in less developed capital markets with lower governance and disclosure standards. Using various standard proxies for capital market development, governance and disclosure quality we do not find significant differences between countries with low and high standards. In fact, focusing on GDP per capita, we document a significantly larger effect for the removal of short selling restrictions for developed compared to emerging market countries.

This paper contributes to two main strands of the finance literature. First, we add a new perspective to the body of research on the effects of short selling on market and real efficiency. Focusing on a sample of U.S. firms around the Regulation SHO experiment, Grullon et al. (2015) find that firms lower their equity issuance and investments following an increase in short selling and lower stock prices.[5] Massa et al. (2015), as well as Fang et al. (2016) show that short selling reduces earnings management, the likelihood of financial fraud, and improves disclosure quality. Further, Bris, Goetzmann, and Zhu (2007), as well as Saffi and Sigurdsson (2010) demonstrate that short selling improves price efficiency regarding the co-movement with local market returns and the deviation of scaled variance ratios from the level implied by a random walk. Boehmer, Jones, and Zhang (2008) show that short sellers as a group add information to stock prices and Boehmer and Wu (2013) provide evidence that short selling accelerates the incorporation of public information into prices and reduces the post-earnings-announcement drift. Recently, Doidge, Cen, and Schiller (2017) show that short selling increases the speed of firm-specific information diffusion in international supply-chain networks. We primarily contribute to the literature on the real effects of short selling by studying an important feature of financial markets in an international setting: the ability to provide information for corporate decision making. We also extend the results in Grullon et al. (2015) to an international level focusing on a different set of short sale constraints.

Second, a quickly growing literature, starting with Durnev, Morck, and Yeung (2004), Luo (2005), and Jiang et al. (2007) studies the feedback effect of prices for managerial decision making. In this vein, Foucault and Frésard (2012) examine the effect of a firm's decision to cross-list abroad

[5] They do not find evidence for an effect of Reg. SHO on the price feedback channel. However, given that Regulation SHO only affected the imposition of the uptick rule for a relatively short period, this difference in results compared to our findings is not surprising.

on the learning channel, and Foucault and Frésard (2014) and Williams and Xiao (2016) show that managers obtain signals and information for investment decisions from stock prices of peers and supply chain partners. By focusing on the introduction of insider trading enforcement around the world, Edmans et al. (2016) distinguish between the types of information contained in prices and their effect on the learning channel. Bai et al. (2016) further show that both, FPE and RPE, have increased significantly over the last decades. To the best of our knowledge, our paper is the first to examine the real effects of short selling with respect to management decision making and the price feedback effect. Since we generally observe a decrease in short selling constraints around the world over time, our results also help explain the findings in Bai et al. (2016).

The rest of this paper is organized as follows. Section 2 introduces our theoretical framework and derives empirical implications. Section 3 describes our data and sample construction. Section 4 presents the main results regarding the effect of short selling on investment-Q sensitivity and M&A deal completion, as well as a host of robustness tests. Section 5 provides further evidence for the learning channel hypothesis, and Section 6 concludes. All variables definitions as well as additional tests mentioned but not reported in the main text for brevity are provided in the appendix.

2. Theoretical Framework and Hypotheses

The information structure and firm investment problem

To fix ideas and motivate our empirical tests, we present a simple, illustrative model of information contained in stock prices and investment-to-price sensitivity in this section. Since the model largely follows Bai et al. (2016) and Edmans et al. (2016), we refrain from fully fleshing out the derivation of the results and predictions and focus on discussing the implications for our research question.[6]

Consider a public firm with assets in place and a growth opportunity. The manager of the firm chooses investment k, given her information set I_m, to maximize total firm value given by:

$$v(z,k) = (1+z)\left(\bar{k}+k\right) - k - \frac{\gamma}{2\bar{k}}k^2 \qquad (1)$$

[6] For the full set of proofs and derivations, the reader is referred to Bai et al. (2016).

where $\bar{k}$ are the assets in place, z is a productivity shock (normalized to have mean zero), and γ represents an adjustment cost parameter. As in Edmans et al. (2016), the firm's productivity has two components, i.e. $z = z_1 + z_2$. The manager's information set I_m is determined in equilibrium and consists of two parts. First, due to his position running the firm, the manager has perfect information about one part of the firm's value, and receives the private signal $s_m = z_1$ at no cost. Second, the manager can derive information about z_2 from the market price p of a traded security linked to the firm's value. For example, one can imagine that the manager is perfectly informed about internal information such as firm profitability (z_1), but not about external information regarding the prospects and risks in certain industries, markets, or product segments (z_2).

In the model, there are two groups of traders: market participants who can choose to acquire information about the firm and liquidity traders. The market participants may choose to acquire information about the value of z_2 at a cost. If they do, they observe a private, noisy signal $s_i = z_2 + \epsilon_z$, where $\epsilon_z \sim N(0, \sigma_z^2)$. Managers ($z_1$) and speculators ($z_2$) therefore have superior information about different components of the firm value.

At the same time, market participants observe information about z_1, for example in the form of disclosure from the firm and the manager to the market. They hence observe a signal $s_d = z_1 + \epsilon_d$, where $\epsilon_d \sim N(0, \sigma_d^2)$ is orthogonal to ϵ_z. Based on these signals, informed speculators trade with exogenous liquidity traders in the secondary market, allowing the price p to be determined by market clearing. Assuming a linear pricing function and a generic payoff of the market security $F(z, k^*)$, the price follows an expression of the form

$$p = \alpha E[F(z, k^*)|s_i, s_d] + \beta u \tag{2}$$

where u is noise trading demand, and α and β are constants determined by the exact model specification. Finally, the manager's information set is given by $I_m = \{s_m, p\}$, and the manager's optimal investment k^* and the maximized ex-post firm value follow as

$$k^* = \frac{\bar{k}}{\gamma} E[z|I_m] = \frac{\bar{k}}{\gamma} E[z_1 + z_2|s_m, p] = \frac{\bar{k}}{\gamma} z_1 + \frac{\bar{k}}{\gamma} E[z_2|s_m, p] \tag{3}$$

and

$$\frac{v(z, k^*)}{\bar{k}} = 1 + z + \frac{z}{\gamma} E[z|I_m] + \frac{1}{2\gamma} E[z|I_m]^2. \tag{4}$$

The "price feedback effect" is captured by the fact that optimal investment k^* depends on the information contained in security price p, and p in turn is determined by informed market participants trading on their private signals. Therefore, Bond et al. (2012) differentiate between two types of price efficiency: forecasting price efficiency (FPE), and revelatory price efficiency (RPE). Normalizing the price by assets in place $\bar{k}$ as $q = \frac{p}{\bar{k}}$, firm value has a first order linear approximation as $\frac{v}{\bar{k}} \approx 1 + z$, from equation (4) and Bai et al. (2016) therefore define FPE as the 'variance of the predictable component of firm value given q':

$$FPE = Var(E[z|q]). \tag{5}$$

FPE, capturing the total amount of information about future payoffs of the firm contained in prices, has been the focus of many previous papers on the price efficiency implications of short selling (Bris et al. 2007, Saffi and Sigurdsson 2010, Boehmer and Wu 2013). On the other hand, RPE captures the improvement in investment efficiency and real allocations due to information contained in prices managers *do not already have* and can be written as follows:

$$RPE = Var(E[z|s_m, q]) - Var(E[z|s_m]) = Var(E[z_2|s_m, q]) - Var(E[z_2|s_m]) \tag{6}$$

where $Var(E[z|s_m, q]) \equiv AE$ is the aggregate efficiency (*AE*) given the manager's information set.[7] RPE is high and investments are efficient, if the ability of managers to estimate productivity shock z, given both s_m and the information contained in prices, is large compared to the accuracy when she has no information from prices (which capture the information of informed traders).

Empirical implications

Following directly from Bai et al. (2016), we can derive the following two predictions for the implications of the information structure (σ_z^2, σ_d^2) for FPE, AE and RPE, which remain robust across different modeling assumptions for the model (preferences of the traders, trading protocol, specification of $F(z, k^*)$):

1) A decrease in σ_z^2 (traders acquire more private information) increases FPE, RPE, and AE.
2) A decrease in σ_d^2 (firm discloses more precise information) increases FPE, but not RPE and AE.

[7] The expected firm value is a function of the manager's ability to forecast z, since Equation (3) can be written conditionally on the manager's information set as $E[v(z, k^*)] = \bar{k} + \frac{\bar{k}}{2\gamma} Var(E[z|I_m])$.

As traders produce more information (signal s_i becomes more precise) RPE and investment efficiency, FPE and AE all increase since the produced information is new to the manager and increases $Var(E[z_2|s_m, q])$. In a model of endogenous information acquisition and trading, Nezafat and Wang (2013) show that short sale constraints reduce the amount of information market participants optimally acquire in equilibrium. In the context of our model, a similar effect can easily be observed. In equilibrium, the number of traders who decide to become informed is determined by the trade-off between the cost of information acquisition F, and the expected utility from investing, $E[U(x^*\,(z-p))|s_i, s_d]$, where x^* is the optimal demand for the security given the traders information set $I_m = \{s_d, s_i\}$. In the case of a short selling ban ($x \geq 0$), an informed trader with a negative private signal s_i would not be able to benefit from his information. In this case fewer traders would decide to become informed, increasing σ_z^2 and decreasing FPE, AE, and RPE, everything else equal.

On the other hand, most papers in this literature (for example Subrahmanyam and Titman 1999, Foucault and Gehrig 2008, Edmans et al. 2016) assume that the traded security is a claim only on the assets in place z, i.e. $F(z, k^*) = z$. This greatly simplifies the analysis as it shuts down the feedback loop "stock price affects investment affects stock price" and helps avoid solving a complex fixed-point problem. However, it also eliminates the malicious manipulation channel introduced in Goldstein and Guembel (2008). If instead $F(\cdot,\cdot)$ depends on both z and k^*, then large uninformed investors may have an incentive to manipulate the security by short selling it even if they are uninformed, as shown in Goldstein and Guembel (2008). In the context of our model, this would decrease FPE and AE.

Further, Massa et al. (2015) and Fang et al. (2016) show that the threat of short selling disciplines managers and improves disclosure quality. In the context of our model, this is equivalent to a decrease in σ_d^2, i.e. the firm discloses more precise information reflected in prices, which leads to an increase in FPE but not RPE and investment efficiency.

Short selling, investment-Q sensitivity, FPE and RPE

Following a long empirical literature pioneered by Chen et al. (2007), we will rely on the cross-sectional relationship between firm-level investment and Tobin's Q to examine the impact of short selling on real efficiency through the price feedback effect in our main tests. In the context of our model, investment-Q sensitivity can be expressed as

$$\beta_{k,p} = \frac{Cov(k,p)}{Var(p)} = \frac{Cov\left(\frac{\bar{k}}{\gamma}E[z|I_m],p\right)}{Var(p)} = \frac{\frac{\bar{k}}{\gamma}Cov(E[z_2|s_m,p],p)}{Var(p)}. \quad (7)$$

In this framework, short selling constraints and restrictions can affect investment-Q sensitivity in two ways. First, and most important for our research question, the legalization of short selling attracts more informed traders, increasing the precision of the price signal (decrease in σ_z^2) and the covariance term in the numerator of equation (7), everything else equal. This results in a positive effect of short selling on investment-Q sensitivity $\beta_{k,p}$ through the price feedback channel (RPE). Second, on the one hand short selling can destabilize the extent to which prices reflect firm fundamentals due to manipulative bear raiders and malicious, uninformed short selling as suggested by Goldstein and Guembel (2008), hence increasing the $Var(p)$ term in the denominator of (7). On the other hand, short selling can increase the correlation between prices and fundamentals as suggested by Boehmer and Wu (2013) and others, and hence decrease the $Var(p)$ term in (7) as shown by Edmans et al. (2016). The FPE-related effect on short selling is therefore ambiguous.

3. Data and Empirical Methodology

Financial firm-level data and short selling regulation

Our sample construction starts with the universe of all non-U.S. firms covered in the Compustat Global database. For each firm-year, we collect financial and accounting variables that serve as proxies for investment, investment opportunities, firm profitability and financing constraints, including market and book value of equity, total assets, capital expenditure, sales, property, plants, and equipment (PPE) and others for the period from 1994 to 2014 and convert all variables into US Dollar values. All variables and data sources are explained in detail in the appendix. We further exclude financial firms (SIC codes 6000 to 6999) and utilities (SIC codes between 9000 and 9999), since their investment and accounting numbers are largely determined by regulatory requirements. We also exclude firm-years without available price and return data, or fewer than 30 weekly return observations for a firm's common stock listings in the Compustat Daily Returns database, as well as firms with negative sales or negative book or market value of equity. To mitigate the impact of reporting errors and outliers all continuous accounting variables are winsorized at the 1% level.

We further obtain data on the proportion of (foreign) institutional investor ownership from the Factset LionShares dataset, as well as the number of analyst forecasts for a given firm in each year from I/B/E/S International. Our data on firm-level cross listings on a United States stock exchange is gathered from various sources, including BNY Mellon, JP Morgan, Citibank, and NASDAQ. We further construct several proxies for price efficiency relying on daily stock returns from the Compustat Global dataset, explained in detail in Section 3.

Tables 1a and 1b present descriptive statistics, including mean, standard deviation, median and interquartile range, as well as the correlation coefficients for the main firm-level variables of interest and explanatory variables in this paper respectively. Consistent with the numbers documented for example in Foucault and Frésard (2012), the median investment (CapEx(t)/PPE(t-1)) for our sample is 0.15 with an interquartile range of 0.07 to 0.30, and the median Tobin's Q is 1.12 with an interquartile range of 0.89 to 1.57. The median firm has total assets of 201.29 Million USD with an interquartile range of 63.48 Million USD to 695.09 Million USD, a return on assets (ROA) of 3%, book leverage ratio of 20%, and is covered by one analyst. Only a small proportion of the firms in our sample, approximately 7.2%, are cross-listed in the United States. Unsurprisingly, firm size is positively correlated with foreign institutional ownership, analyst coverage, the speed of information diffusion, and cross-listings in the U.S., and negatively correlated with market return co-movement and the deviation of the 5-week variance ratio from 1, as shown in Table 1b.

[Insert Table 1 about here.]

Our main data source for short selling regulation and regulation changes is provided by Jain et al. (2013). Their paper covers the introduction and repeal of bans on short selling, naked short selling, and the enforcement of an uptick rule when short selling is legal of more than 90 countries since the early 1990s, as summarized in Table A1.

