

***Ex ante* Competitive Threats and Stock Price Crash Risk: Evidence from International Antitrust Leniency Programs**

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ABSTRACT

We examine empirically the impact of *ex ante* competitive threats from potential rivals on bad news hoarding/disclosure decisions utilizing the staggered adoption and strengthening of international antitrust leniency laws as an exogenous shock to U.S. product markets. We find that stronger antitrust enforcement around the world causes domestic U.S. firms to exhibit lower levels of stock price crash risk. This is consistent with managers disclosing rather than hoarding bad news to deter entry from potential rivals. Our findings are robust to a battery of sensitivity tests including temporal dynamics test, and staggered DiD bias correction methods. Cross-sectional analyses show that the negative impact of competitive threats from potential rivals on stock price crash risk is more salient for firms with a lower percentage of independent board directors, longer tenure of board directors, higher levels of CFO equity incentives and entrenchment index, higher level of risk-taking activities and more opaque financial reporting. Overall, our evidence indicates that competitive threats from potential rivals constrain firms' stock price crash risk, in line with the view that competition serves as an important governance mechanism in curbing agency problems.

JEL Classifications: D22, D43, D82, G12, G32, G34, G38, L44.

Keywords: *ex ante* competitive threats; crash risk; natural experiment; international antitrust programs.

Data Availability Statement

The data that support the findings of this study are available from the sources identified in the paper. Restrictions apply to the availability of these data, which were used under license for this study.

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We examine empirically the impact of *ex ante* competitive threats from potential rivals on bad news hoarding/disclosure decisions utilizing the staggered adoption and strengthening of international antitrust leniency laws as an exogenous shock to U.S. product markets. We find that stronger antitrust enforcement around the world causes domestic U.S. firms to exhibit lower levels of stock price crash risk. This is consistent with managers disclosing rather than hoarding bad news to deter entry from potential rivals. Our findings are robust to a battery of sensitivity tests including temporal dynamics test, and staggered DiD bias correction methods. Cross-sectional analyses show that the negative impact of competitive threats from potential rivals on stock price crash risk is more salient for firms with a lower percentage of independent board directors, longer tenure of board directors, higher levels of CFO equity incentives and entrenchment index, higher level of risk-taking activities and more opaque financial reporting. Overall, our evidence indicates that competitive threats from potential rivals constrain firms' stock price crash risk, in line with the view that competition serves as an important governance mechanism in curbing agency problems.

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“Monopoly, besides, is a great enemy to good management, which can never be universally established but in consequence of that free and universal competition which forces everybody to have recourse to it for the sake of self-defence.”

-Adam Smith (1776), *The Wealth of Nations*

1. Introduction

The economics of competition in the face of agency conflicts has a long history, dating back to Adam Smith (1776) (see, for example, Hicks 1935; Jensen and Meckling 1976; Holmstrom 1982; Hart 1983; Scharfstein 1988; Stein 1989; Schmidt 1997). However, this literature has yielded contradictory empirical results. Some studies document that competition curtails agency problems (e.g., Giroud and Mueller 2010, 2011), whereas other studies show the opposite (e.g., Shleifer 2004). While prior research focuses primarily on existing or close to immediate competition, such as is likely to arise from large tariff reductions or exchange rate movements, our study focuses on potential *ex ante* competitive threats as reflected in the passage of international antitrust leniency laws and programs.¹ Specifically, we examine the influence of international leniency programs on U.S. corporate decision making with respect to bad news disclosure/hoarding.²

Recent academic literature finds that U.S. domestic firms respond to competitive threats from potential competition by up-grading productivity, raising innovation, altering product composition, fostering tacit coordination, and augmenting capital financing (e.g., Bao and Chen 2018; Dasgupta and Zaldakos 2019). Although this research yields valuable insights, it fails to

¹ We use the term *ex ante* competitive threats to denote threats that take time to actualize into full blown competition, not those threats involving immediate or near-immediate competition.

² In this study, international antitrust leniency laws are used interchangeably with foreign (antitrust) leniency laws.

speak to negative information disclosure as an alternative mechanism for dealing with competitive threats. In this study, we posit that US domestic firms use financial reporting and alternative forms of *negative* news disclosure to dissuade potential rivals from competing in US markets.³ Disclosure theory underpins this prediction. In particular, the entry game model by Darren and Stoughton (1990) demonstrates that incumbent firms are often incentivized to disseminate negative news to deter potential competitors from entering the industry. We predict that, in equilibrium, the agency incentives of managers to hoard bad news because of career and compensation concerns are offset by the need to deter potential competition.

This study tests the hypothesis that US domestic firms disclose rather than hoard negative news when facing *ex ante* competition from potential rivals. Testing this hypothesis raises empirical challenges, not the least of which is measuring potential competition. How are we as researchers to know which firms constitute potential rivals, especially if these potential rivals eschew competing in US domestic markets because of negative news disclosures? Moreover, competition from potential rivals and U.S. firms' disclosure choices are likely endogenously determined. To overcome these issues, we utilize the staggered adoption and strengthening of antitrust leniency laws in foreign countries as an instrument to capture *ex ante* competitive threats from potential rivals in U.S. markets (Chen and Rey 2013, Dasgupta and Žaldokas 2019, Bourveau, She and Žaldokas 2020). Leniency laws empower market regulators or courts to offer (partial) amnesty to companies participating in a cartel if they cooperate by revealing information about their collusive practices. The adoption of foreign leniency laws reduces cartelization and facilitates the potential entry of rivals into US domestic markets and, hence, serves as an exogenous shock to the *ex ante* competitiveness of U.S. firms. The adoption of foreign leniency

³ Potential rivals include domestic and foreign potential entrants into the industry.

laws is particularly suited for testing our hypothesis because their adoption leads initially to competitive threats, but not immediate or intensified competition per se, insofar as it takes time for international cartels to break down and foreign antitrust authorities to interact with their U.S. counterpart. (e.g., Harrington 2008; Miller 2009; Borrell, Jimenez, and Garcia 2014; Harrington and Chang 2015; Zhou 2016). This differs significantly from other measures of competition as discussed further below.

Another empirical challenge involves measuring bad news hoarding and disclosure. There are many avenues by which bad economic news can be hoarded. For example, bad news hoarding can be achieved through earnings manipulation, classification shifting, off-balance sheet devices, opaque financial statement notes, earnings guidance and conference calls. By the same token, there are many ways by which negative news can be disclosed including financial statements, notes to financial statements, earnings guidance, trade journals, and conference calls to name a few. How are we, as researchers, to know which medium was used to deter entry by potential rivals or, contrariwise, to hoard bad news? An analysis based on a specific channel metric may well yield false inferences regarding a firm's bad news hoarding/disclosure decision. To overcome this issue, we use stock price crash risk to measure bad news hoarding/disclosure. Extant research on crash risk (i.e., Jin and Myers 2006; Hutton, Marcus, and Tehranian 2009; Kim, Li, and Zhang 2011a and 2011b; Callen and Fang 2015) shows that the market-based risk measure of stock price crash risk is a comprehensive metric that reflects all manner of bad news hoarding/disclosure irrespective of the specific channels by which bad news is suppressed or disclosed.

We predict that stock price crash risk of US domestic firms facing *ex ante* competitive threats from the passage of international antitrust leniency legislation will be reduced relative to US domestic firms unaffected by the passage of antitrust leniency regulations. This follows

because foreign antitrust leniency legislation facilitates cooperation among foreign and US antitrust authorities to break up international cartels, decreasing collusion among firms and industry barriers to entry. With a reduction in cartels and collusion and the reduced barriers to entry, potential rivals are more likely to enter US markets. Facing this competitive threat, affected US firms will strategically disclose bad news on a timely basis instead of hoarding it to prevent entry by potential rivals, leading to a concomitant reduction in price crash risk, by comparison to US firms unaffected by antitrust leniency regulations.

Using a large sample of U.S. public firms from 1989 to 2013, we find results consistent with the bad news disclosure to deter potential entry by firms facing *ex ante* competitive threats from potential rivals. Specifically, based on staggered Difference-in-Differences (DiD) analyses, we provide empirical evidence that after foreign leniency laws are passed or strengthened, US firms that operate in industries exposed to foreign imports exhibit significantly lower levels of stock price crash risk relative to firms unaffected by foreign leniency laws. Our findings are robust to a battery of sensitivity tests including temporal dynamics test, alternative model specifications, additional controls, and falsification tests. To mitigate concerns about potential biases induced by staggered DiD analyses, we follow Baker, Larcker, and Wang (2021) and implement bias correction methods, and our findings hold.

To further buttress and enrich our understanding of the relation between foreign antitrust leniency laws and bad news hoarding/disclosure, we conduct cross-sectional analyses that exploit settings where the negative relation between foreign leniency laws and stock price crash risk is predictably stronger. We expect that the impact of *ex ante* competitive threats from foreign leniency laws on stock price crash risk is more pronounced (more negative) for firms with severe agency conflicts. The latter firms are prone *ex ante* to disclose bad news and thus are likely to experience a greater decrease (increase) in bad news hoarding (disclosure) following the passing

or strengthening of foreign leniency laws than their counterparts. The cross-sectional results conform to intuition. We show that the impact of competitive threats arising from antitrust leniency programs on stock price crash risk is more pronounced for firms with weaker corporate governance, higher level of risk-taking activities and more opaque financial reporting. These cross-sectional findings shed light on how competition interacts with agency concerns to affect the impact of bad news hoarding/disclosure decisions on capital markets.

Further, we complement our primary results by documenting three important benefits of timely bad news disclosure. First, we find that firms face less competitive pressure after the adoption of foreign leniency laws, confirming the effectiveness of timely bad news disclosure in preempting future potential entry. Second, we show that the impact of foreign leniency laws on stock price crash risk has asset pricing implications. In particular, the adoption of antitrust leniency laws reduces firms' costs of equity capital through a chain effect on stock price crash risk. This evidence helps to corroborate that stock price crash risk is an important consequence of foreign antitrust leniency laws. Last, we document that the passage of foreign leniency laws has important implications for the growth of U.S. imports from foreign trading partners.

Our study provides important contributions to the literature. First, our paper builds on the growing literature linking international antitrust enforcement regulation with a variety of corporate decision-making activities, including real investment, equity financing, capital structure, and product market information sharing with peers (e.g., Dasgupta and Žaldokas 2019). To the best of our knowledge, this is the first study to assess the linkage between international antitrust enforcement regulation and incentives for bad news disclosures. Specifically, by emphasizing a unique perspective, higher moments of the stock return distribution (i.e., extreme negative returns), we provide new evidence concerning the direct economic consequences of antitrust enforcement regulation on capital markets.

More specifically, our findings contribute to the decades-long debate on the benefits and costs of antitrust enforcement regulation (e.g., Kauper 1980; Van Den Berg and Slivinski 2009; Sawyer 2019; Morton 2020) by identifying a material benefit that antitrust legislation brings to firms and their shareholders. Recent studies (e.g., Xing, Zhang, and Zhao 2010; Yan 2011; Kelly and Jiang 2014) emphasize the importance of understanding the impact of extreme outcomes in the equity market on the welfare of investors. Our empirical evidence helps to advance understanding of the governance role that antitrust enforcement regulations play in enhancing investor welfare through the mitigation of agency problems.

Relatedly, our study contributes to the literature regarding the impact of legal institutions on corporate information disclosure practices (e.g., La Porta, Lopez-de-Silanes, Shleifer and Vishny 1998; Leuz, Nanda and Wysocki 2003) by focusing on antitrust enforcement regulations and their impact on bad news hoarding/disclosure. Our results echo the view of Ball, Robin and Wu (2003) that institutions play an important role in shaping corporate disclosure practice.

Second, this study adds to recent research on the bad-news hoarding theory of stock price crash risk. A series of recent studies on crash risk suggest that, due to agency conflicts, firm-level characteristics are important determinants of bad-news hoarding activities (e.g. Hutton et al. 2009; Kim, Li, and Zhang 2011a, 2011b; Kim and Zhang 2015; Callen and Fang 2013 and 2015b). An, Chen, Li and Xing (2018) further argue that individualism, as an important national culture characteristic, shapes managerial perception of bad news hoarding, thus leading to crash risk, and they find supporting evidence. Nevertheless, research on the potential impact of regulatory enforcement on stock price crash risk remains scant. The only other study we are aware of is that of Fang Liu, and Xin (2010), which examines the impact of the Sarbanes-Oxley Act (SOX) on stock price crash risk. By approaching the issue from a unique angle, that is, exogenous shocks around the world to competitive threats from potential rivals, our study provides a new regulatory

determinant of stock price crash risk, namely, the impact of international antitrust enforcement regulations on managerial bad-news disclosure behavior in U.S. public firms. Importantly, although the price crash risk literature provides extensive evidence showing that managers tend to withhold bad news because of career, short-term compensation, and other personal concerns, our evidence indicates that these tendencies are mitigated by suitably enforced antitrust regulations.