After merging these datasets, we are left with a sample of 215,858 firm-year observations (26,596 unique firms) from 45 countries, including 172,000 firm-year observations for our main proxy of firm investments (CapEx/PPE(t-1)). The countries with the largest number of observations in our sample are Japan (44,768), China (19,625), the UK (16,063), India (15,176), Taiwan (13,880), Australia (13,553), and South Korea (10,116). Table 2 presents descriptive statistics by country for the firms in our sample. In most developed countries, short selling and

naked short is permitted throughout our sample period, except for Japan which introduced a naked short selling ban following the financial crisis in 2008. In total, short selling is legal for approximately 98% of the 119,451 firm-year observations from countries with a high GDP per capita (20 countries), and naked short selling is legal for about 67%. In contrast, regarding our sample of firms from low income countries (25 countries), short selling and naked short is legal for about 57% and 23% respectively of our 96,407 firm-year observations.

[Insert Table 2 about here.]

Figure 1 provides an illustration of the evolution of our sample over time. We see that generally both the sample of firm-year observations with and without short selling bans is increasing over time due to expanding coverage by Compustat Global. We further observe changes to short selling regulation, i.e. bans and legalizations of (naked) short selling in several countries with a broad coverage in our sample at different points in time, including Argentina, Chile, Finland, India, Indonesia, Malaysia, South Korea and Thailand, among others. These switches, which are key for our identification, are also visible in the increase in firm-year observations with and simultaneous decrease in observations without short sale bans at different points in time in Figure 1, most notably during and after the financial crisis.

We further obtain data on GDP per capita from the World Bank, and stock market development including the total market capitalization per GDP and per capita, and the number of stock listings per population from the World Federation of Exchanges (WFE). We additionally collect civil and common law origins and the pervasiveness of earnings management from La Porta, Lopez-de-Silanes, and Shleifer (2006), as well as the Anti-Self-Dealing Index from Djankov, La Porta, Lopez-de-Silanes, and Shleifer (2008) for the countries in our sample. Unsurprisingly, stock market development, as proxied by number of stock listings relative to population, is much higher in high-income countries. Market capitalization per GDP is higher on average for the group of low income countries. This average however is dominated by the outlier value and large number of firm-level observations from India. The average values for the Anti-Self-Dealing Index and the Earnings Management Index are relatively similar across the two groups of countries.

M&A transactions and deal completion

Our sample of mergers and acquisitions (M&A) is obtained from the SDC Platinum database and covers the period from 1987 to 2014. We retain all 'disclosed dollar value' M&A deals by non-U.S. companies with a deal value of at least 25 Million U.S. Dollars, which were either officially completed or withdrawn by the acquiring firm, and drop transactions in which the acquirer or target firm were private. We obtain deal specific information regarding percentage sought, deal attitude, target identity, payment details, and anti-takeover provisions from SDC. To construct M&A announcement day excess returns, we collect the daily stock return of the acquiring firm on the M&A announcement day from Compustat Global, and the respective market return on the same day in the country of the acquirer from Datastream. All variables and data sources are explained in detail in the appendix. Table 1a, Panel B summarizes the main variables in our M&A deal sample. In total, the sample includes 10,157 deals undertaken by 4,290 unique acquiring firms with an average deal value of 668.75 Million USD. Approximately 7% of the deals in our sample are withdrawn and not completed, acquirers on average seek to purchase an 82% stake in the target firm and use, on average, a 22% stock payment.

4. Empirical Results

4.1. Short selling and investment-to-price sensitivity

Equation (7) suggests that short selling has a positive effect on investment-to-Q sensitivity by increasing RPE and enhancing the price feedback effect. To test this prediction, we compare the investment-to-Q sensitivity of firms in countries which introduced (repealed) short sale constraints before and after these changes to firms in countries which did not change institutional short selling constraints. We test this prediction we estimate models of the following form,

$$Inv_{i,c,t} = \alpha + \beta_0 * Q_{i,c,t-1} + \beta_1 * Short\ Legal_{c,t-1} + \beta_2 * \left(Q_{i,c,t-1} \times Short\ Legal_{c,t-1}\right) + \delta_1 * X_{i,c,t-1} + \delta_2 * \left(Q_{i,c,t-1} \times X_{i,c,t-1}\right) + \gamma_t + \theta_c + \mu_{s(i)} + \varepsilon_{i,c,t} \quad (8)$$

where $Inv_{i,c,t}$ is a proxy for corporate investment by firm i from country c in year t. In our baseline specifications, we follow Foucault and Frésard (2012) and use capital expenditure (CapEx) normalized by one-year lagged tangible assets (property, plant, and equipment; PPE) as our measure for investment. Variable $Q_{i,c,t-1}$ is the one-period lagged Tobin's Q of firm i, calculated as the sum of total assets (AT) and market capitalization minus book value of equity, scaled by total assets, and $Short\ Legal_{c,t-1}$ is a dummy variable taking the value of one if short selling was

permitted in country c in year $t-1$, as reported in Jain et al. (2013). We include year fixed effects (γ_t), as well as country (θ_c) and industry fixed effects ($\mu_{s(i)}$) in all regression specifications to capture time-invariant country or industry characteristics and changing investment opportunities or constraints over time.[8] Standard errors are adjusted for heteroscedasticity and clustering at the firm level in all specifications.[9]

Our main variable of interest is the interaction of $Q_{i,c,t-1}$ and $Short\ Legal_{c,t-1}$. In line with the empirical predictions in Section 2, a significantly positive coefficient estimate of β_2 would indicate that the reduction of short selling constraints, e.g. the permission of short selling, increases investment-Q sensitivity. This will be the focus of our empirical tests. At the same time, Grullon et al. (2015) suggest that short selling lowers stock prices and equity issuance, and ultimately reduces firm investment. In this sense, β_1 measures the difference-in-difference effect of country-level institutional short selling constraints on firm-level investment, as we include country and time fixed effects in all specifications. Following Grullon et al. (2015), we would expect a significant negative coefficient estimate for β_1 in our model.

Following Chen et al. (2007) and Foucault and Frésard (2012), we additionally include a vector of controls $X_{i,c,t-1}$ with firm level proxies well-known to affect investment, including firm size (measured as the logarithm of total assets, AT), leverage (sum of short and long term debt scaled by total assets) and cash flow (sum of D&A and income before extraordinary items, scaled by lagged total assets), as well as the logarithm of the GDP in country c in the given year.[10] We further extend the models in the previous literature by adding the interaction terms of the vector of controls $X_{i,c,t-1}$ with Q.

Figure 2 presents the results from sequentially estimating the model in equation (8) year by year for the group of firms in countries with and without restrictions on short selling individually. Consistent with our main regression results discussed below, investment-Q sensitivity is higher for countries in which short selling is legal for each year in our sample, except for 2002.

[8] In additional robustness tests, we include firm fixed effects (which then replace the country fixed effects), and country-by-year-by-industry fixed effects to control for changing investment opportunities and financing constraints within-industry, within-country and across-years, as well as time-invariant firm specific characteristics.
[9] We additionally conduct our tests clustering standard errors at the country level and our results remain robust.
[10] We additionally estimate our models including alternative controls for investment constraints, such as the Kaplan-Zingales Index of financial constraints with four components, as well as cash holdings and obtain very similar results.

Investment-Q sensitivity is relatively constant around 0.10 over time for the group of countries without short selling bans. In contrast, we observe more time-series variation for the firms from countries with restrictions to short selling, with an average of approximately 0.01. The large variation over time in part is likely due to the overall smaller sample size. Around the time following the financial crisis in 2008, we see an increase in investment-Q sensitivity for the countries with short selling restrictions and a simultaneous decrease for the countries without such bans in Figure 2. This is consistent with our observation that some countries with high investment-Q sensitivity, such as South Korea, introduced a ban on short selling and 'switched groups' around this time.

[Insert Table 3 about here.]

Table 3 summarizes our main results for the effect of short selling regulation on the investment-Q sensitivity. Column (1) presents OLS estimates of equation (8) including year, country, and industry fixed effects: First, in line with prior literature (for example Baker, Stein, and Wurgler 2003 and Chen et al. 2007), investment is positively and significantly (*t*-stat of 10.09) related to firms' Tobin's Q, for firms both in countries with and without permission of short selling. The regression coefficient for $Q_{i,t-1}$ of 0.0384 in our baseline specification in column (1) is very robust across specifications (between 0.0342 to 0.0511), and implies a 21.09% increase in investment (CapEx/PPE(t-1)) relative to the unconditional mean for a one standard deviation increase in Tobin's Q, after controlling for short selling restrictions. The marginal effect of Tobin's Q ($\hat{\beta}_0 + \hat{\beta}_2 * Mean(Q) = 0.0342 + 0.0513 * 1.12$) on investment is 0.09576, in line with the estimates reported for example in Foucault and Frésard (2012). The coefficient estimates for our control variables are similarly consistent with the previous literature: we find a statistically significant, negative relationship between firm size (i.e. log(AT)) and investment, as well as a negative effect of leverage and country level GDP for investment and a highly significant, positive investment to cash-flow sensitivity.

Second, consistent with Grullon et al. (2015), we find a negative impact of lower short selling constraints on firm investment. The coefficient estimate for $Short\ Legal_{i,t-1}$ is negative (between -0.0770 in column (1) and -0.0379 in column (6)) and significant at the 1% level across all specifications, and captures the difference-in-difference impact of short selling availability on firm investment, conditional on short selling constraints and the interaction with Q. The marginal effect

of short-selling availability on firm investment is negative on average ($\hat{\beta}_1 + \hat{\beta}_2 * Mean(Q) = -0.0770 + 0.0513 * 1.12 = -0.0195$) and implies an economically meaningful decrease of firm investment of up to 13% following the removal of short selling restrictions.

Third, regarding our main research question, we find that the removal of short selling constraints enhances the sensitivity of investment to Q. The coefficient estimate on $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$ is positive (0.0513 in column 1) and statistically significant at the 1% level across all specifications in columns (1) to (6). In other words, when short selling is legal, firms' investments are much more sensitive to stock price signal Q, consistent with the price feedback learning channel. Compared to the coefficient estimate for $Q_{i,t-1}$ of 0.0384 in column (1), this estimate implies that investment is more than twice as sensitive to Tobin's Q in countries when short selling is permitted compared to when it is not. For a one standard deviation increase in Tobin's Q, investment increases significantly for firms both in countries with and without short selling ($\frac{\hat{\beta}_0 * sd(Q)}{mean(Inv)} = 0.2109$). However, the effect is more than twice as large ($\left(\hat{\beta}_0 + \hat{\beta}_2\right) * \frac{sd(Q)}{mean(Inv)} = (0.0384 + 0.0513) * \frac{0.824}{0.15} = 0.5378$) for firms in countries that allow short selling for our baseline model in column (1).

For robustness, we re-estimate the model in equation (8) using various estimation methods, additional controls, and fixed effects. Column (2) of Table 3 presents the results of estimating equation (8) allowing for random effects at the firm level using GLS. In column (3) we estimate a multi-level mixed effects (MLME) model, allowing for both varying intercepts and $Q \times Short$ coefficient estimates at the country level, to account for country-level variation in the effect of short selling on investment-Q sensitivity. The magnitude, direction, and statistical significance remain virtually unchanged in both specifications ($\hat{\beta}_2$ estimate of 0.0431 and 0.0488 respectively) compared to our baseline model.

In columns (4) and (5) we further replace the country and industry fixed effects from our baseline OLS model with year-by-country-by-industry fixed effects and firm fixed effects respectively, to control for any unobservable differences within the same industry, in the same country, across years as well as any time-invariant firm characteristics. Of course, in column (4), the year-by-country-by-industry fixed effects subsume the coefficients for short selling legality and GDP. The estimates for Q, and the interaction of Q and Short Selling however are virtually

unchanged compared to column (1). Similarly, when including firm fixed effects in column (5), the direction and significance of our variables of interest as well as the controls remain unchanged with a slightly smaller coefficient estimate for $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$ at 0.0319. Further, including interactions of our control variables (one-period lagged leverage and cash flow) with Tobin's Q in column (6) does not alter the results. Given the small differences in our results when including firm fixed effects, we use the specification with year, country, and industry fixed effects as our baseline model in most of our analyses to maximize the efficiency of estimates, following Foucault and Frésard (2012).

Robustness tests with alternative proxies and subsample periods

One potential concern might be that our results are due to the specific measure of firm investment used in Table 3. In Table 4a we therefore re-estimate our baseline model from equation (8) using four alternative investment proxies: the change in total assets (ΔAT(t)/AT(t-1)), the change in property, plants, and equipment (ΔPPE(t)/PPE(t-1)), capital expenditure and research and development (R&D) expenses scaled by assets (CapEx+R&D)(t)/AT(t-1)), as well as R&D scaled by PPE (R&D(t)/PPE(t-1)). Across all four specifications summarized in columns (1) to (4), the results confirm our findings from Table 3. Considering these four proxies, investment is positively related to Tobin's Q and increases by 35%, 34%, 25%, and 75% respectively for a one standard deviation increase in Q, relative to the unconditional sample mean in countries with and without short selling. As in Table 3, we find that the permission of short selling has a negative direct effect on investment in line with Grullon et al. (2015). Most importantly, the coefficient estimate for $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$ is positive and significant at the 1% level (except for column 3), indicating that firm investment is between 35% to 60% more sensitive to Tobin's Q in countries which permit short selling, compared to countries with bans on short selling. Including R&D measures significantly reduces our sample size, which is potentially why we do not find a statistically significant effect of short selling legality on investment-Q sensitivity in column (3).

[Insert Table 4 about here.]

As shown in Appendix A1, many of the changes to short selling restrictions in our sample occurred around the time of the financial crisis in 2007/08 and were reverse in the following years. At the same time, it is possible that the financial crisis affected firms' financing constraints and investment opportunities, as well as the informativeness of stock prices and the price feedback

mechanism. In Table 4b we therefore present results for our baseline model using different subsamples and estimation periods. Our results in columns (1) to (3) of Table 4b suggest that the impact of short selling legality on the investment-Q sensitivity is stable over time and robust for the subsample periods before 2005, from 2005 to 2009, and from 2010 to 2014. In each specification (1) to (3), we obtain a positive coefficient estimate for $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$ (with *t*-stats between 3.18 to 7.88) indicative of an approximately 60% to 90% higher sensitivity of investment to Q for countries which allow short selling and a marginal effect of Q on investment in line with the results from our baseline models in Table 3 (between 0.09124 to 0.1035). Further, to alleviate the concern about distortions during the financial crisis, we exclude the years 2007 and 2008 from our sample in column (4) and re-estimate equation (8). The results are virtually unchanged compared to the baseline specification.