Third, our analysis complements broad research on the impact of product market competition on various firm outcomes. While most extant studies on product market competition focus on immediate or near-immediate competition, our research utilizes the introduction of antitrust leniency laws to offer *causal* evidence that *ex ante* competitive threats help to shape corporate bad news disclosure policy, especially for firms with severe agency problems. These findings are in line with the view in Giroud and Mueller (2010) that competition serves as an important disciplining force on managers and substitutes for conventional governance mechanisms in mitigating managerial manipulative behavior and agency problems.⁴

The recent literature on competition tends to employ import tariffs and currency exchange rates as instruments for import penetration to capture exogenous shocks to the firm's competitive environment (e.g., Fresard 2010; Valta 2012; Xu 2012, among many). However, our setting is fundamentally different from import penetration used in the prior literature. First, tariff shocks and currency exchange rate movements result in almost immediate competition in product

⁴ Nevertheless, there are differences between our setup and that of Giroud and Mueller (2010). Giroud and Mueller (2010) analyze managerial intent once competition has materialized, unlike the passage of foreign leniency laws for which competition is only an *ex ante* anticipatory threat. Their argument, that competition acts to mitigate management slack and increase operating performance and efficiency, is not relevant to the adoption of foreign leniency laws in the crash risk setting, whereas strategic responses to their adoption such as timely disclosure of bad news is more pertinent.

markets and thus adverse economic consequences to U.S. incumbents.⁵ Managers will try to protect their careers and remuneration from these negative impacts by hoarding bad news, increasing crash risk. By contrast, the passage of international leniency laws is not immediate in terms of their economic impact on product market competition and U.S. incumbents do not experience direct worsening of performance. Instead, incumbents are incentivized to strategically disclose bad news to deter potential threat of future entry and competition, leading to a decrease in crash risk. Hence, our setting provides a powerful testing ground for the impact of *ex ante* competitive threats on capital markets. Furthermore, cartel agreements typically involve price fixing, customer/market allocation, supplier division, bid rigging, and cartel activities other than just restrictions or controls over production or sales of goods (e.g., import and export) (U.S. Department of Justice and Federal Trade Commission 2017). Thus, international leniency laws affect the overall market equilibrium along various dimensions (including non-trade barriers) and are far more comprehensive than import penetration per se. Indeed, the arguments underlying our study are based on *ex ante* competitive threats that are more general than specific import-related threats from say tariff reductions.⁶

Finally, our paper is the first to study empirically the equilibrium tradeoffs that firms face when looking at their bad news disclosure/hoarding decisions. On the one hand, managers have short-termism incentives to hoard bad news to protect their careers and compensation. On the other hand, managers have longer-term *strategic* incentives to disclose bad news to prevent entry

⁵ Hong (2024) compares the crash risk of U.S. stocks three years before and after the event of granting Permanent Normal Trade Relations status to China by the U.S.

⁶ *Ex ante* competitive threats could arise from rivals' future potential for direct investment in monetary or non-monetary forms, predatory pricing, competitive bidding, product market invasion, competition along the supply chain including customers and suppliers, and others.

into their industry by potential rivals and thereby protect their firms' market share. Our results show that resulting equilibrium tradeoff impacts capital markets.

The paper proceeds as follows. Section 2 reviews prior literature and develops our hypotheses. Section 3 describes the sample, variable measurement, and research design. Empirical results are presented in Section 4. Section 5 concludes.

2. Literature Review and Hypotheses Development

2.1. Competition of Potential Rivals

Porter (1979, 1980) maintains that firms formulate strategies that take into account competitive forces in the industries and markets in which they operate. He maintains that the threat of new entrants is one of the most important competitive forces that firms face.⁷ Porter further argues that the potential entrants' anticipation about the reaction of incumbent rivals will have an impact on the former's decision as to whether or not to compete. The potential entrant is more likely to have second thoughts about competing if the following exemplary scenarios occur: 1) incumbents appear likely to reduce prices to maintain market share or because of industry-wide excess capacity; and 2) slow industry growth reduces incumbents' ability to accommodate new arrivals and potentially leads to a decline in the financial performance of all the parties involved. Thus, based on Porter (1979, 1980), the effectiveness of competitive threats from potential entrants is conditional on how existing incumbents in the industry respond to these potential threats.

In their seminal paper, Darrough and Stoughton (1990) model competition in the context of an entry game, in which one firm contemplates entering the market where an incumbent firm

⁷ He states that "new entrants to an industry bring new capacity, the desire to gain market share, and often substantial resources ... The seriousness of the threat of entry depends on the barriers present and on the reaction from existing competitors that entrants can expect."

already manufactures a product. Assuming that the incumbent is endowed with private information, Darrough and Stoughton find that, when entry costs are low (that is, a high threat of potential rival entry), the incumbent discloses bad news to deter the potential rival. In contrast, when entry costs are high, the incumbent is less likely to be incentivized to reveal bad news. The intuition underlying their model is that, given that the entry of new firms weakens the incumbent's competitive advantage, the threat of entry induces the incumbent to release bad news more quickly to discourage new entrants into the industry. Consequently, in equilibrium, firms facing higher competitive threats from potential rivals disclose bad news in a timely fashion. Darrough and Stoughton (1990) suggest the product market considerations act as a countervailing force against managerial incentives for bad news hoarding by inducing managers to disclose bad news in a timely fashion to deter entry. Indeed, Darrough and Stoughton (1990)'s entry game model offers an important theoretical underpinning regarding the incentives for bad news disclosure and against bad news hoarding in our setting.

2.2. Measuring Competition from Potential Foreign Rivals

2.2.1. Standard Measures of Competition

Most extant empirical studies use industry-structure metrics to capture product market competition (e.g., Lang and Stulz 1992; Harris 1998; Rennie 2006; Karuna 2007; Tookes 2008; Ali, Klasa and Yeung 2009; Giroud and Mueller 2010 and 2011; Hadlock and Sonti 2012; Datta, Iskandar-Datta and Singh 2013; Ali, Klasa, and Yeung 2014; Stoughton, Wong and Yi 2017). But the theoretical link between these metrics and competition, including competitive threats from potential rivals, is not at all clear. That is hardly surprising because the implications of these metrics for competition are complicated and ambiguous, depending on the specific type of

competition firms face (Beyer, Cohen, Lys and Walther 2010; Lang and Sul 2014).⁸ Moreover, industry-structure measures themselves are static, reflecting the level of existing competition in the industry at a specific point in time, and so are much less likely to be useful in capturing the competitive threat from potential rivals. Furthermore, identification is a major impediment in testing the relation between competition from potential rivals and price crash risk, if the researcher's intent is to make causality predictions. As noted by Sutton (1991) and Raith (2003), industry structure may be endogenously determined by firm-level characteristics, including competitive business strategy, so that the statistical relation between industry structure and managerial bad news disclosure may be uninformative about the *causal* link between the two. Li, Lin, and Zhang (2018) even challenge the notion of industry structure by arguing that industry structure may simply be capturing neglected capital market benefits or agency costs of disclosure. Recent studies (e.g., Li, Lundholm, and Minnis 2013; Hoberg, Phillips, and Prabhala 2014; Li and Zhan 2019) use firm-level measures of competitive threat from potential entrants, but they too suffer from potential endogeneity issues. These firm-level measures are likely to be endogenous to various firm-level characteristics including firm size, performance, growth opportunities, and even the industry-wide competitive environment.

Li and Zhang (2019) examine the effect of product market competition on firms' stock price crash risk and find that firms facing more competition are more likely to suffer from stock price crashes. They argue that the already-intensified competitive pressure from the product market exacerbates managerial tendencies to hide negative information due to their career and

⁸ For example, Land and Sul (2014) point out that "more concentrated industries are also more competitive based on the theoretical analysis in Raith (2003). This prediction seems surprising since a more traditional view of industry concentration is that higher levels of concentration tend to be associated with lower levels of competition. The basic argument in Raith (2003) is that, as competition through greater product substitutability increases, prices will fall, lowering profits and causing firms to exit the industry."

compensation concerns, which leads to stock price crashes in the future. Li and Zhang (2019) focus on two firm-specific measures of competitive pressure based on firms' 10k filings developed by Hoberg et al. (2014) and Li et al. (2013), respectively. Such measures likely capture *ex post* competitive threats from existing rivals, while our study centers on *ex ante* competitive threats from potential entrants. Essentially, Li and Zhan (2019) and our study focus on two different types of competition (i.e., *ex ante* competitive threats versus *ex post* intensified competition) leading to different crash risk predictions.⁹

2.2.2. Antitrust Leniency Laws as a Measure of Competition from Potential Rivals

Given the ubiquitous nature of cartels and their anti-welfare implications for consumers, governments around the world have devoted substantial resources to break-up cartels and encourage competition. One of the most effective mechanisms for breaking-up cartels involves leniency programs (or leniency laws).¹⁰ Leniency laws give market regulators or courts the power to grant (partial) amnesty to firms belonging to a cartel that cooperate in providing information about collusive behavior. As documented in Chen and Rey (2013), leniency programs are one of the most important antitrust developments for cartel detection and deterrence in recent decades. By decreasing monetary penalties or even offering immunity to cartel members who collaborate in conviction cases (e.g., providing crucial evidence to cartel prosecutors), leniency laws aggravate the costs of forming cartels and increase the benefits of breaking them up (Dasgupta and Žaldokas 2019). A typical leniency law states that the first firm in the cartel to provide

⁹ Our analysis differs from Li and Zhan (2019) in other ways as well. They limit their analysis to the sample period 1998-2009 for firms in the Hoberg-Phillips Data Library. Instead, our study investigates the entire sample of U.S. publicly listed firms available on Compustat for the longer period of 1989 to 2013. Most importantly, in contrast to Li and Zhan (2019), our study is based on a natural experiment reflecting exogenous shocks to competitive threats from potential rivals, which is much less likely to suffer from endogeneity concerns. Indeed, as we shall see, our findings are very different from those of Li and Zhan (2019).

¹⁰ Marvao and Spagnolo (2016) provide a survey of leniency laws and their effectiveness.

considerable evidence to the regulator about cartel behavior (if the latter has not yet obtained enough evidence to prosecute the cartel) is granted automatic amnesty. In countries where the firm's managers, employees and directors face criminal charges for being part of a collusive agreement, amnesty also applies to waiving such charges. Hammond (2005) and Miller (2009) suggest that U.S. leniency laws, introduced in 1993, were in fact quite successful in dismantling existing cartels and preempting the formation of new cartels. This experience encouraged other countries to follow suit and enact similar laws. Dong, Massa, and Žaldokas (2018) show that the world-wide passage of leniency laws significantly destabilized collusion among cartel members, increasing conviction rates and cutting gross margins of affected firms.

Not only are foreign leniency laws costly to foreign firms but U.S. firms that might consider colluding in foreign markets would find it more costly to form international cartels with industry peers, because the passage of more leniency programs makes coordination between antitrust authorities easier.¹¹ In addition, even if U.S. antitrust authorities are unable to bring actions against the suspected cartel in the U.S., the conviction of cartels in foreign jurisdictions can aid affected parties in bringing private civil actions within the U.S.¹² Overall, the passage of international leniency laws increases the costs of collusion in the U.S. and reduces entry barriers into U.S. markets, thus facilitating the entrance of potential rivals.

Though the passage of leniency laws is usually influenced by specific economic and political conditions in each respective country, the timing of the passage of leniency laws does not appear to be driven by one specific trend or reason. Some countries passed leniency laws in response to widely known collusion cases. For instance, Hungary passed leniency laws after facing

¹¹ See <https://www.wsj.com/articles/u-s-competition-policy-is-aligning-with-europe-and-deeper-cooperation-could-follow-11626334844>.

¹² U.S. firms often collude with foreign firms that produce similar products and export them to the U.S. markets.

considerable criticism regarding its antitrust probe against the domestic mobile telephone industry. Switzerland did the same after failing to prosecute firms involved in the vitamin cartel. Other countries passed leniency laws because of significant pressure from outside parties (Lipsky 2009). For example, Mexico passed leniency laws in 2006 following recommendations of the OECD (Organization for Economic Co-operation and Development) Peer Review in 2004. The European Union also influenced its member states to pass leniency laws. The International Monetary Fund and the World Bank regularly require a thorough examination of antitrust laws before providing country loans and other funding (Bradford 2012).¹³ Overall, the staggered passage of leniency laws in different countries is much less likely to be related to decision-making processes of U.S. firms and, thus, should prove useful as an instrument for measuring competitive threats from potential rivals.

2.3. Bad News Hoarding/Disclosure and Crash Price Risk

Prior literature maintains that managers delay the disclosure of bad news as long as possible from outsiders because of career and short-term compensation concerns (e.g., Basu 1997; Graham, Harvey, and Rajgopal 2005; Kothari, Shu, and Wysocki 2009; Ball 2009). Jin and Myers (2006) were the first to explore how information disclosure decisions lead to stock price crash risk. They posit that managers control the disclosure of negative information about the firm to the public, and that a threshold level exists at which managers will stop withholding bad news. They argue that managers are willing to personally absorb limited downside risks and losses related to temporary bad performance by hiding firm-specific bad news. They further point out that “If a sufficiently long run of bad firm-specific news is encountered, insiders give up and all the bad news comes out at once. Giving up means a large negative outlier in the distribution of returns [a

¹³ In some instances, the passage of leniency laws was controversial. See the discussions in Dong, Massa, and Žaldokas (2018), Dasgupta and Žaldokas (2019) and Bourveau, She and Žaldokas (2020).

crash].” Similarly, Hutton et al. (2009) maintain that “firms managing reported earnings shelter bad information up to a point, but that once a threshold is crossed, the information comes out in one fell swoop, at which point a price crash is observed.” In sum, these studies indicate that a stock price crash reflects managerial propensity to withhold or delay the disclosure of bad news.