4.2. FPE vs. RPE, and investment-Q sensitivity

The main goal of our paper is to study the impact of short selling restrictions on the price feedback mechanism of stock prices for corporate decision making. As outlined in section 2, while short selling is expected to have a positive impact on the 'learning channel' through revelatory price efficiency (RPE), short selling can also have a positive effect on investment-Q sensitivity by increasing the total amount of information incorporated in prices (FPE). Previous literature (e.g. Boehmer and Wu 2013) has shown that short selling can enhance the speed at which information is incorporated into prices, and improve disclosure quality and reduce fraudulent reporting (Massa et al. 2015). One might therefore be concerned that our results in Tables 3 and 4 are not due to the impact of short selling on the price feedback mechanism (RPE), but instead the result of a higher FPE following the removal of short selling constraints. If prices reflect fundamental information already available to managers more accurately and managers make investment decisions based on these fundamentals, we might observe an increase in investment-Q sensitivity when unrelated to the 'price feedback channel'.

To alleviate these concerns, we control directly for a host of proxies designed to capture forecasting price efficiency (FPE) and the total amount of information contained in prices, which have previously been used in the literature to study the impact of short selling on price efficiency. First, following Morck, Yueung, and Yu (2000), Fernandes and Ferreira (2009), and Chen et al. (2007), we calculate the firm-specific return variation, $(1 - \mathrm{R2})$, using the weekly returns from

Wednesday to Wednesday for each firm $R_{i,t}$ and the respective local market return $R_{m,t}$. We separately estimate the following market model for each firm-year in our sample: $R_{i,t} = \alpha + \sum_{k=0}^{4} \beta_k R_{m,t-k} + \varepsilon_{i,t}$. We include four weekly lags in addition to the contemporaneous market return to account for the possibility of slow information dissemination. The measure for firm-specific return variation is then obtained as one minus the R^2 of the above market regression. As suggested by Roll (1988), the R^2 from the above equation captures the percentage of variation in a firm's weekly stock returns explained by market returns. The lower this value, and hence the higher (1- R^2), the more stock price variation is due to firm-specific information. Following this idea, a long literature has relied on (1- R^2) as a measure for the informativeness and information contained in prices.

Second, we calculate the 'speed of information diffusion' in the tradition of Hou and Moskowitz (2005) as the ratio of the R^2 from the above regression including only the contemporaneous market return, divided by the R^2 of the same regression model including all four lagged returns, i.e. $Speed = \frac{R^2_{\beta_k=0,\forall k>0}}{R^2}$. The larger this measure, the less additional variation in stock returns is explained by including lagged market returns, and hence the higher the speed of information diffusion. Variations of this measure have been used to capture price and information efficiency e.g. in Hou (2007), Griffin, Kelly, and Nardari (2010), Saffi and Sigurdsson (2010), and Boehmer and Wu (2013).

Third, we add the FHT measure from Fong, Holden, and Trzcinka (2017), as well as the proportion of zero returns per number of trading days as measures of trading costs and liquidity for each firm-year observation in our sample, since liquidity is a key determinant of price efficiency, following Edmans et al. (2016). Last, as explained for example in Griffin et al. (2010), under the null hypothesis of a random walk with uncorrelated increments, variance ratios should equal one at all lags. Variance ratios different from one hence indicate serial correlation and a departure from a random walk and informationally efficient markets. We therefore calculate the absolute value of the five-week variance ratio minus one, i.e. (|VR5-1|), for each firm-year in our sample as another measure of price efficiency commonly used in the literature (e.g. Saffi and Sigurdsson 2010) in the context of short selling restrictions.

[Insert Table 5 about here.]

We augment our baseline specification by including these five proxies and their interactions with Tobin's Q, as well as firm fixed effects *and* year-by-country-by-industry fixed effects. If the above proxies indeed capture forecasting price efficiency (FPE), as previously suggested in the literature, and the primary effect of short selling on investment-Q sensitivity is through this FPE channel, the direct inclusion of these controls and interactions should largely soak up the impact of short selling on investment-Q sensitivity. Although this is obviously not a perfect test for distinguishing RPE and FPE, we should at least observe a significant drop in the effect of short selling on the sensitivity of investment to Q. The results summarized in Table 5 do not provide evidence for this alternative explanation. The coefficient estimate for $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$ remains positive and statistically significant at the 1% level across all specifications in columns (1) to (5), and the magnitude of the estimate is virtually unchanged compared to the fixed effects specification in Table 3. In fact, none of our five FPE proxies is significantly correlated with firm investment or investment-Q sensitivity, suggesting that the total amount of information contained in prices and price efficiency are unrelated to the learning channel in our sample.[11]

4.3. Types of short selling constraints and investment-Q sensitivity

So far our analysis has focused on the impact of short selling bans in countries around the world as provided in Jain et al. (2013). However, in the wake of the financial crisis some countries moved to introduce and later repeal bans on *naked* short selling, or enforced the use of an uptick rule.[12] These institutional restrictions to short selling are less severe compared to an outright ban on short selling. Nonetheless they pose a constraint on the ability to short sell stocks and should therefore have a similar impact on the incentives of market participants to acquire private information and the price feedback mechanism for firm investment decisions, albeit to a lesser extent. Our theoretical framework from section 2 therefore suggests that we should find a smaller effect of naked short selling on the learning channel. Further, relying on the Regulation SHO experiment[13] as an exogenous shock to the enforcement of the uptick rule in the U.S. stock market, Grullon et al. (2015) do not find evidence supporting a relationship between the uptick rule and

[11] The only exception is the interaction of Q and the 'speed of information diffusion', which is negatively associated with short selling, significant at the 10% level.
[12] The uptick rule is a trading restriction stating that a short sale transaction can only be entered at a price that is higher than the previous trade, i.e. on an 'uptick'.
[13] During the Regulation SHO experiment, the uptick rule was suspended for 1,000 randomly selected stocks in the Russell 3000 index for two years from 2004 to 2006.

the price feedback effect. To test these additional hypotheses, we re-estimate the baseline model from equation (8) using the interaction of Tobin's Q with the legality of naked short selling and the enforcement of the uptick rule respectively as our main variables of interest. Since the uptick rule can only be enforced if short selling is generally permitted, we condition our sample to countries which allow short selling in our uptick rule tests.

[Insert Table 6 about here.]

Table 6 reports the results. Consistent with the notion that short selling reduces equity issuance and investment (Grullon et al. 2015), and in line with our findings in Tables 3 and 4, column (2) suggests a negative effect of naked short selling on investment. We obtain a coefficient estimate of -0.0280 for naked short selling, statistically significant at the 1% level. Compared to short selling in general (reported in column 1), this estimate is much smaller in magnitude, supporting our conjecture that less restrictive short selling constraints have a smaller direct effect on firm investment. More importantly, we find a similar result regarding our main variable of interest, the effect of (naked) short selling on investment-Q sensitivity. As shown in column (2), the coefficient for $Q_{i,c,t-1} \times NakedShort\ Legal_{c,t-1}$ is 0.0225, significant at the 1% level, indicating a 29% increase in investment-Q sensitivity in countries permitting naked short sales compared to countries with bans on naked short selling. The economic magnitude of this effect is about half the size of general short selling permission and bans, supporting our conjecture that a less severe short selling constraints has a smaller impact on the price feedback mechanism.

In columns (3) and (4) we consider only countries and firm-years in which short selling was generally permitted, and study the marginal effect of naked short selling and the use of an uptick rule. Consistent with the results in Grullon et al. (2015), we do not find any evidence that naked short selling or the enforcement on an uptick rule affect the price feedback effect, if short selling is generally permitted. We obtain a positive coefficient estimate for $Q_{i,c,t-1} \times NakedShort\ Legal_{c,t-1}$, and a negative estimate for $Q_{i,c,t-1} \times UptickRule_{c,t-1}$, suggesting that constraints to short selling are negatively related to the sensitivity of investment to price, however, the estimates are not significant at conventional levels.

4.4. Short selling and the completion of announced M&A deals

When a CEO announces the plan to acquire another firm in a corporate takeover, it is likely because she believes this M&A transaction will create value for her shareholders. If, however the market reacts to the M&A announcement by trading down the firm's stock, and the manager in the following decides to withdraw the deal, it can be interpreted as 'learning' from the market price (Luo 2005). The information signal revealed in the price reaction of the stock in this context is likely new and previously unknown to the manager, since she would not have announced an M&A deal in the first place if it was expected to reduce firm value. The completion of M&A deals is therefore an appropriate setting to study the impact of short selling on RPE and the feedback mechanism.

Studying M&A transactions as an important and material firm investment decision therefore helps to alleviate potential concerns that short selling primarily affects investment-Q sensitivity through price efficiency in general or firms' disclosure quality. If a manager had private information that a planned investment, i.e. the acquisition of another firm, is destroying firm value, she would not go forward and announce the plan to the public. In our M&A tests, we focus specifically on the announcement day stock return in excess of the respective market return to further minimize the importance of total price efficiency (FPE) as opposed to RPE. To the extent that the speed at which prices incorporate information varies across firms and impacts the efficacy of the feedback mechanism, we are biasing our tests against finding significant results.

Our regressions specifications follow the basic framework of Luo (2005), augmented with an indicator variable capturing whether short selling was permitted or not at the time of the M&A announcement. We estimate models of the following form,

$$Completed_{i,c,t} = \alpha + \beta_1\left(ExcRet_{i,c,t} \times ShortLegal_{c,t}\right) + \beta_2\left(ExcRet_{i,c,t} \times ShortNotLegal_{c,t}\right) + \delta_1 X_{i,c,t} + \gamma_t + \theta_c + \mu_{s(i)} + \varepsilon_{i,c,t} \quad (9)$$

where $Completed_{i,c,t}$ is an indicator variable that takes the value of one if the M&A deal announced by acquirer firm *i* on day *t* was successfully completed, and zero otherwise. $ExcRet_{i,c,t}$ is the stock return of the acquiring firm on announcement day minus the respective local market return in country *c*, and $ShortLegal_{c,t}$ and $ShortNotLegal_{c,t}$ are indicators taking the value of one if short selling is legal (not legal) in country *c* on the day of announcement *t*. We further add a vector of firm and deal characteristics $X_{i,c,t}$, including firm size, deal value, percentage of shares sought, payment details, target characteristics, deal attitude, and anti-takeover provisions, as well

as monthly time fixed effects, and country and industry fixed effects for the acquirer and target firm. To avoid an incidental variables problem and issues with interactions in non-linear models, we estimate equation (9) as a linear probability model in each specifications.

[Insert Table 7 about here.]

Table 7 reports our findings. In column (1) we first confirm the findings of Luo (2005) for our sample of international acquirers disregarding the effect of short selling.[14] The excess return of the acquiring firm on announcement day is positively associated with the subsequent completion of the M&A deal. The coefficient estimate of 0.335 is statistically significant at the 1% level (we cluster standard errors at the firm level) and indicates an approximately 1.5% increase in the probability that the deal will be completed for a one standard deviation increase in excess return. In line with previous literature, we also obtain a significantly negative relationship between M&A deal completion and the ratio of deal value to firm value of the acquirer, the percentage of shares sought, the percentage of stocks used as payment, and a hostile deal attitude from the target firm.

In column (2), we focus on our main research question and estimate equation (9), differentiating between countries with and without restrictions to short selling. The results show that the sensitivity of M&A deal competition to the stock price reaction on announcement day is significantly higher in countries that allow short selling. The coefficient estimate for $ExcRet_{i,c,t} \times ShortLegal_{c,t}$ of 0.409 is statistically significant at the 1% level and indicates a 2% increase in the probability of M&A deal completion for a one standard deviation increase in announcement day excess return. In contrast, the interaction term $ExcRet_{i,c,t} \times ShortNotLegal_{c,t}$ is slightly negative and insignificant. The difference between the two interaction terms is statistically significant at the 5% level (F-Statistic of 5.91). In column (3) we additionally include a multitude of controls for anti-takeover provisions, including the presence of a white knight, asset lockup, or poison pill. The coefficient on our main variable of interest, the interaction of excess return and short selling, remains positive and statistically significant at the 1% level. Similarly, the difference between firms from countries which permit short selling and countries which do not remains large and statistically significant. In column (3), we additionally test the impact of naked short selling

[14] The sample in our paper is about 10 times larger (10,000 deal observations compared to 1,000) than the sample used in Luo (2005) and covers only M&A deal announcements where the acquiring firm is incorporated or listed outside the United States.

on the sensitivity of M&A deal completion on announcement day excess returns. Consistent with our results in Table 6, we do not find a significant difference. M&A deal completion is positively, significantly related to the announcement day excess returns of the acquiring firms. However, the coefficient estimate is approximately similar for firms from both groups of countries, independent of naked short selling restrictions.

Taken together, our results from the investment-Q sensitivity regressions and the M&A deal competition regressions suggest that short selling has a positive impact on the feedback effect from stock prices to firm investment decisions. This price feedback mechanism does not appear to be due to higher overall price efficiency. As shown in Table A3 in the appendix, consistent with the learning channel, the link between M&A deal completion and excess return on announcement day is similarly stronger for stocks that contain more private information (i.e. are cross-listed) and have higher liquidity. In the next section, we further explore this learning channel by examining cross sectional differences in the ex-ante incentives of market participants to get informed, the profitability of firm investments, cross-listings, and the country-level institutional environment.

5. The Price Feedback Channel

5.1. Ex-ante incentives to acquire information

The results so far show that the sensitivity of investment to stock price is higher for firms in countries that permit short selling than for firms in countries that do not. Our theoretical framework in Section 2 and equation (7) suggest that this effect is – at least in part – driven by the fact that the possibility of short selling provides an incentive for market participants to produce private information that otherwise would not be available to managers, i.e. revelatory price efficiency (RPE). As this information is impounded into stock prices through trading, they become a more reliable source of information for managers, resulting in a higher sensitivity of investments to stock prices.

If this learning channel is indeed driving our results, the availability of short selling should have a larger effect on firms with ex-ante higher incentives for market participants to acquire private information. Testing this proposition is challenging as there is no well-established measure for information acquisition incentives. We therefore take three different approaches to exploit cross-sectional variation between firms. First, we rely on a proxy shown in the literature to

encourage the production of RPE-relevant information new to managers. Following Fernandes and Ferreira (2009), we use firm specific stock return variation (1-R2) as a proxy for the amount of information contained in a firm's stock prices. Roll (1988) suggests that informed trades based on firm specific information increase the idiosyncratic risk of a stock. To the extent that at least some of this information is new to managers, we expect (1-R2) to be correlated with market participants' incentive for information acquisition. The impact of short selling legalization on investment-Q sensitivity should therefore be increasing in firm specific return variation.