The empirical evidence supports the bad-news hoarding theory underlying crash risk (e.g., Jin and Myers 2006; Hutton et al. 2009; Kim et al. 2011a, 2011b; Callen and Fang 2013, 2015a, 2015b). Since there are so many options to hoard/disclose bad news in ways not captured by specific channel metrics, this study follows the literature and employs the market-based risk measure of stock price crash risk as a comprehensive metric that should reflect *all* manner of bad news hoarding/disclosure.

2.4. Hypothesis

In equilibrium, managers will trade off the benefits arising from timely disclosure of bad news to thwart entry by potential rivals with personal career and compensation benefits from hoarding bad news. Based on the discussion above, we expect that firms facing *ex ante* competitive threats from potential rivals are more likely to disclose bad news in a timely manner rather than suppress it for an extended period and, as a consequence, are less likely to experience future stock price crashes.¹⁴ This leads to our main hypothesis stated in the alternative:

¹⁴ Compared with collusive equilibria, competitive equilibria should have more “real” bad news events arising from the product market interactions. For example, if sales are anticipated by incumbents to decline, they would have hoarded this information in the absence of FLL. However, with the FLL shock, incumbents are motivated to disclose this information on a timely basis to pre-empt potential entry, thereby leading to more bad news forthcoming. Thus, FLL could change the frequency of bad news in addition to the incentive to hoard it.

Hypothesis: Firms facing *ex ante* competitive threats from potential rivals are associated with lower levels of stock price crash risk.

3. Variables Measurement, Sample, and Descriptive Statistics

3.1. Treatment Variable

We follow the literature (i.e., Dasgupta and Žaldokas 2019; and Bourveau, She and Žaldokas 2020) and construct a treatment variable that measures the exposure of a U.S. firm to the passage of leniency laws in foreign countries from which the firm's industry obtains a portion of its imports. This treatment variable captures *ex ante* competitiveness threats to U.S. firms from potential rivals. Specifically, Foreign Leniency Law (*FLL*) is a continuous variable measured as the weighted average of the passage of leniency laws in all countries other than the United States:

$$FLL_{i,T} = \sum_k w_{k,i} L_{k,T}, \quad (1)$$

where k denotes the foreign country, i the industry (SIC two-digit code) and T the year, respectively. Here, $w_{k,i}$ is the proportion of industry i 's imports from country k . $L_{k,T}$ is equal to a value of one if country k has passed a leniency law by year T , and zero otherwise. To mitigate spurious correlation arising from changes in industry structure or classification, following the literature, time variation is removed by basing the weights on data for the year 1990. The variable *FLL* varies from zero, if no country from which the industry imports passed leniency laws, to one if all foreign countries from which the industry imports passed leniency laws and the industry imports all of its products. Empirically, *FLL* measures the exogenous *ex ante* competitiveness threat at time T from potential rivals to U.S. firms operating in industries with import exposure. See Appendix A Panel A for the full list of foreign countries that passed antitrust leniency laws and the year in which the law passed as replicated from Dasgupta and Žaldokas (2019). The

setting of antitrust leniency programs provides ample variation of time series and cross-industry information to identify the effect of competitive shocks on stock price crash risk.

3.2. Crash Risk Metrics

Following prior literature on crash risk (Chen et al. 2001; Jin and Myers 2006; Hutton et al. 2009), we employ three firm-year specific measures of (*ex post*) stock price crash risk: 1) the negative coefficient of skewness of firm-specific weekly returns (*NCSKEW*); 2) the down-to-up volatility of firm-specific weekly returns (*DUVOL*); and 3) the difference between the number of weeks with negative extreme firm-specific weekly returns and the number of weeks with positive extreme firm-specific weekly returns (*COUNT*). Appendix B describes how to compute these measures of stock price crash risk.

3.3. Control Variables

Following prior literature on crash risk, we control for the following variables: prior period price crash risk measured by $NCSKEW_{T-1}$; KUR_T defined as the kurtosis of firm-specific weekly returns in fiscal year T ; $SIGMA_T$ defined as the standard deviation of firm-specific weekly returns in fiscal year T ; RET_T defined as the cumulative firm-specific weekly return in fiscal year T ; ROA_T defined as the earnings before extraordinary items divided by total assets at the end of fiscal year T ; MB_T defined as the market-to-book ratio at the end of fiscal year T ; LEV_T defined as the book value of total liabilities divided by the total assets at the end of fiscal year T ; $LNSIZE_T$ defined as the log of market value of equity at the end of fiscal year T ; $CVOLA$ and $EVOLA$ defined as the standard deviation of cash flow from operations and earnings excluding extraordinary items and discontinued operations, deflated by the lagged total assets over the current and prior 4 years, respectively; and $DTURN_T$ defined as the average monthly share turnover over fiscal year T minus the average monthly share turnover over the previous year $T-1$, where monthly share

turnover is calculated as the monthly share trading volume divided by the number of shares outstanding over the month.

Following Hutton, Marcus, and Tehranian (2009), we further control for a firm-level earnings manipulation measure of financial reporting opaqueness ($OPAQUE_T$), measured as the absolute value of annual performance-adjusted discretionary accruals--as developed by Kothari, Leone, and Wasley (2005)--for fiscal year T . Following Callen and Fang (2017), we control for the audit-firm-client relationship ($TENURE_T$), measured as the number of consecutive fiscal years that the auditor has been retained by the client, up to and including the current year (in the case of audit firm mergers, the incumbent auditor-client relationship is considered to be the continuation of the prior auditor). We also include as a control an indicator equal to one if the firm is audited by a Big-N auditor, and zero otherwise ($BIGN_T$). In addition, we follow Chen et al. (2001) to control for the number of analysts following the firm (ANA_T) in order to capture short-term earnings pressures from analysts. Chen et al. (2001) find that firms with more analysts following have more crashes in the future. Based on Callen and Fang (2013), we control for institutional ownership measured by the percentage of a firm's shares held by transient (TRA_T), dedicated (DED_T), and quasi-indexer (QIX_T) institutional investors, respectively. Following Bourveau, She and Žaldokas (2020), we also control for extant industry-level competition, i.e., Herfindahl-Hirschman industry index (HHI_T) and industry-level import penetration ($IMPORTPENE_T$). Appendix B summarizes the variable definitions used in this study.

3.4. Data Sources and Sample

The initial sample comprises US listed firms on Compustat for the period of 1989 to 2013. We collect sample data from the following main sources: 1) CRSP stock files to estimate the measures of firm-specific crash risk; 2) Compustat annual files for firm-level accounting data; (3) I/B/E/S for financial analyst data; and 4) Thompson-Reuters Institutional Holdings Database for

institutional ownership.¹⁵ In addition, we obtain information regarding country-level imports from Schott's data library, which is available from the year 1989. The sample data ends in 2013, the year after the most recent passage of a foreign leniency law. We exclude observations with non-positive total assets and equity book values, observations with year-end share price less than one dollar, and observations with fewer than six months of CRSP return data available. We also exclude financial services and utility firms. After imposing these data requirements, our final sample consists of 38,337 firm-year observations.