Second, we follow the "crowding out hypothesis" of Easley, O'Hara, and Paperman (1998): The presence of sell-side analysts producing publicly available firm reports and forecasts reduces the returns market participants can expect from acquiring and trading on private information. Intuitively, all else equal, this effect should lower the incentives of speculators to collect private, firm-specific information. Indeed, Easley et al. (1998) show that the PIN measure, a widely-used proxy for the likelihood of privately informed trading, is negatively associated with analyst coverage. Hence, we would expect that the impact of short selling on investment-Q sensitivity is inversely related to firms' analyst coverage, i.e. the crowding out of private information acquisition by investors due to analyst information production. Following the same idea, we use the proportion of shares held by foreign institutional investors from the FactSet LionShares database as an additional measure for the ex-ante incentives of investors to acquire information. We hypothesize that a larger proportion of shares held by sophisticated, foreign institutional investors reduces the returns market participants can expect from getting informed, since many of these profit opportunities will be competed away quickly by the institutional investors.

Third, we consider firms with a cross-listing on a U.S. stock exchange. Managers of cross-listed firms have an additional "information source" when making investment decisions as they can rely on the RPE-relevant information revealed in the cross-listed security, even if the firm's stock in the home market is short selling restricted. Indeed, Foucault and Frésard (2012) show that cross-listing on a stock exchange in the U.S. increases the sensitivity of investment to Tobin's Q due to an enhanced price feedback effect. If this is the case, we would expect the introduction of short selling to have a smaller effect on cross-listed firms, as they already benefit from a strong learning channel through their securities listed abroad.

We test these predictions by augmenting our baseline model from equation (8) with an additional layer of interactions as follows,

$$Inv_{i,c,t} = \alpha + \beta_0 * Q_{i,c,t-1} + \beta_1 * Short\ Legal_{c,t-1} + \beta_2 * ProxyInfAcq_{i,c,t-1} + \beta_3 * \left(Q_{i,c,t-1} \times Short\ Legal_{c,t-1}\right) + \beta_4 * \left(Q_{i,c,t-1} \times ProxyInfAcq_{i,c,t-1}\right) + \beta_5 * \left(Q_{i,c,t-1} \times Short\ Legal_{c,t-1} \times ProxyInfAcq_{i,c,t-1}\right) \delta_1 * X_{i,c,t-1} + \delta_2 * \left(Q_{i,c,t-1} \times X_{i,c,t-1}\right) + \gamma_t + \theta_c + \mu_{s(i)} + \varepsilon_{i,c,t} \quad (10)$$

where $ProxyInfAcq_{i,c,t-1}$ is our proxy for the ex-ante incentives of market participants to acquire information (i.e. (1-R2), analyst coverage, foreign institutional ownership, cross-listing indicator), and all other variables are defined as before. Our main variable of interest is the triple interaction term $Q_{i,c,t-1} \times Short\ Legal_{c,t-1} \times ProxyInfAcq_{i,c,t-1}$, capturing how the impact of short selling on investment-Q sensitivity is related to information acquisition incentives.

[Insert Table 8 about here.]

Table 8 reports the results. In line with our baseline results in Tables 3 and 4, the coefficient for $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$ remains positive and statistically significant across all specifications in columns (1) to (3), even after introducing the additional interactions terms and control variables. Investment-Q sensitivity in column (1) is significantly positively related to firm specific return variation (1-R2), confirming the results of Chen et al. (2007) for our sample of international firms. More importantly, we find that the effect of short selling on investment-Q sensitivity is increasing in firm specific return variation, suggesting that the impact of short selling is larger when firms attract more private information collection. The coefficient estimate for $Q_{i,c,t-1} \times Short\ Legal_{c,t-1} \times (1 - R2)_{i,c,t-1}$ of 0.0342 is significant at the 5% level, indicating that a one standard deviation increase in $(1 - R2)_{i,c,t-1}$ results in an approximately 25% increase in the impact of short selling legality on investment-Q sensitivity compared to the coefficient estimate for $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$.

Next, consistent with the "crowding out hypothesis", column (2) shows a negative effect of analyst coverage on investment-Q sensitivity (significant at the 1% level), confirming the finding reported in Chen et al. (2007). Importantly, in line with our conjecture, we further find that the impact of short selling availability on investment-Q sensitivity is decreasing in analyst coverage, significant at the 1% level. The coefficient estimate for $Q_{i,c,t-1} \times Short\ Legal_{c,t-1} \times$

$Analysts_{i,c,t-1}$ of -0.00296 implies that for every additional analyst, the impact of short selling legality on investment-Q sensitivity decreases by 4%. The results or foreign institutional ownership presented in column (3) confirm this finding. Again, we find that $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$ is decreasing in the incentives of market participants to get informed, i.e. the proportion of shares held by foreign institutional investors. The coefficient estimate of the triple interaction term is large and significant at the 1% level.

In column (4), we consider the differential effect of short selling availability on investment-Q sensitivity for cross-listed firms compared to all other firms in our sample. $Cross\ Listed_{i,c,t-1}$ is a dummy variable indicating that firm *i* from country *c* was cross-listed in period *t-1*. We find two main results. First, our results confirm the finding of Foucault and Frésard (2012): all else equal, the sensitivity of investment to Tobin's Q is significantly larger for cross-listed firms. The coefficient estimate for $Q_{i,c,t-1} \times Cross\ Listed_{i,c,t-1}$ of 0.0257 implies that investment-Q sensitivity is about 50% higher for cross-listed firms compared to the coefficient estimate for $Q_{i,c,t-1}$ of 0.0491. Second, in line with our hypothesis, we find that the positive effect of short selling availability on investment-Q sensitivity is almost completely mitigated for firm with a cross listing. The coefficient estimate for $Q_{i,c,t-1} \times Short\ Legal_{c,t-1} \times Cross\ Listed_{i,c,t-1}$ is negative (-0.0607) and statistically significant at the 1% level. Importantly, the magnitude of this estimate is very close in absolute terms to the coefficient estimate for $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$ of 0.0516, suggesting that short selling does not attract more private information acquisition and strengthen the feedback mechanism for cross-listed firms, as these market participants could already trade on their private information using the securities cross-listed abroad.

Taken together, the results in this section indicate that the relation between short selling restrictions and investment-to-price sensitivity varies across firms in the cross-section in the way predicted by the price feedback mechanism: Firms with characteristics that should encourage the acquisition of RPE-related information are associated with a relatively higher investment-to-price sensitivity when short selling is permitted in the firm's country of origin.

5.2. Short selling and operating performance

If the price feedback effect for management decision making is indeed driving our results, we would expect managers with a stronger price feedback channel, i.e. when short selling is

permitted, to be able to select better investment projects for their firm. This improved investment efficiency in turn should in the following translate into improved operating performance. We would therefore expect ex-post firm performance to be positively related to the change in investment-Q sensitivity when short selling is permitted. Since we have too few observations to estimate the coefficient for $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$ for each firm separately, we test this proposition using a two-step approach following the methodology of Durnev (2010) and Foucault and Frésard (2012). In the first step, we re-estimate our baseline model from equation (8) without the interaction term of Tobin's Q and short selling and collect the residuals from this regression for each firm-year. As explained in Foucault and Frésard (2012), everything else equal, positive (negative) residuals indicate firm-year observations for which investments are more (less) related to the corresponding stock price. We then define a dummy variable $Positive_{i,c,t-1}$ taking the value of one, if the regression residual of firm *i* in year *t-1* is positive, and negative otherwise. In the second step, we regress company performance on $Positive_{i,c,t-1}$ and a set of controls: short selling legality, Tobin's Q, Return on Assets (ROA), sales growth, logarithm of asset size, leverage, cash flow, and logarithm of GDP per capita, all lagged by one period. We also include year and firm fixed effects in every specification. If the increase in investment-Q sensitivity is driven by a price feedback effect and results in managers choosing more profitable projects, we would expect the coefficient estimate on $Positive_{i,c,t-1}$ to be positive.

Table 9 reports the results. Firm operating performance is measured by return on assets (ROA) in columns (1) and (2), and (5) and (6) and by sales growth in specifications (3) and (4), and (7) and (8). In each regression, we skip one year between the observation of $Positive_{i,c,t-1}$ and our measure of firm performance, i.e. in column (1) we regress ROA in year *t+1* on *Positive* in year *t-1,* to avoid confounding effects between our dependent variable and several controls due to common denominators.[15] We also use the average of the following three years for ROA and sales growth as the dependent variable in columns (5) to (8) in Panel B.

[Insert Table 9 about here.]

Focusing on ROA in column (1), the coefficient on $Positive_{i,c,t-1}$ is 0.00238, significant at the 1% level, which amounts to an approximately 8% increase in ROA compared to the

[15] In untabulated tests we confirm that the magnitude and statistical significance of our findings is similar or even stronger if we do not skip a year and consider firm operating performance in year *t* as the dependent variable instead.

unconditional sample median. The result remains virtually unchanged when we control for short selling legality in column (2). We observe a similar increase in performance when sales growth is considered in columns (3) and (4), as well as (7) and (8): independent of whether the short selling variable is included or not, the coefficient estimate on $Positive_{i,c,t-1}$ is positive and statistically significant. For example, companies that have a positive residual in our first-stage regression experience an increase in sales growth of 0.65 percentage points in the following year (column 4), and 0.53 percentage points in the following three years, which is approximately equivalent to a 5% and 4% increase respectively, compared to the unconditional sample mean. The results support our conjecture that the feedback channel from stock price signals enhances the allocative efficiency of management decisions, consistent with our theoretical framework in Section 2.

5.3. Market development, information environment, and short selling

Our last set of tests further explores the idea of the stock price feedback effect for management decision making by ruling out alternative explanations revolving around country-level differences in stock market development and information environment. One alternative explanation for our findings is that the threat of short selling improves disclosure quality and reduces the likelihood of financial misconduct, as argued by Massa et al. (2015) and Fang et al. (2015). Further, Leuz, Nanda, and Wysocki (2003) and Dyck and Zingales (2004) provide evidence that countries with less developed financial markets and lower governance standards have a lower quality of financial disclosure, as well as higher extraction of private benefits of control. The threat of short selling might hence also discipline managers to choose profitable investment projects instead of extracting private benefits. In both cases, an increase in investment-Q sensitivity might be due to a higher forecasting price efficiency (FPE), and unrelated to the management learning channel. If this is indeed driving our results, we would expect the impact of short selling on the sensitivity of investment to stock price to be larger for firms from less developed countries with weaker governance and less developed stock markets.

On the other hand, Karpoff and Lou (2010) show that short sellers have a strong ability to detect and exploit financial misconduct for profitable trading opportunities. If short selling is not available, it might therefore deter market participants from acquiring information and potentially becoming short sellers, resulting in less private information production and lower RPE.

We test these propositions by collecting various proxies for market development and quality of information environment widely used in the literature. Following Foucault and Frésard (2012) we collect GDP per capita in USD from the World Bank, and the number of stock listings and the total market capitalization per population from the World Federation of Exchanges (WFE) as proxies for market development. Further, we add country legal origin (civil or common law) and the Anti-Self-Dealing Index from Djankov et al. (2008) to capture the potential extraction of private benefits by firm managers, as well as the Earnings Management Pervasiveness Index as a proxy for disclosure quality on the country level. For each proxy, we define a dummy variable $High_{c,t-1}$, which takes the value of one if the value is above the median in country *c* in year *t-1*, indicating high market development, governance and disclosure standards, and zero otherwise[16], and a dummy variable $Low_{c,t-1}$ taking the value of one if it is below the median. Then, we re-estimate the baseline model in equation (8) by interacting $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$ with *High* and *Low*. Table 10 reports the results.

[Insert Table 10 about here.]

In line with the findings in Tables 3 and 4, investment-Q sensitivity is positively related to short selling, irrespective of whether the firm is listed in a country from the *High* or *Low* group. Both, the triple interaction coefficients on $Q_{i,c,t-1} \times Short\ Legal_{c,t-1} \times High_{c,t-1}$, and $Q_{i,c,t-1} \times Short\ Legal_{c,t-1} \times Low_{c,t-1}$ are positive and highly statistically significant across all specifications (1) to (6). More importantly, we do not find evidence that our results for investment-Q sensitivity are driven by alternative explanations revolving around disclosure quality, extraction of private benefits by managers or governance standards. The coefficient estimates on $Q_{i,c,t-1} \times Short\ Legal_{c,t-1}$ are very similar in magnitude for both the *High* and *Low* groups across columns (2) to (6), and statistically indistinguishable at standard confidence levels (see the F -tests at the bottom of Table 10). In fact, considering the estimates shown in column (1), we find that the effect of short selling on investment-Q sensitivity is larger for firms from the *High* GDP per capita group of countries (estimate of 0.0554), compared to the *Low* GDP per capita group (0.0339). The difference is statistically significant at the 1% level.

[16] Legal origin, the anti-self-dealing index and earnings management pervasiveness are time invariant and we define our dummy variables accordingly.

6. Conclusion

In this paper, we explore the role of short selling regulation and restrictions for the price feedback mechanism for corporate decision making. We introduce a simple theoretical model following Bai et al. (2016) and Edmans et al. (2016), suggesting that the availability of short selling can incentivize market participants to acquire private information. If managers do not have perfect information about firm investment prospects, the information signal contained in prices due to informed trading can help inform management investment decisions and result in higher allocative efficiency. On the other hand, if short selling enables uninformed 'bear raiders' to reduce the reliability of price signals and harm price informativeness overall, short selling can have a negative effect on the efficacy of the price feedback mechanism and detrimental 'real' effects.

We combine data on institutional short selling bans and restrictions from Jain et al. (2013) with a large sample of firms from 45 countries over the period from 1994 to 2014, and test these hypotheses by estimating the effect of short selling legality on the sensitivity of firm investment to Tobin's Q. We find that the availability of short selling is associated with an increase in investment-Q sensitivity by 27% to 35%, significant at the 1% level. Our finding is robust to the use of various estimation methodologies, the inclusion of country-by-industry-by-year fixed effects, as well as the inclusion of firm fixed effects and alternative subsample periods and investment proxies.

To alleviate concerns that our findings are a result of overall higher price efficiency when short selling is permitted and hence unrelated to the management learning channel, we control for a wide range of commonly used proxies of price informativeness and forecasting price efficiency (FPE) and show that our results remain unaffected. We further consider one particularly important investment decision specifically, corporate takeover (M&A) announcements, and show that the sensitivity of M&A deal completion to the market reaction on announcement day is significantly higher when short selling is permitted. Further, the impact of short selling on investment-Q sensitivity is significantly larger when the incentives of market participants to acquire firm specific private information and trade on it are higher, for example, for example when the firm has low analyst coverage, institutional ownership, and is not cross-listed on an exchange in the United States. Moreover, firms that benefit from the increase in investment-Q sensitivity due to short selling exhibit better operational performance down the road. We also show that alternative

explanations, such as the level of market development, governance standards, and disclosure quality between the countries in our sample cannot explain our main result.

Our findings provide a novel perspective on the importance of short selling regulation for managerial decision making in particular, and the impact of financial market frictions for 'real' allocative efficiency in general. By affecting the efficacy of the price feedback mechanism for corporate investment decisions, financial market and short selling regulation can have potentially large implications for firm investment and profitability beyond the direct impact on firm's ability to raise financing and the disciplining effect on managers previously shown in the literature.