To offer some perspective on our sample composition, we populate Table 1 with the number of observations by year. The fraction of observations rises at first and declines after 1997. The sample is relatively evenly distributed across the whole sample period so that our empirical evidence is unlikely to be driven by any specific year. In Table 1, we also report the average *FLL* across the sample years. The average *FLL* increases relatively monotonically over the years. The variation of the average *FLL* over the years should provide sufficient power to identify the role that foreign leniency laws play in shaping crash price risk.

-----Insert Table 1 -----

3.5. Descriptive Statistics

Table 2, Panel A presents descriptive statistics for the variables used in the regression models. The mean values of the stock price cash risk measures are -0.0581, -0.0947, and -0.0238, respectively. The standard deviations of these measures are 0.8226, 0.7345, and 0.6357, respectively, and are similar to those documented in the literature (e.g., Chen et al. 2001; Callen and Fang 2013, 2015a, 2015b). The mean value and standard deviation of *FLL* are 0.0591 and 0.1086, respectively, comparable to the statistics reported in prior studies (e.g., Bourveau, She and

¹⁵ Data on corporate ownership for transient institutional investors starting from 1981 are available from Brian Bushee's website at <http://acct3.wharton.upenn.edu/faculty/bushee/>.

Žaldokas 2020). In addition, the industrial composition of our sample is similar to that of Bourveau, She and Žaldokas (2020). Panel B presents a Pearson correlation matrix for the key variables used in this study. *NCSKEW*, *DUVOL*, and *COUNT* are significantly and positively correlated with each other at the one percent level. While these future stock price crash risk measures are calculated differently from firm-specific weekly returns, they appear to contain much the same information.

-----Insert Table 2 -----

4. Empirical Analyses

4.1. Main Results

To examine the effect of foreign leniency laws on firm-specific stock price crash risk, we estimate the following DiD regression equation:

$$CRASHRISK_{j,T} = \alpha_0 + \alpha_1 FLL_{i,T} + \theta X_{ji,T} + \gamma_T + \mu_j + \varepsilon_{j,T}, \quad (2)$$

where *CRASHRISK* is measured by one of *NCSKEW*, *DUVOL*, and *COUNT*. The treatment variable *FLL* captures the staggered effect of exposure to foreign leniency laws on treatment firms' crash risk relative to control firms not exposed to these foreign leniency laws. $X_{ji,T}$ represents a vector of firm-level and industry-level control variables, described in Section 3.3. All regressions control for year (γ_T) and firm fixed-effects (μ_j). Regression equations are estimated using pooled ordinary least squares with White standard errors corrected for industry (i.e., SIC two-digits) clustering.¹⁶ Our focus is on the effect of *FLL* on stock price crash risk, that is, on the coefficient α_1 .

¹⁶ To control for potential outliers, we winsorize the top and bottom 1% regressor outliers--but not the dependent variables following Jin and Myers (2006) and Hutton et al. (2009). However, the results remain essentially unchanged without winsorizing. In addition, our results are robust to controlling for lagged crash risk and clustering standard errors by firm.

When analyzing the impact of foreign leniency laws on stock price crash risk, it is possible that foreign leniency laws and stock price crash risk are simultaneously determined by other variables. Endogeneity concerns arise because of potential omitted unobservable firm characteristics. Hence, we implement firm fixed-effects regressions to help mitigate the concern that omitted time-invariant firm characteristics may be driving the results. Regarding time-varying factors, we included a long list of control variables in the regression model. The year fixed-effects account for unobserved aggregate fluctuations across time.

Table 3 provides the estimation results of regression equation (2). First, Models 1-3 present the results from regressing *CRASHRISK* on *FLL* without controlling for firm-specific variables.¹⁷ The estimated coefficients on *FLL* are negative and significant at the 1% level across all crash risk specifications (*t*-statistics = -3.79, -4.49, and -4.77, respectively). Next, in Models 4-6, we provide the full estimation results of Eq. (2) inclusive of all firm-specific controls. The coefficients on *FLL* are significantly negative at the 1% level (*t*-statistics = -4.07, -4.94, and -4.26, respectively) across all three crash risk measures. These results indicate that competitive threats from potential rivals are negatively associated with stock price crash risk, in line with our hypothesis. These findings are consistent with the view that firms facing *ex ante* competitive threats from potential rivals are more likely to disclose bad news on a timely basis than suppress them for an extended period, yielding less future stock price crashes.

In gauging the economic significance of our results, we compare crash risk across a one standard-deviation shift of the *FLL* distribution, while holding all other variables at their mean values. The results are economically material. Specifically, we estimate an average 0.03365

¹⁷ It is possible that the control variables here (e.g., profitability, firm size, and leverage) are directly affected by the competition shock, which might lead to a problem of “bad controls” (See Angrist and Pischke 2009). Thus, we implement the specification without controls to ensure that the key results hold.

decrease in crash risk across the three alternative measures with a one-standard-deviation shift in the *FLL* distribution. The specific decreases for *NCSKEW*, *DUVOL* and *COUNT* are 0.03938, 0.03398, and 0.02760, respectively. The percentage economic impact of *FLL* is substantial relative to the standard deviation of the crash risk measures. Stock price crash risk decreases by about 4.58% on average with a one-standard-deviation increase in foreign leniency laws. The percentage shift in crash risk corresponding to a one-standard-deviation shift in the distribution of *RET*, *DTURN*, *ANA*, *TRA* and *QIX* is 1.20%, 2.06%, 3.07%, 3.13%, and 4.71%, respectively. Comparing these results indicates that, in terms of economic significance, the estimated impact of foreign leniency laws on firm-specific crash risk is either similar to or larger than the impact of each of the control variables on crash risk.

Turning to the other control variables, we find that the coefficients on *RET*, *DTURN*, *ANA*, and *TRA* are highly significant and are consistent with predictions in the literature. More specifically, consistent with Blanchard and Watson (1982) and Chen et al. (2001), the coefficient on *RET* is significantly positive, suggesting that stock-price bubbles increase future crash risk. In addition, the positive signs on the coefficient of *DTURN* are in line with the explanation of stock price crash in Chen et al. (2001) that when disagreement among investors (and hence trading volume) is large, bearish investors are more likely to sit out the market with their private information incompletely incorporated in prices. This in turn sets the stage for negative skewness in subsequent rounds of trade. In addition, following Chen et al. (2001) that stocks with more analysts following are more prone to crash, the coefficients on *ANA* are significantly positive across all crash risk specifications. Lastly, in consonance with the monitoring theory in Callen and Fang (2013), the coefficients on *TRA* are significantly positive across all crash risk specifications.