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Figure 1: Firm-year Observations by Year and Short-Selling Restrictions.

This figure presents the number of observations in our sample by year and short selling restrictions. The black bars (grey bars) represent firm-year observations for which short selling was (not) permitted in the given year.

Figure 2: Investment-Q Sensitivity by Year and Short-Selling Legality.

In this figure, we estimate the sensitivity of investment to Tobin's Q separately, year by year, for the group of firms listed in countries which permit short selling (solid black line), and the group of firms listed in countries which do not (dashed grey line). The dependent variable in each regression is investment, defined as capital expenditures divided by lagged property, plant, and equipment (PPE) and Q is Tobin's Q as defined in the appendix. All regressions include country fixed effects and industry fixed effects based on two-digit SIC codes, as well as similar control variables and interactions of controls with Tobin's Q as our baseline model in equation (8). Due to insufficient sample size, we do not report the estimates for 1994 and 1995. The y-axis represents the coefficient estimate for Tobin's Q in each year.

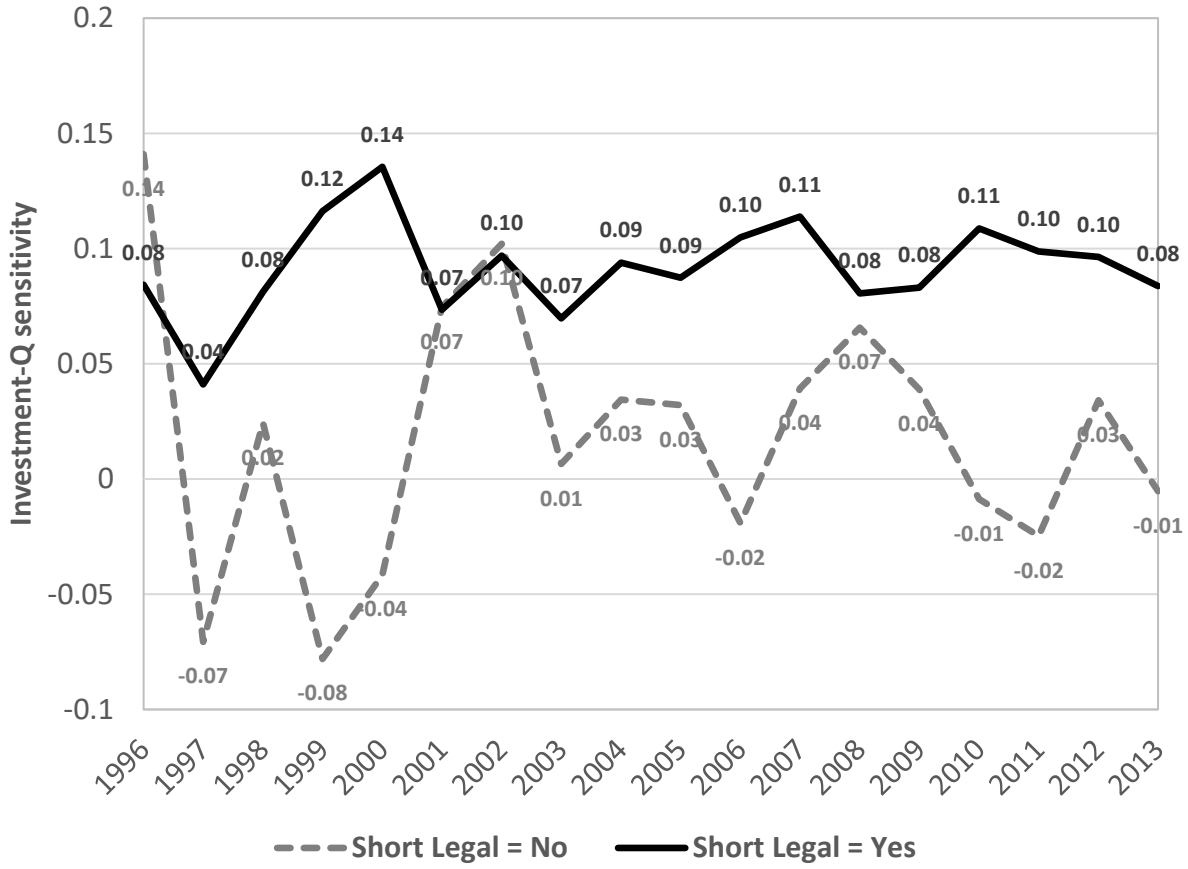

Table 1a: Firm-Level Summary Statistics.

This table reports the number of observations *N*, the mean, and standard deviation, as well as the 5%, 25%, 50%, 75%, 95% percentiles of the main variables used in the analysis. Panel A presents summary statistics for the main variables used in our investment-Q sensitivity tests, Panel B reports the summary statistics for the main variables from our M&A deal completion sample. All variables are defined in detail in the appendix. The sample period is from 1994 to 2014 in Panel A, and from 1987 to 2014 in Panel B.

Panel A: Investment-Q Sensitivity Sample

Variables	N	Mean	SD	p5	p25	p50	p75	p95
Investment:								
CapEx(t)/PPE(t-1)	172635	0.275	0.441	0.01	0.07	0.15	0.30	0.90
ΔAT(t)/AT(t-1)	185790	0.094	0.299	-0.26	-0.06	0.05	0.18	0.58
ΔPPE(t)/PPE(t-1)	184894	0.125	0.549	-0.34	-0.08	0.03	0.18	0.79
(CapEx+R&D)(t)/AT(t-1)	63940	0.084	0.079	0.01	0.03	0.06	0.11	0.24
R&D(t)/PPE(t-1)	66658	0.615	2.137	0.00	0.01	0.06	0.21	3.05
Independent Variables:								
Tobin's Q	207709	1.387	0.824	0.63	0.89	1.12	1.57	3.15
Total Assets (AT) (MM USD)	215581	1621.360	6649.086	9.87	63.48	201.29	695.09	6057.12
Mkt. Capitalization (MM USD)	212933	884.972	2797.114	5.39	31.70	116.36	443.24	3997.45
Book Leverage	214660	0.227	0.193	0.00	0.05	0.20	0.35	0.58
ROA	197258	-0.002	0.164	-0.26	0.00	0.03	0.06	0.14
Sales Growth	183472	0.124	0.475	-0.36	-0.07	0.06	0.22	0.71
Cash Flow (CF) (MM USD)	179699	0.056	0.129	-0.16	0.02	0.06	0.11	0.22
Inst. Own. Foreign (%)	109806	0.046	0.075	0.000	0.002	0.013	0.059	0.196
Analyst Coverage	188398	3.768	6.384	0	0	1	5	18
(1-R2)	215585	0.752	0.158	0.44	0.66	0.78	0.88	0.95
Speed of Information Diffusion	215585	0.567	0.309	0.02	0.31	0.64	0.84	0.96
FHT Trading Costs	214551	0.035	0.192	0.00	0.00	0.01	0.02	0.12
|VR5 -1|	215585	0.494	0.450	0.04	0.19	0.40	0.64	1.34
Cross Listed (y/n)	215858	0.072	0.259	0	0	0	0	1

Panel B: M&A Completion Sample

Variables	N	Mean	SD	p5	p25	p50	p75	p95
Deal Completed (y/n)	10157	0.928	0.258	0.00	1.00	1.00	1.00	1.00
Excess Return [0,0]	10157	0.006	0.044	-0.04	-0.01	0.00	0.02	0.08
Short Legal	10042	0.902	0.298	0.00	1.00	1.00	1.00	1.00
Naked Short Legal	10042	0.610	0.488	0.00	0.00	1.00	1.00	1.00
Deal Value (MM USD)	10157	668.75	3116.328	29.10	54.80	129.43	395.00	2287.11
MCap Acquirer (MM USD)	10033	391031.10	1396478	178.14	1676.44	8822.76	67902.94	2134649.00
Deal Value / Mkt. Cap. Acq.	10033	0.146	0.352	0.00	0.00	0.02	0.11	0.71
% sought	10033	82.398	26.894	25.00	60.00	100.00	100.00	100.00
% stock payment	10157	21.999	40.001	0.00	0.00	0.00	0.00	100.00
Hostile (y/n)	10157	0.014	0.118	0.00	0.00	0.00	0.00	0.00
White Knight (y/n)	10157	0.002	0.043	0.00	0.00	0.00	0.00	0.00
Block Purchase (y/n)	10157	0.007	0.085	0.00	0.00	0.00	0.00	0.00
Challenged Deal (y/n)	10157	0.030	0.171	0.00	0.00	0.00	0.00	0.00
Unsolicited (y/n)	10157	0.023	0.149	0.00	0.00	0.00	0.00	0.00
Asset Lockup (y/n)	9142	0.002	0.044	0.00	0.00	0.00	0.00	0.00
Poison Pill (y/n)	9142	0.001	0.035	0.00	0.00	0.00	0.00	0.00

Table 1b: Pearson Correlation Coefficients.

This table reports Pearson correlation coefficients of the main explanatory variables used in the following tests. The sample period is from 1994 to 2014. All variables are defined in the appendix.

		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
(1)	Tobin's Q	1.000												
(2)	Total Assets (AT) (MM USD)	-0.030	1.000											
(3)	Mkt. Capitalization (MM USD)	0.130	0.790	1.000										
(4)	Book Leverage	-0.220	0.080	0.010	1.000									
(5)	ROA	0.030	0.050	0.110	-0.120	1.000								
(6)	Cash Flow (CF) (MM USD)	0.130	0.060	0.150	-0.140	0.880	1.000							
(7)	Inst. Own. Foreign (%)	0.100	0.280	0.390	-0.010	0.100	0.140	1.000						
(8)	Analyst Coverage	0.130	0.480	0.600	0.010	0.140	0.190	0.510	1.000					
(9)	(1-R2)	0.080	-0.240	-0.260	-0.060	-0.130	-0.120	-0.200	-0.270	1.000				
(10)	Speed of Information Diffusion	-0.020	0.160	0.200	0.050	0.170	0.150	0.160	0.220	-0.600	1.000			
(11)	FHT Trading Costs	-0.020	-0.040	-0.050	-0.010	-0.160	-0.130	-0.040	-0.080	0.110	-0.140	1.000		
(12)	|VR5-1|	0.100	-0.050	-0.060	0.010	-0.030	-0.030	-0.100	-0.090	0.090	-0.070	0.040	1.000	
(13)	Cross Listed (y/n)	0.010	0.380	0.430	0.040	0.040	0.050	0.370	0.390	-0.200	0.150	-0.030	-0.050	1.000

Table 2: Country-Level Summary Statistics.

This table presents summary statistics for the firm-year observations in our sample by country of origin. Panel A presents summary statistics defined at the country-year level for high income countries with a per capita GDP in USD above the median value, Panel B reports the same summary statistics for low income countries in our sample. *No. Obs.* is the number of firm-year observations for each country, *short legal* and *naked legal* are the average number of firm-year during which (naked) short selling is permitted. *GPD/Cap* and *Pop (MM)* are the mean per capita GDP in USD, and country population in Millions from the World Bank. *Mkt. Cap/GDP* and *Stock Listings/Pop* are the mean total market capitalization per GDP and mean number of stock listings per population, both from the World Federation of Exchanges (WFE). *Civil Law* is a dummy capturing the legal origin of the country obtained from Djankov et al. (2008), *ASD* is the Anti-Self-Dealing Index from Djankov et al. (2008), defined at the country level, and *EMGT* is the Earnings Management Pervasiveness Index from La Porta et al. (2006). The sample period is 1994 to 2014.

Panel A: High Income Countries

Country	No. Obs.	Short Legal (y/n)	Naked Legal (y/n)	GDP / Cap (USD)	Pop (MM)	Mkt. Cap / GDP	Stock Listings / Pop	Civil Law (y/n)	ASD	EMGT
Australia	13553	0.93	0.11	46783	21.4	22.33	0.82	0	0.76	4.80
Belgium	1228	1.00	1.00	37222	10.6	6.18	0.16	1	0.54	19.50
Cyprus	341	0.00	0.00	27916	1.1	0.24	1.03			
Denmark	1376	1.00	1.00	48759	5.4	2.60	0.42	1	0.46	16.00
Finland	1543	0.91	0.91	39353	5.3	5.91	0.25	1	0.46	12.00
France	8138	1.00	1.00	34091	63.0	33.16	0.12	1	0.38	13.50
Germany	8002	1.00	1.00	35747	82.0	36.63	0.09	1	0.28	21.50
Hong Kong	1662	1.00	0.00	29605	6.8	49.57	1.60	0	0.96	19.50
Ireland	541	0.55	0.55	44565	4.2	2.20	0.14	0	0.79	5.10
Israel	1886	1.00	0.00	28642	7.5	5.68	0.72	0	0.73	
Italy	2973	0.86	0.86	30904	58.2	23.59	0.05	1	0.42	24.80
Japan	44768	1.00	0.63	37548	127.0	119.42	0.27	0	0.50	20.50
Luxembourg	265	1.00	1.00	91697	0.5	0.75	0.74			
Netherlands	2155	1.00	0.90	38398	16.2	15.86	0.14	1	0.20	16.50
New Zealand	1202	1.00	1.00	27458	4.2	1.51	0.38	0	0.95	
Norway	1746	1.00	1.00	72620	4.7	2.46	0.42	1	0.42	5.80
Singapore	5568	1.00	0.00	38682	4.7	10.25	0.93	0	1.00	21.60
Sweden	3984	1.00	1.00	45669	9.2	9.59	0.28	1	0.33	6.80
Switzerland	2457	1.00	0.87	60117	7.5	15.65	0.34	1	0.27	22.00
UK	16063	1.00	1.00	35463	60.8	70.82	0.16	0	0.95	7.00
Total	119451	0.98	0.67	39031	71.0	64.53	0.35	0.28	0.58	15.05

Panel B: Low Income Countries

Country	N. Obs.	Short Legal (y/n)	Naked Legal (y/n)	GDP / Cap (USD)	Pop (MM)	Mkt. Cap / GDP	Stock Listings / Pop	Civil Law (y/n)	ASD	EMGT
Argentina	568	0.89	0.00	9549	39.1	10.49	0.05	1	0.34	
Brazil	833	1.00	0.00	10381	195.0	111.01	0.02	1	0.27	
Chile	962	0.93	0.00	10975	16.8	18.79	0.14	1	0.63	
China	19625	0.00	0.00	3769	1320.0	677.60	0.01	1	0.76	
Colombia	191	0.00	0.00	5750	45.7	26.32	0.02	1	0.57	
Czech Rep.	133	1.00	1.00	11978	10.3	2.31	0.06	1	0.33	
Greece	2476	0.84	0.00	22795	11.1	4.88	0.26	1	0.22	28.30
India	15176	0.78	0.00	1197	1190.0	1752.66	0.06			
Indonesia	2493	0.86	0.86	2300	230.0	93.95	0.02	1	0.65	18.30
Malaysia	9815	0.55	0.00	6805	26.2	38.80	0.34	0	0.95	14.80
Mexico	832	1.00	0.00	7912	112.0	36.25	0.01	1	0.17	
Pakistan	1786	1.00	0.07	1034	171.0	43.28	0.04			
Peru	469	0.00	0.00	4755	28.9	13.94	0.07	1	0.45	
Philippines	1128	0.91	0.00	1839	88.6	55.52	0.03	0	0.22	8.80
Poland	2833	0.95	0.95	11170	38.1			1	0.29	
Portugal	614	1.00	1.00	17857	10.4	4.66	0.12	1	0.44	25.10
Russia	722	0.95	0.95	12560	143.0	81.28	0.03	1	0.44	
South Africa	2684	1.00	0.00	5457	48.0	100.11	0.09	0	0.81	
South Korea	10116	0.84	0.84	19962	48.8	36.45	0.32	0	0.47	
Spain	1729	1.00	0.00	23764	43.1	42.76	0.58	1	0.37	18.60
Sri Lanka	952	0.00	0.00	2484	20.4	5.79	0.13	0	0.39	
Taiwan	13880	0.39	0.39	18315	23.0	37.20	0.49	1	0.56	22.50
Thailand	4328	0.85	0.00	3811	65.1	46.20	0.08	0	0.81	18.30
Turkey	1961	1.00	1.00	8433	69.9	22.12	0.04	1	0.43	
Venezuela	101	0.00	0.00	7792	26.6	2.30	0.05	1	0.09	
Total	96407	0.57	0.23	8930	489.8	439.39	0.18	0.52	0.52	7.36
Total	**215858**	**0.83**	**0.48**	**25584**	**258.0**	**241.83**	**0.29**	**0.42**	**0.60**	**16.47**

Table 3: Impact of Short Selling Legality on Investment-Q Sensitivity.