Overall, the findings in Table 3 uniformly support our hypothesis that firms facing *ex ante* competitive threats from potential rivals are associated with lower levels of stock price crash risk. These findings are consistent with the view that competitive threats from potential rivals mitigate managerial bad news hoarding by U.S. domestic firms. The results are robust to multiple measures of firm specific crash risk, after controlling for a variety of crash risk determinants.¹⁸

-----Insert Table 3 -----

4.2. Robustness Checks

4.2.1. Temporal Dynamics Test

We implement a temporal dynamics test (i.e., pre-trend analysis) to ensure that no confounding trends affected the outcome variable of stock price crash risk even without foreign leniency programs. We follow prior literature (e.g., Bertrand and Mullainathan 2003) and examine the dynamics of stock price crash risk surrounding the *FLL* events. Specifically, in equation (1), for each foreign country, we introduce a set of dummy variables for years before and after the event year of leniency law passage. We then recalculate *FLL* as the weighted average of the set of new dummies: FLL^{-t} denotes t years prior to and including the event year ($t=0, 1, 2, 3$), respectively. FLL^{1+} denotes at least one year after the event year. The coefficient estimates of FLL^{-t} ($t=1, 2, 3$) are especially important because their significance and magnitude indicate whether there are any confounding trends that affect stock price crash risk even without foreign leniency programs. We estimate the parallel trend based on the model:

$$CRASHRISK_{j,T} = \alpha_0 + \alpha_1 FLL^{-3}_{i,T} + \alpha_2 FLL^{-2}_{i,T} + \alpha_3 FLL^{-1}_{i,T} + \alpha_4 FLL^0_{i,T} + \alpha_5 FLL^{1+}_{i,T} + \sum_l \alpha_l Control_l + \gamma_T + \mu_j + \varepsilon_{j,T}, \quad (3)$$

¹⁸ Based on data from the Factiva News database collected by Dasgupta and Žaldokas (2019), we replace the year of FLL adoption with the year when the first discussion of leniency laws was started by policy makers. Untabulated results are robust to this change. We are grateful to these authors for providing us with the data.

The regression results for stock price crash risk are reported in Table 4 Panel A. We find that none of the coefficient estimates for the prior-event periods is significant across all three measures of crash risk. We also test whether the summation of the coefficient estimates for the prior-event period is significantly different from zero. Untabulated F-tests show the insignificance of the coefficient summations (p values = 0.47, 0.52, and 0.54, respectively). Collectively, these results indicate that there is no pre-existing trend in stock price crash risk prior to the passage of foreign leniency laws. In contrast, we find that the event and post-event coefficient estimates are negative and significant at the 10 percent level or less, consistent with our main findings. In addition, the estimated coefficients for the event and post-event periods are economically much larger on average than those for the pre-event periods. In sum, the evidence presented reinforces our main findings that there was a significant decrease in stock price crash risk for U.S. firms after stronger antitrust enforcement around the world.

4.2.2. Alternative Measures of *FLL*

The construction of *FLL* is based on imports from all foreign countries in each industry. As a sensitivity analysis, we recalculate the measure of foreign leniency law based on the major import countries in each industry, that is the top 5 and top 10 import countries (*FLL_TOP5* and *FLL_TOP10*), respectively. Table 4 Panel B, Models (1) and (2) present the estimation results after replacing *FLL* with *FLL_TOP5* and *FLL_TOP10* in regression Equation (2), respectively. The coefficients on *FLL_TOP5* and *FLL_TOP10* remain significant and negative at the 1 percent level across all three crash risk specifications.

4.2.3. Falsification Tests

To further support our main results, we implement a placebo test. Specifically, for each foreign country that passed a leniency law, we randomly select a pseudo-*FLL* adoption year and define a new variable as *Placebo-FLL*, based on the same procedure as in our main tests. We then

re-estimate regression Equation (2) by using *Placebo-FLL* as an independent variable. We repeat this procedure 5,000 times and plot the distribution of the placebo coefficients in Figure 1 with *NCSKEW* as the dependent variable. Figure 1 shows that the placebo coefficient largely follows a normal distribution (mean = -0.4419, standard error = 0.00144). Model (3) in Table 4, Panel B shows that, in both economic and statistical terms, the actual *FLL* coefficient is significantly greater than the mean of the placebo coefficients at less than 0.001 level. Similarly, we find a normal distributional pattern when *DUVOL* and *COUNT* are used as the dependent variables. Here too the actual *FLL* coefficient is significantly greater than the mean of the placebo coefficients at less than 0.001 level.

-----Insert Figure 1 -----

In addition, we recalculate the measure of foreign leniency laws based on the five least important countries in terms of imports (*FLL_LEAST5*). Since there exists limited trade between these industries and countries, leniency laws in these countries should have little impact on U.S. firms' bad news disclosure decisions. As shown in Table 5 Panel B, Model (4), we do not find a significant impact on stock price crash risk as expected.

4.2.4. Export-based *FLL*

In our previous tests, we follow the literature by measuring an industry's exposure to *FLL* based on import shares, since U.S. firms are likely most concerned about protecting their domestic markets. As a sensitivity test, we also measure the industry's exposure to the foreign leniency laws with respect to its export shares and re-estimate *FLL*. Table 5 Panel B, Model (5) presents the regression results after replacing *FLL* with this export-based measure of *FLL*. The coefficient on this measure loads negatively at the 1 percent level or better across all the crash risk specifications, consistent with our main findings.

4.2.5. Bad News Formation

One of the underlying assumptions of our study is that under the exogenous shock of the passage of foreign antitrust leniency legislation, U.S. firms' incentives to disclose/hoard bad news changed. However, it is possible that after the adoption of FLL, firms take on more risky strategies, which ultimately lead to a competition-related change in bad news formation. To further make sure that our main finding regarding the negative relation between the passage of foreign leniency laws and stock price crash risk is due to the change in U.S. firms' incentives to disclose/hoard bad news rather than the change in bad news formation, we re-estimate Equation (2) after controlling for a series of measures of bad news formation including loss, missing earnings forecast, investment inefficiency and analyst forecast bias. Table 4 Panel B, Model (6) provides the regression results. Despite a significant sample attrition, the coefficients on *FLL* remain negative and significant (t -statistics= -2.35, -2.04, and -1.86, respectively). Overall, our evidence indicates that, conditional on changes in bad news formation, our main finding on the relation between the passage of foreign leniency laws and stock price crash risk remains.