This table reports the OLS estimation results for our baseline model in equation (8) on the link between short selling constraints and investment-Q sensitivity. The dependent variable in all specifications is investment, defined as capital expenditures divided by lagged property, plant, and equipment (PPE). *Short Legal* is a dummy variable taking the value of one if short selling is permitted in the given year and *Q* is Tobin's Q as defined in the appendix. *Log(AT)* is the logarithm of a firm's total book assets, *Leverage* is the book leverage defined as short term and long term debt over total assets, cash flow (*CF*) is the sum of D&A plus income before extraordinary items, and *log(GDP)* is the logarithm of the per capita GDP in USD. All variables and data sources are explained in detail in the appendix. Each regression includes an intercept not reported for brevity. The industry fixed effects are based on two-digit SIC codes. We winsorize all continuous covariates at the 1% level. *t*-statistics provided in parentheses are computed based on standard errors clustered at the firm level in all specifications, except for column 3 where we cluster standard errors on the country level due to the MLME model specification. ∗, ∗∗ and ∗∗∗ indicate statistical significance at the 10%, 5% and 1% level, respectively.

	Dependent Variable: CapEx (t) / PPE (t-1)					
	OLS	RE	MLME	OLS	Firm FE	Firm FE
	(1)	(2)	(3)	(4)	(5)	(6)
Q (t-1)	0.0384***	0.0511***	0.0380***	0.0342***	0.0462***	0.0486***
	(10.09)	(13.75)	(3.13)	(7.34)	(10.62)	(8.45)
Short Legal (t-1)	-0.0770***	-0.0529***	-0.0686***		-0.0390***	-0.0379***
	(-9.50)	(-7.29)	(-2.98)		(-4.72)	(-4.65)
Q (t-1) * Short Legal (t-1)	**0.0513*** **	**0.0431*** **	**0.0488*** **	**0.0577*** **	**0.0319*** **	**0.0312*** **
	(10.11)	**(8.84)**	**(3.47)**	**(9.42)**	**(5.73)**	**(5.68)**
log(AT) (t-1)	-0.0206***	-0.0367***	-0.0251***	-0.0193***	-0.0566***	-0.0564***
	(-18.23)	(-23.55)	(-5.02)	(-15.46)	(-11.59)	(-11.56)
Leverage (t-1)	-0.128***	-0.259***	-0.171***	-0.138***	-0.278***	-0.316***
	(-13.62)	(-24.02)	(-6.73)	(-13.91)	(-17.56)	(-11.42)
CF (t-1)	0.104***	0.226***	0.121***	0.0899***	0.279***	0.411***
	(4.88)	(11.20)	(2.80)	(3.83)	(11.60)	(11.65)
log(GDP) (t-1)	0.0200***	0.0177***	0.0195		0.0267***	0.0274***
	(3.13)	(9.38)	(1.51)		(3.60)	(3.69)
Q (t-1) * Leverage (t-1)						0.0318*
						(1.71)
Q (t-1) * CF (t-1)						-0.0766***
						(-4.05)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	No	Yes	Yes
Industry FE	Yes	No	No	No	No	No
Country FE	Yes	No	Yes	No	No	No
Year-by-Country-by-Industry FE	No	No	No	Yes	No	No
Firm FE	No	No	No	No	Yes	Yes
Clustered SE	Firm	Firm	Country	Firm	Firm	Firm
Number of firm-years	130403	133829	133829	122924	131298	131298
R^2	0.138			0.243	0.479	0.479
Adj. R^2	0.137			0.156	0.397	0.397

Table 4a: Robustness Tests using Alternative Measures of Investment.

In this table we report robustness tests for our baseline model in equation (8) considering several alternative proxies for firm investment. The dependent variable in column 1 is the change in total assets, in column 2 we use the change in tangible assets (PPE), column 3 uses capital expenditure and research and development (R&D) expenses, and column 4 uses R&D expenses as a measure for firm investment. All variables are defined at the firm-year level and explained in detail in the appendix. Similar to Table 3, *Short Legal* is a dummy variable taking the value of one if short selling is permitted in the given year and *Q* is Tobin's Q as defined in the appendix. Each regression includes an intercept and similar additional firm-level controls and interactions as Table 3, i.e. the logarithm of a firm's total book assets (*log AT*), book leverage (*Leverage*), cash flow (*CF*), interactions of these variables with Tobin's Q and the logarithm of country GDP, all lagged by one year. All continuous covariates are winsorized at the 1% level. *t*-statistics provided in parentheses are computed based on standard errors clustered at the firm level in all specifications. ∗, ∗∗ and ∗∗∗ indicate statistical significance at the 10%, 5% and 1% level, respectively.

	Dependent Variable: Investment			
	ΔAT(t)/ AT(t-1)	ΔPPE(t)/ PPE(t-1)	(CapEx+R&D)(t)/ AT(t-1)	R&D(t)/ PPE(t-1)
	(1)	(2)	(3)	(4)
Q (t-1)	0.0410***	0.0516***	0.0253***	0.561***
	(12.55)	(9.01)	(7.00)	(7.41)
Short Legal (t-1)	-0.0154***	-0.0351***	-0.00771	-0.404***
	(-2.82)	(-3.46)	(-1.42)	(-4.45)
Q (t-1) * Short Legal (t-1)	**0.0246*** **	**0.0327*** **	**0.00435**	**0.205*** **
	(7.47)	**(5.71)**	**(1.23)**	**(3.26)**
log(AT) (t-1)	-0.00527***	-0.0120***	-0.00249***	-0.120***
	(-10.01)	(-12.30)	(-6.40)	(-13.40)
Leverage (t-1)	-0.174***	-0.155***	-0.0431***	0.927***
	(-14.12)	(-7.10)	(-5.13)	(4.66)
CF (t-1)	0.543***	0.578***	0.0893***	-2.061***
	(27.78)	(15.52)	(7.41)	(-5.15)
log(GDP) (t-1)	-0.0212***	-0.0462***	-0.0119***	0.305***
	(-5.09)	(-6.29)	(-3.57)	(4.12)
Q (t-1) * Leverage (t-1)	0.107***	0.0584***	0.0156**	-1.761***
	(10.66)	(3.40)	(2.14)	(-9.91)
Q (t-1) * CF (t-1)	-0.128***	-0.152***	-0.0565***	-1.250***
	(-12.49)	(-7.98)	(-10.73)	(-6.41)
Year FE	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes
Number of firm-years	133162	132677	47018	48189
R^2	0.142	0.052	0.261	0.318
Adj. R^2	0.141	0.051	0.259	0.316

Table 4b: Robustness Tests using Subsample Periods.

In this table we report robustness tests for our baseline model in equation (8) considering several subsample periods. The dependent variable in all specifications is investment, defined as capital expenditures divided by lagged property, plant, and equipment (PPE). *Short Legal* is a dummy variable taking the value of one if short selling is permitted in the given year. *Q* is Tobin's Q as defined in the appendix. In column 1 we include firm-year observations from 1994 to 2004, and in columns 2 and 3 we consider the subsample periods from 2005 to 2009, and 2010 to 2014, respectively. In column 4 we use the full sample of observations excluding the years of the Financial Crisis in 2007 and 2008. All variables are defined at the firm-year level and explained in detail in the appendix. Each regression includes an intercept and similar additional firm-level controls and interactions as Table 3, i.e. the logarithm of total book assets (log AT), book leverage (Leverage), cash flow (CF), interactions of these variables with Tobin's Q and the logarithm of country GDP, all lagged by one year. All continuous covariates are winsorized at the 1% level. *t*-statistics provided in parentheses are computed based on standard errors clustered at the firm level in all specifications. ∗, ∗∗ and ∗∗∗ indicate statistical significance at the 10%, 5% and 1% level, respectively.

	Dependent Variable: CapEx (t) / PPE (t-1)			
	Subsample:			
	Before 2005	2005 to 2009	2010 to 2014	Excl. Crisis
	(1)	(2)	(3)	(4)
Q (t-1)	0.0526***	0.0467***	0.0395***	0.0522***
	(4.52)	(6.05)	(5.42)	(9.69)
Short Legal (t-1)	-0.0133	-0.0624***	-0.0711***	-0.0643***
	(-0.64)	(-5.15)	(-5.03)	(-7.66)
Q (t-1) * Short Legal (t-1)	**0.0345*** **	**0.0545*** **	**0.0571*** **	**0.0424*** **
	(3.18)	**(7.26)**	**(7.41)**	**(7.88)**
log(AT) (t-1)	-0.0162***	-0.0238***	-0.0216***	-0.0201***
	(-10.26)	(-13.27)	(-13.71)	(-17.69)
Leverage (t-1)	-0.132***	-0.159***	-0.118***	-0.123***
	(-3.79)	(-4.59)	(-3.93)	(-5.78)
CF (t-1)	0.364***	0.400***	0.271***	0.344***
	(5.88)	(6.75)	(5.64)	(9.88)
log(GDP) (t-1)	0.0175	-0.0747***	-0.00743	0.0166**
	(1.01)	(-5.10)	(-0.28)	(2.55)
Q (t-1) * Leverage (t-1)	0.00235	0.0318	-0.0199	-0.00894
	(0.08)	(1.24)	(-0.87)	(-0.54)
Q (t-1) * CF (t-1)	-0.104***	-0.127***	-0.121***	-0.125***
	(-2.63)	(-4.56)	(-4.89)	(-6.64)
Year FE	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes
Number of firm-years	33412	41913	51992	110266
R^2	0.161	0.161	0.126	0.134
Adj. R^2	0.158	0.159	0.124	0.133

Table 5: FPE vs. RPE and the Price Feedback Effect for Firm Investment.

This table reports regression estimates of our baseline model from equation (8) on the link between short selling and investment-Q sensitivity, controlling for the total amount of information in prices and forecasting price efficiency (FPE). The dependent variable in all specifications is investment, defined as capital expenditures divided by lagged property, plant, and equipment (PPE). *Short Legal* is a dummy variable taking the value of one if short selling is permitted in the given year. *Q* is Tobin's Q as defined in the appendix. Following Chen et al. (2007), *(1-R2)* is the firm specific return variation as a proxy of total information contained in prices, *Speed* is the speed of information diffusion measure from Hou and Moskowitz (2005), *FHT* is the trading costs and frictions measure from Fong et al. (2017), *Prop Zero* is the proportion of zero return days per number of trading days in a given year following Saffi and Sigurdsson (2010), and *(|VR5-1|)* is the absolute value of the five-week variance ratio minus one, following Griffin et al. (2011). All variables are defined at the firm-year level and explained in detail in the appendix. Each regression includes an intercept and similar additional firm-level controls and interactions as the previous tables not tabulated for brevity, i.e. the logarithm of total book assets (log AT), book leverage (Leverage), cash flow (CF), interactions of these variables with Tobin's Q and the logarithm of country GDP, all lagged by one year. We additionally include firm fixed effects and country-by-year fixed effects. The sample period is from 1994 to 2014, all continuous covariates are winsorized at the 1% level. *t*-statistics provided in parentheses are computed based on standard errors clustered at the firm level in all specifications. ∗, ∗∗ and ∗∗∗ indicate statistical significance at the 10%, 5% and 1% level, respectively.

	Dependent Variable: CapEx (t) / PPE (t-1)				
	(1)	(2)	(3)	(4)	(5)
Q (t-1)	0.0477***	0.0549***	0.0453***	0.0429***	0.0423***
	(3.38)	(6.70)	(6.78)	(6.09)	(5.55)
Q (t-1) * Short Legal (t-1)	**0.0298*** **	**0.0287*** **	**0.0295*** **	**0.0286*** **	**0.0312*** **
	(4.54)	**(4.39)**	**(4.48)**	**(4.30)**	**(4.59)**
Q (t-1) * (1-R2) (t-1)	-0.00201				
	(-0.13)				
Q (t-1) * Speed (t-1)		-0.0146*			
		(-1.89)			
Q (t-1) * FHT (t-1)			0.0377		
			(0.94)		
Q (t-1) * Prop Zero (t-1)				0.0244	
				(1.18)	
Q (t-1) * (|VR 5 -1|) (t-1)					0.00447
					(0.98)
(1-R2) (t-1)	0.00280	0.000682	0.000308	0.000348	0.000247
	(0.14)	(0.08)	(0.04)	(0.04)	(0.03)
Speed (t-1)	-0.00768	0.0121	-0.00766	-0.00764	-0.00762
	(-1.60)	(1.16)	(-1.59)	(-1.59)	(-1.59)
FHT (t-1)	0.00641	0.00661	-0.0234	0.00695	0.00652
	(0.85)	(0.89)	(-0.76)	(0.95)	(0.87)
Prop Zero (t-1)	-0.00678	-0.00504	-0.00930	-0.0350	-0.00690
	(-0.38)	(-0.28)	(-0.52)	(-1.23)	(-0.39)
|VR 5-1| (t-1)	0.00414	0.00400	0.00407	0.00405	-0.00260
	(1.47)	(1.43)	(1.45)	(1.44)	(-0.42)
Firm Controls & Interactions	Yes	Yes	Yes	Yes	Yes
Country-by-Year FE	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes
Number of firm-years	127742	127742	127742	127742	127742
R^2	0.489	0.489	0.489	0.489	0.489
Adj. R^2	0.403	0.404	0.403	0.403	0.403

Table 6: Types of Short-Selling Constraints.

This table reports regression estimates of our baseline model from equation (8) on the link between short selling and investment-Q sensitivity, considering different types of short selling constraints and restrictions. The dependent variable in all specifications is investment, defined as capital expenditures divided by lagged property, plant, and equipment (PPE). *Short Legal*, *Naked Short Legal*, and *Uptick Rule* are dummy variables taking the value of one if (naked) short selling is permitted in the given year and if an uptick rule, restricting short sales to prices above the last trade, is enforced (in countries that allow short selling), respectively. Columns 1 and 2 consider the complete sample of firm-year observations, columns 3 and 4 condition the sample to firm-years in which short selling was permitted. Each regression includes an intercept and similar additional firm-level controls and interactions as the previous tables, i.e. the logarithm of a firm's total book assets (*log AT*), book leverage (*Leverage*), cash flow (*CF*), interactions of these variables with Tobin's Q and the logarithm of country GDP, all lagged by one year. The sample period is from 1994 to 2014. The industry fixed effects are based on two digit SIC codes, continuous covariates are winsorized at the 1% level. *t*-statistics provided in parentheses are computed based on standard errors clustered at the firm level in all specifications. ∗, ∗∗ and ∗∗∗ indicate statistical significance at the 10%, 5% and 1% level, respectively.

	Dependent Variable: CapEx (t) / PPE (t-1)			
	Sample: All		Sample: Short Legal = Yes	
	(1)	(2)	(3)	(4)
Q (t-1)	0.0502***	0.0775***	0.0934***	0.0978***
	(9.80)	(15.63)	(15.16)	(16.99)
Short Legal (t-1)	-0.0731***			
	(-9.06)			
Q (t-1) * Short Legal (t-1)	**0.0471*** **			
	(9.37)			
Naked Short Legal (t-1)		-0.0280***	-0.00139	
		(-3.67)	(-0.16)	
Q (t-1) * Naked Short Legal (t-1)		**0.0225*** **	**0.00555**	
		(4.04)	**(0.85)**	
Q (t-1) * Uptick Rule (t-1)				**-0.00548**
				(-0.87)
log(AT) (t-1)	-0.0209***	-0.0205***	-0.0231***	-0.0231***
	(-18.31)	(-17.97)	(-18.84)	(-18.83)
Leverage (t-1)	-0.127***	-0.114***	-0.140***	-0.140***
	(-6.16)	(-5.50)	(-5.69)	(-5.74)
CF (t-1)	0.354***	0.372***	0.329***	0.326***
	(10.52)	(11.05)	(9.18)	(9.09)
log(GDP) (t-1)	0.0195***	0.0170**	-0.0191*	-0.0167*
	(3.04)	(2.34)	(-1.79)	(-1.68)
Q (t-1) * Leverage (t-1)	-0.000540	-0.0106	0.00970	0.0106
	(-0.03)	(-0.66)	(0.49)	(0.54)
Q (t-1) * CF (t-1)	-0.127***	-0.137***	-0.132***	-0.130***
	(-7.10)	(-7.65)	(-6.93)	(-6.82)
Year FE	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes
Number of firm-years	127317	127317	106119	106119
R^2	0.141	0.140	0.150	0.150
Adj. R^2	0.140	0.139	0.149	0.149

Table 7: The Impact of Short-Selling on M&A Deal-Completion-to-Return Sensitivity.

This table reports regression estimates of equation (9) on the importance of short selling availability for the link between the completion of announced M&A transactions and the market reaction on announcement day. The dependent variable in all specifications is the dummy variable *M&A Deal Completed*, which takes the value of one, if an announced M&A deal was ultimately completed, and zero otherwise. Despite the binary dependent variable, we estimate this model as a linear probability regression to avoid issues with incidental variables and interaction terms in non-linear models in al specifications. In columns 1 to 3, *Short Legal* is a dummy variable taking the value of one if short selling is permitted in the given year and country and zero otherwise. In column 4 *Short Legal* is defined similarly using naked short selling instead. *Exc. Ret [0,0]* is the stock return of the acquiring firm on the announcement day of the M&A transaction minus the market return in the local stock market on the same day. All further firm and deal specific control variables are explained in detail in the main text and in the appendix. Each regression includes an intercept and similar additional firm-level controls and interactions as the previous tables, i.e. the logarithm of a firm's total book assets (*log AT*), book leverage (*Leverage*), cash flow (*CF*), interactions of these variables with Tobin's Q and the logarithm of country GDP, all lagged by one year. The sample period is from 1987 to 2014. In the last lines of the table, we report the p-value and the test statistic of an F-test that evaluates whether the coefficients on ExcRet × Short Legal and ExcRet × Short Not Legal are equal. The industry fixed effects are based on two digit SIC codes, continuous covariates are winsorized at the 1% level. *t*-statistics provided in parentheses are computed based on standard errors clustered at the firm level in all specifications. $*$, $**$ and $***$ indicate statistical significance at the 10%, 5% and 1% level, respectively.

	Dependent Variable: M&A Deal Completed (Yes/No)			
	Short Selling			Naked Short Selling
	(1)	(2)	(3)	(4)
Short Legal	0.0143	0.0112	0.00756	**-0.0172**
	(0.71)	(0.54)	(0.47)	**(-1.46)**
Exc. Ret. [0,0]	**0.335*** **			
	(3.94)			
A: Exc. Ret. [0,0] * (Short Legal = Yes)		**0.409*** **	**0.355*** **	**0.306** **
		(5.39)	**(5.23)**	**(2.66)**
B: Exc. Ret. [0,0] * (Short Legal = No)		**-0.0573**	**0.0333**	**0.312** **
		(-0.33)	**(0.20)**	**(2.49)**
Mkt. Cap Acq.	5.21e-09	4.99e-09	5.03e-09	5.75e-09
	(1.67)	(1.60)	(1.37)	(1.46)
Deal Value / Mkt. Cap. Acq.	-0.117***	-0.118***	-0.0916***	-0.0912***
	(-9.38)	(-9.28)	(-5.02)	(-5.15)
% sought	-0.000340***	-0.000340***	-0.000189**	-0.000230***
	(-3.41)	(-3.47)	(-2.37)	(-2.76)
% stocks used as payment	-0.000456***	-0.000443***	-0.000435***	-0.000446***
	(-3.17)	(-3.14)	(-3.93)	(-3.68)
Hostile (y/n)	-0.413***	-0.411***	-0.0960*	-0.0974*
	(-8.47)	(-8.43)	(-1.77)	(-1.78)
White Knight (y/n)			-0.0678	-0.0704
			(-0.34)	(-0.35)
Block Purchase (y/n)			0.00924	0.00859
			(0.23)	(0.20)
Challenged Deal (y/n)			-0.427***	-0.426***
			(-9.78)	(-9.80)
Unsolicited (y/n)			-0.284***	-0.283***
			(-7.35)	(-7.44)
Asset Lockup (y/n)			0.0979**	0.0945**
			(2.65)	(2.50)
Poison Pill (y/n)			-0.448*	-0.439*
			(-1.88)	(-1.81)
Intercept	Yes	Yes	Yes	Yes
Additional Firm Controls	Yes	Yes	Yes	Yes
Monthly Time FE	Yes	Yes	Yes	Yes
Acquirer Country FE	Yes	Yes	Yes	Yes
Acquirer Industry FE	Yes	Yes	Yes	Yes
Target Country FE	Yes	Yes	Yes	Yes
Firm-level Clustered SE	Yes	Yes	Yes	Yes
Number of deals	8151	8151	7358	7355
R^2	0.163	0.164	0.234	0.240
Adj. R^2	0.107	0.108	0.186	0.184
p-value (A = B)		5.91**	2.87*	0.9753
F-Stat		0.0192	0.0973	0.000

Table 8: Short Selling and the Managerial Learning Channel.

This table reports OLS regression estimates of equation (10) on the importance of short selling availability for the managerial learning channel. We introduce an additional layer of interactions to our baseline model in equation (8), using proxies for the ex-ante incentives of market participants to acquire private information to estimate cross-sectional differences for the link between short selling availability and firms' investment to Tobin's Q sensitivity. The dependent variable in all specifications is investment, defined as capital expenditures divided by lagged property, plant, and equipment (PPE). In column 1, the proxy for total informativeness of prices, *(1-R2)*, is the firm-specific return variation as in Chen et al. (2007). In columns 2 and 3 we use *Analyst Coverage* is from I/B/E/S, *Foreign Institutional Ownership* is from FactSet LionShares to test the 'crowding out hypothesis', and in column 4 *Cross Listed* is a dummy variable taking the value of one if the firm is cross listed on a U.S. stock exchange in the given year, following Foucault and Frésard (2012). *Short Legal* and *Q* are defined similarly as in Tables 3 and 4. Each regression includes an intercept and similar firm-level controls and interactions as the previous tables, i.e. the logarithm of a firm's total book assets (*log AT*), book leverage (*Leverage*), cash flow (*CF*), interactions of these variables with Tobin's Q and the logarithm of country GDP, all lagged by one year. The sample period is from 1994 to 2014. The industry fixed effects are based on two digit SIC codes. All explanatory variables are defined in the appendix, continuous covariates are winsorized at the 1% level. *t*-statistics provided in parentheses are computed based on standard errors clustered at the firm level in all specifications. $*$, $**$ and $***$ indicate statistical significance at the 10%, 5% and 1% level, respectively.

	CapEx(t)/PPE(t-1)			
	(1)	(2)	(3)	(4)
Q (t-1)	0.0254*	0.0500***	0.0481***	0.0491***
	(1.95)	(8.91)	(6.71)	(9.19)
Short Legal (t-1)	-0.0654***	-0.0754***	-0.0646***	-0.0705***
	(-7.81)	(-9.05)	(-6.00)	(-8.32)
Q (t-1) * Short Legal (t-1)	0.0217*	0.0682***	0.0565***	0.0516***
	(1.79)	(11.70)	(8.04)	(9.37)
Q (t-1) * (1-R2) (t-1)	**0.0297***			
	(1.92)			
Q (t-1) * Short Legal * (1-R2) (t-1)	**0.0342****			
	(2.35)			
Q (t-1) * Analyst Coverage (t-1)		**-0.0112*****		
		(-3.69)		
Q (t-1) * Short Legal * Analyst Coverage (t-1)		**-0.00296*****		
		(-11.64)		
Q (t-1) * Inst. Own. For. (t-1)			**0.00301**	
			(0.84)	
Q (t-1) * Short Legal (t-1) * Inst. Own. For. (t-1)			**-0.129*****	
			(-3.54)	
Q (t-1) * Cross Listed (t-1)				**0.0257***
				(1.67)
Q (t-1) * Short Legal (t-1) * Cross Listed (t-1)				**-0.0607*****
				(-3.79)
(1-R2) (t-1)	-0.0905***	-0.0152*	-0.0129	-0.0154*
	(-4.31)	(-1.86)	(-1.32)	(-1.90)
Analyst Coverage (t-1)	0.000411	0.00418***	0.000222	0.0000510
	(1.26)	(9.61)	(0.51)	(0.16)
Inst. Own. Foreign (t-1)			0.250***	
			(4.32)	
Cross Listed (t-1)				0.0109
				(0.57)
Short Legal (t-1) * Cross Listed (t-1)				0.0486**
				(2.20)
Firm Controls and Interactions	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes
Number of firm-years	114804	114804	72785	116804
R^2	0.146	0.148	0.164	0.145
Adj. R^2	0.145	0.147	0.163	0.144

Table 9: Short Selling and Operating Performance.

This table reports second stage OLS regression estimates of the link between short selling availability and firms' ex post operating performance. The dependent variable in Panel A is the return on assets (ROA) in columns 1 and 2, and sales growth in columns 3 and 4. In Panel B we use the average ROA and sales growth over the next three years as the dependent variable. Our main variable of interest is the dummy variable *Positive*, which is constructed as follows. In a first stage regression, we estimate our main model from equation (1), without the interaction term Q × Short Legal. *Positive* then takes the value of one, if the regression residual from the first stage regression is positive, and zero otherwise. In each specification we 'skip one year', i.e. our dependent variable is observed in year t+1 and *Positive* is observed in t-1, to avoid collinearity issues due to common denominators. In columns 2, 4, 6, and 8 we additionally include the dummy variable *Short Legal*, which takes the value of one if short selling is permitted in the given year and country and zero otherwise. Each regression includes an intercept and similar firm-level controls as the previous tables, i.e. *log(AT)*, *Leverage*, *CF*, and country GDP, all lagged by one year, as well as firm and year fixed effects. All explanatory variables are defined in the appendix, continuous covariates are winsorized at the 1% level. *t*-statistics provided in parentheses are computed based on standard errors clustered at the firm level in all specifications. ∗, ∗∗ and ∗∗∗ indicate statistical significance at the 10%, 5% and 1% level, respectively.

	Panel A: Performance (following year)				Panel B: Performance (Average of following 3 years)			
	ROA (t+1)		Sales Growth (t+1)		Avg. of ROA(t+1 to t+3)		Avg. Sales Growth (t+1 to t+3)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Positive (t-1)	**0.00238*** **	**0.00236*** **	**0.00637***	**0.00654***	**0.00188*** **	**0.00188*** **	**0.00542****	**0.00530****
	(2.86)	**(2.84)**	**(1.87)**	**(1.93)**	**(3.59)**	**(3.59)**	**(2.54)**	**(2.48)**
Short Legal (t-1)		-0.00159		0.0195**		-0.000544		-0.0378***
		(-0.79)		(2.26)		(-0.32)		(-6.99)
Q (t-1)	0.0114***	0.0114***	-0.0000368	0.000273	0.00374***	0.00372***	-0.00176	-0.00264
	(9.04)	(9.02)	(-0.01)	(0.06)	(3.84)	(3.83)	(-0.45)	(-0.68)
ROA (t-1)	-0.0546**	-0.0545**	0.0893	0.0888	-0.0486***	-0.0486***	0.0237	0.0246
	(-2.44)	(-2.44)	(1.21)	(1.20)	(-3.85)	(-3.85)	(0.60)	(0.63)
Sales Growth (t-1)	0.00321*	0.00321*	-0.109***	-0.109***	0.00439***	0.00439***	-0.0587***	-0.0586***
	(1.84)	(1.84)	(-12.89)	(-12.89)	(4.07)	(4.07)	(-11.85)	(-11.86)
log(AT) (t-1)	-0.0337***	-0.0338***	-0.188***	-0.188***	-0.0331***	-0.0331***	-0.153***	-0.154***
	(-13.92)	(-13.94)	(-24.38)	(-24.35)	(-14.93)	(-14.95)	(-23.41)	(-23.62)
Leverage (t-1)	0.0123*	0.0123*	0.0296	0.0301	0.0350***	0.0350***	0.0473***	0.0445**
	(1.84)	(1.84)	(1.23)	(1.25)	(6.44)	(6.43)	(2.66)	(2.51)
CF (t-1)	0.0795***	0.0795***	-0.170**	-0.169**	0.0202	0.0202	-0.0324	-0.0328
	(3.57)	(3.57)	(-2.05)	(-2.05)	(1.63)	(1.62)	(-0.71)	(-0.72)
log(GDP) (t-1)	0.0219***	0.0222***	0.0488***	0.0458***	0.0180***	0.0181***	0.0380***	0.0457***
	(7.65)	(7.72)	(4.67)	(4.36)	(6.70)	(6.72)	(4.19)	(5.02)
Firm and Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Number of firm-years	80852	80852	81720	81720	54751	54751	55967	55967
R^2	0.561	0.561	0.281	0.281	0.781	0.781	0.553	0.554
Adj. R^2	0.484	0.484	0.156	0.156	0.739	0.739	0.469	0.470

Table 10: Market Development, Corporate Governance, and Disclosure Quality.

This table reports OLS regression estimates of the link between short selling availability and firms' investment to Tobin's Q sensitivity, for several proxies of market development, governance, and information quality proxies at the country level. The dependent variable in all specifications is investment, defined as capital expenditures divided by lagged property, plant, and equipment (PPE). *Short Legal* is a dummy variable taking the value of one if short selling is permitted in the given year and country and zero otherwise. In each column, we define two dummy variables *High* and *Low* taking the value of one if the firm's country is above (below) the median in market development, governance standards, or information quality. In column 1 we use GDP per capita in USD obtained from the World Bank to partition our sample. Columns 2 and 3 use the number of stock listings, and total market capitalization per population from the World Federation of Exchanges to define *High* and *Low*. In columns 4 and 5 we use the legal origin (common law origin is defined high) and the Anti Self-Dealing Index from Djankov et al. (2008), and in column 6 we use the Earnings Management Pervasiveness Index from La Porta et al. (2006) to split the sample. Each regression includes an intercept and similar firm-level controls and interactions as the previous tables, i.e. the logarithm of a firm's total book assets (*log AT*), book leverage (*Leverage*), cash flow (*CF*), interactions of these variables with Tobin's Q and the logarithm of country GDP, all lagged by one year. The industry fixed effects are based on two digit SIC codes. In the last lines of the table, we report the p-value and the test statistic of an F-test that evaluates whether the coefficients on Q × Short Legal × Low and Q × Short Legal × High are equal. All explanatory variables are defined in the appendix, continuous covariates are winsorized at the 1% level. *t*-statistics provided in parentheses are computed based on standard errors clustered at the firm level in all specifications. ∗, ∗∗ and ∗∗∗ indicate statistical significance at the 10%, 5% and 1% level, respectively.

	CapEx(t)/PPE(t-1)					
	GDP / Capita (t-1)	Stock Listings / Population (t-1)	Mkt Cap. / Population (t-1)	Legal Origin	Anti Self-Dealing Index	Earnings Mgt. Pervasiveness
	(1)	(2)	(3)	(4)	(5)	(6)
Q (t-1)	0.0487***	0.0503***	0.0501***	0.0445***	0.0444***	0.0646***
	(9.51)	(9.83)	(9.79)	(8.37)	(8.37)	(5.03)
Short legal (t-1)	-0.0590***	-0.0729***	-0.0744***	-0.0686***	-0.0680***	-0.0411**
	(-7.24)	(-9.03)	(-9.05)	(-8.33)	(-8.26)	(-2.54)
A: Q (t-1) * Short Legal (t-1) * High (t-1)	0.0554***	0.0460***	0.0453***	0.0544***	0.0538***	0.0317**
	(10.34)	(8.67)	(8.48)	(9.67)	(9.68)	(2.30)
B: Q (t-1) * Short Legal (t-1) * Low (t-1)	0.0339***	0.0482***	0.0493***	0.0629***	0.0661***	0.0383***
	(6.48)	(9.12)	(8.99)	(8.12)	(7.96)	(2.95)
Firm Controls and Interactions	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Number of firm-years	127317	127317	127317	117839	117839	89625
R^2	0.141	0.141	0.141	0.146	0.146	0.160
Adj. R^2	0.140	0.140	0.140	0.145	0.145	0.159
(A) = (B) (*p*-value)	0.000	0.450	0.980	0.254	0.124	0.362
F-Stat	36.00***	0.50	0.32	1.30	2.37	0.83

Appendix

Table A1: Short Selling Regulation around the World from Jain et al. (2013).

Country	Short Selling Legal	Naked Short Selling Legal	Uptick Rule
Argentina	Since 1999	No	Yes
Australia	Yes, except 2008	No since 2001	n/a
Belgium	Since inception	Since inception	No
Brazil	Since inception	No	n/a
Chile	Since 1999	No	Yes
China	No	No	n/a
Colombia	No	No	n/a
Cyprus	No	No	n/a
Czech Rep.	Since inception	Since inception	No
Denmark	Since inception	Since inception	No
Finland	Since 1998	Since 1998	No
France	Since inception	Since inception	No
Germany	Since inception	Since inception	No
Greece	Yes, except 2008 and 2009	No	Yes
Hong Kong	Since 1994	No	Yes
India	Since 2008	No	n/a
Indonesia	Yes, except 2008 and 2009	n/a	n/a
Ireland	Yes, until 2008	n/a	n/a
Israel	Since inception	No	n/a
Italy	Yes, except 2008 and 2009	Yes, except 2008 and 2009	No
Japan	Since inception	Yes, until 2008	Yes
Luxembourg	Since inception	Since inception	No
Malaysia	Yes, before 1997 and since 2007	No	Yes
Mexico	Since inception	No	Yes
Netherlands	Since inception	Yes, except 2008 and 2009	No
New Zealand	Since 2000	n/a	n/a
Norway	Since 1992	No	No
Pakistan	Since inception	Yes, until 2002	Yes
Peru	No	No	n/a
Philippines	Since 1998	No	Yes
Poland	Since 2000	Since 2000	n/a
Portugal	Since inception	Since inception	No
Russia	Yes, except 2008 and 2009	Yes, except 2008 and 2009	Yes
Singapore	Since inception	No	No
South Africa	Since inception	No	No
South Korea	Yes, 1996 to 2007, since 2009	Yes, until 2000	Yes
Spain	Yes, since 1992	No	n/a
Sri Lanka	No	No	n/a
Sweden	Yes, since 1991	Yes, since 1991	No
Switzerland	Since inception	Yes, except 2008	No
Taiwan	Yes, until 2008	Yes, until 2008	Yes
Thailand	Yes, since January 2001	No	Yes
Turkey	Since inception	Since inception	Yes
UK	Since inception	Since inception	No
Venezuela	No	No	n/a

Table A2: Variable Definitions.

Variable	Description	Source
CapEx(t)/PPE(t-1)	Capital expenditure in year t, scaled by one year lagged tangible assets.	Compustat Global
ΔAT(t)/AT(t-1)	Change in total assets, scaled by one year lagged total assets.	Compustat Global
ΔPPE(t)/PPE(t-1)	Change in tangible assets, scaled by one year lagged tangible assets.	Compustat Global
(CapEx+R&D)(t)/AT(t-1)	Capital expenditure plus research and development expenses in year t, scaled by one year lagged total assets.	Compustat Global
R&D(t)/PPE(t-1)	Research and development expenses, scaled by lagged tangible assets.	Compustat Global
Tobin's Q	(Total assets + market capitalization - shareholders' equity)/total assets.	Compustat Global
log(AT)	Logarithm of total assets in U.S. dollars.	Compustat Global
Mkt. Cap.	Market capitalization in U.S. dollars.	Compustat Global
Book Leverage	(Long term debt + short term debt)/total assets.	Compustat Global
Sales Growth	One-year change in total revenues, scaled by lagged total revenues.	Compustat Global
ROA	Return on Assets, calculated as net income over total assets.	Compustat Global
CF	Cash Flow, calculated as (depreciation and amortization + income before extraordinary items)/lagged total assets.	Compustat Global
Inst. Own. Foreign	Percentage of outstanding shares held by foreign institutional investors.	FactSet LionShares
Analyst Coverage	Number of analysts issuing at least one forecasting in the given year.	I/B/E/S International
(1-R2)	Firm-specific return variation, defined as the one minus the R2 of a regression of firm-level weekly equity returns on local market returns. This measure is defined at the firm-year level.	Datastream and Compustat Global
Speed of Information Diffusion	Ratio of the R^2 of a regression of firm-level weekly returns on local market returns without lags over the R^2 of the same regression with four weekly lags of the local market return. This measure is defined at the firm level.	Datastream and Compustat Global
FHT Trading Costs	Fong-Holden-Trzcinka measure of trading costs, calculated as $2\sigma N^{-1}(\frac{1+z}{2})$, where $N^{-1}()$ is the inverse cumulative normal distribution, z is the ratio of zero return days to calendar days, and σ is the standard deviation of non-zero returns. The measure is defined at the firm-year level.	Compustat Global
|VR5 -1|	Deviation from random walk, calculated as the absolute value of the five-week variance ratio (1/5 times the variance of 5 week return, divided by one-week return variance) minus one. The measure is defined at the firm-year level.	Compustat Global
Cross Listed (y/n)	Indicator variable taking the value of one if the firm has a cross-listed security on a U.S. stock exchange in the given year.	BNY Mellon, JP Morgan, Citibank, and NASDAQ
log(GDP)	Logarithm of current GDP per capita in USD.	World Bank
Excess Return [0,0]	Stock return of acquiring firm minus local market return on announcement day.	Datastream and Compustat Global
Short Legal	Indicator variable taking the value of one if short selling is permitted.	Jain et al. (2013)

continued

Deal Value	Value of Transaction in Million USD.	SDC Platinum
MCap Acquirer	Market capitalization of acquiring firm in U.S. dollars	SDC Platinum
% sought	Percentage of shares of target firm sought by acquirer.	SDC Platinum
% stock payment	Percentage of transaction paid for in stock by acquirer.	SDC Platinum
Hostile (y/n)	Deal attitude, dummy variable taking the value of one if attitude was hostile.	SDC Platinum
White Knight (y/n)	Dummy variable taking the value of one if 'white knight' was involved.	SDC Platinum
Block Purchase (y/n)	Dummy variable taking the value of one if transaction was a block purchase.	SDC Platinum
Challenged Deal (y/n)	Dummy variable taking the value of one if deal was challenged.	SDC Platinum
Unsolicited (y/n)	Dummy variable taking the value of one if deal was unsolicited	SDC Platinum
Asset Lockup (y/n)	Dummy variable taking the value of one if target had 'asset lockup' anti-takeover provision.	SDC Platinum
Poison Pill (y/n)	Dummy variable taking the value of one if target had 'poison pill' anti-takeover provision.	SDC Platinum

Table A3: Information Environment, Liquidity, and M&A Deal-Completion-to-Return Sensitivity.

This table reports regression estimates of equation (9) on the importance of private information in stock prices for the link between the completion of announced M&A transactions and the market reaction on announcement day. The dependent variable in all specifications is the dummy variable *M&A Deal Completed*, which takes the value of one, if an announced M&A deal was ultimately completed, and zero otherwise. Despite the binary dependent variable, we estimate this model as a linear probability regression to avoid issues with incidental variables and interaction terms in non-linear models in al specifications. *Excess Return [0,0]* is the stock return of the acquiring firm on the announcement day of the M&A transaction minus the market return in the local stock market on the same day. *(Cross Listed = No)* is a dummy variable that takes the value of one if the firm is *not* cross-listed on a U.S. stock exchange in the given year, and *Illiquidity* is the proportion of zero return days of the firm's stock per total number of trading days in a given year. All further firm and deal specific control variables are explained in detail in the main text and in the appendix. Each regression includes an intercept and similar additional firm-level controls and interactions as the previous tables. The sample period is from 1987 to 2014. The industry fixed effects are based on two digit SIC codes, continuous covariates are winsorized at the 1% level. *t*-statistics provided in parentheses are computed based on standard errors clustered at the firm level in all specifications. $*$, $**$ and $***$ indicate statistical significance at the 10%, 5% and 1% level, respectively.

	Dependent Variable: M&A Deal Completed (Yes/No)	
	(1)	(2)
Excess Return [0,0]	0.366***	0.246***
	(2.90)	(3.74)
Excess Return [0,0] * (Cross Listed = No)	**-0.272****	
	(-2.00)	
Excess Return [0,0] * Illiquidity		**-0.441****
		(-2.10)
Cross Listed (y/n)	-0.00155	-0.000106
	(-0.36)	(-0.03)
Illiquidity	0.0358**	0.0405**
	(2.04)	(2.30)
Mkt. Cap Acquirer	4.21e-09***	4.28e-09***
	(2.77)	(2.82)
Deal Value / Mkt. Cap. Acquirer	-0.0813***	-0.0804***
	(-8.45)	(-8.39)
% sought	-0.000257***	-0.000260***
	(-4.27)	(-4.32)
% stocks used as payment	-0.000497***	-0.000500***
	(-8.24)	(-8.28)
Hostile (y/n)	-0.0841	-0.0835
	(-1.49)	(-1.48)
Target Private (y/n)	0.00557	0.00567
	(1.56)	(1.59)
White Knight (y/n)	-0.0976	-0.0948
	(-0.75)	(-0.73)
Block Purchase (y/n)	0.0253	0.0252
	(1.36)	(1.35)
Challenged Deal (y/n)	-0.414***	-0.414***
	(-12.28)	(-12.26)
Unsolicited (y/n)	-0.292***	-0.293***
	(-7.59)	(-7.61)
Lockup (y/n)	0.0772**	0.0767**
	(2.03)	(2.01)
Poison Pill (y/n)	-0.460***	-0.458***
	(-2.70)	(-2.68)
Deal was rumoured (y/n)	-0.0106	-0.0112
	(-1.02)	(-1.08)
Intercept	Yes	Yes
Monthly Time FE	Yes	Yes
Acquirer Country FE	Yes	Yes
Acquirer Industry FE	Yes	Yes
Target Country FE	Yes	Yes
Firm-level Clustered SE	Yes	Yes
Number of firm-years	16772	16772
R^2	0.188	0.188
Adj. R^2	0.159	0.159