4.2.6. DiD Bias Correction

Recent methodological developments in the econometric theory literature (e.g., Callaway and Sant'Anna, 2020; Goodman-Bacon, 2019; de Chaisemartin and D'Haultfœuille, 2020; Sun and Abraham 2021) suggest that parameter estimates from a conventional staggered DiD approach are potentially biased when treatment heterogeneities across cohort groups or over time are present. Baker, Larcker, and Wang (2021) summarize this literature and propose alternative DiD methods applicable to finance, accounting and law issues. We follow the suggestion by Baker et al. (2021) to correct for potential DiD estimation bias by focusing on a set of major passages of foreign leniency programs specific to each U.S. industry in our sample. Specifically, for each industry, we choose the most important country (i.e., a country having the largest import volume in that industry into the U.S.) over our sample period so that each industry is treated only once.

We then create a binary variable of foreign leniency laws equal to one if the country specific to that industry has adopted foreign leniency laws, and zero otherwise. In the empirical analysis, we weight the binary variable by the average proportion of the industry's imports from that country.

Appendix A shows significant variation in treatment timing across cohorts, suggesting potential biases in the TWFE staggered DiD estimators. Following Baker et al.'s (2021) recommendation, we implement the stacked regressions with and without time-varying controls as alternative approaches to the TWFE staggered DiD estimation.¹⁹ Specifically, we first stack cohort-specific datasets that include all firms in U.S. industries that experienced a major passage of foreign leniency laws in a treatment cohort year as well as all other not-yet-treated firm-year observations as controls. We keep only observations within the estimation window (4 years before and after the treatment event) and fit the stacked regression specification on the dataset, controlling for stack-specific firm fixed-effects and stack-specific year fixed-effects.²⁰ Models (7) and (8) of Table 4 report the treatment effect estimate using the stacked regressions without and with the inclusion of control covariates, respectively. The results show that the treatment effect estimates are similar in the magnitude to those of our main findings, statistically significant at the 5 percent level or less across all the crash risk models.

-----Insert Table 4 -----

4.3. Cross-sectional Analyses: Agency Conflicts

¹⁹ Baker et al. (2021) also proposed two alternative approaches, Callaway and Sant'Anna (2021) and Sun and Abraham (2021). However, these two approaches are not appropriate in our setting. See the detailed discussions in Baker et al. (2021), Callaway and Sant'Anna (2021) and Sun and Abraham (2021).

²⁰ Alternatively, it is possible to fit the complete set of time periods. However, the length of time in the event window included in the analysis can affect the estimation of aggregate treatment effects. The downside of an excessively long event window is that event-study estimates under examination could be contaminated by the causal effects of other potentially omitted events during the sample period. The results are qualitatively similar if we try 3-year and 5-year event windows.

In this section, we seek to identify and understand the underlying economic factors that lead to cross-sectional differences in crash risk arising out of *ex ante* competitive threats from rivals. The hypothesized relation between *ex ante* competitive threats from rivals and firm-specific crash risk is based conceptually on agency conflicts between managers and shareholders that ultimately lead to managerial bad news hoarding behavior (Jensen and Meckling 1976). Firms with few agency conflicts between managers and shareholders are likely to be well-governed (Jensen and Meckling, 1976) and disclose bad news on a timely basis irrespective of the passage of foreign leniency laws. As a result, the exogenous FLL shock will have little or no impact on their disclosure strategy. By contrast, the exogenous passage of foreign leniency laws is more likely to motivate poorly governed firms, those with more agency problems, to engage in bad news disclosure to deter potential competition and avoid being forced out of business. In fact, being poorly governed, these firms will tend to have more scope for bad news disclosure arising out of past bad news hoarding. Poorly governed firms will also have more incentive to disclose bad news to deter potential competition because they are more likely to perform relatively poorly with intensified competition. In other words, when facing the exogeneous shock of FLL, poorly governed firms are more likely to re-weigh the cost and benefits of bad news hoarding/ disclosure and leverage their discretion by strategically disclosing bad news to deter potential competition. Thus, the role that *ex-ante* competitive threats from FLL play in mitigating stock price crash risk should be more salient for poorly governed firms. In this study, we focus on three different aspects of agency conflicts: 1) corporate governance mechanisms; 2) risk-taking behavior; and 3) corporate information asymmetry.

First, we use a series of direct governance measures including board independence (*NON_IND_BD*), board tenure (*TENURE_BD*), CFO equity incentives (*CFO_INCENTIVE*) and the entrenchment index (*E_INDEX*) of Bebchuk, Cohen, and Ferrell (2009) to proxy for corporate

governance mechanisms. Andreou, Antoniou, Horton and Louca (2016) suggest that internal corporate governance plays a material role in mitigating crash risk. The literature indicates that the higher *IND_BD* and the lower *TENURE_BD*, the stronger is corporate governance (e.g., Anderson, Mansi, and Reeb 2004).²¹ Kim, Li and Zhang (2011b) provide strong evidence that the sensitivity of a chief financial officer's (CFO) equity incentives to stock price is significantly and positively related to the firm's future stock price crash risk. In contrast, they find only weak evidence for the chief executive officer (CEO)'s equity incentives. Here, we follow Bergstresser and Philippon (2006) and Kim, Li and Zhang (2011b) to measure *CFO_INCENTIVE* based on the dollar change in the value of a CFO's equity holdings that would come from a one-percentage-point increase in the stock price. Following prior literature, we also use *E_INDEX* to measure cross-sectional variation in managerial power versus investor protection. The *E_INDEX* is based on six out of 24 corporate governance provisions of Gompers, Ishii, and Metrick (2003) that are followed by the Investor Responsibility Research Center (IRRC), including staggered boards, limits to shareholder by-law amendments, poison pills, golden parachutes, and supermajority requirements for mergers and charter amendments.²² The lower *E_INDEX* values correspond to a lower level of managerial power and a higher level of investor protection.

Second, Callen and Fang (2015a) argue that managers of firms with high levels of risk-taking are more likely to conceal bad news because bad news may be perceived by investors to be the realization of excessive risk-taking behavior by managers. Excessive risk-taking behavior

²¹ We obtain similar results when we replace board independence by board cooption as in Coles, Daniel and Naveen (2014) and Cassell, Myers, Schmardebeck and Zhou (2018).

²² Bebchuk, Cohen and Ferrell (2009) document that the rise in the level of the *E_INDEX* is associated with a significant decrease in firm value and large negative abnormal returns. However, no such findings obtain for the rest of the eighteen provisions. Hence, the *E_INDEX* represents a less noisy metric of the corporate governance index originally developed by Gompers, Ishii and Metrick (2003). Further, relative to single-dimension measures of corporate governance, the *E_INDEX* tends to reflect the multiple aspects of corporate governance, thereby minimizing false inferences.

may exacerbate agency problems in the firm by prompting managers to hoard bad news information. We expect that competitive threats from existing rivals should play a greater role in preempting stock price crash risk for firms with higher risk-taking levels. Empirically, we use volatility of firm-specific weekly returns (*SIGMA*), volatility of cash flow from operations (*CVOLA*), volatility of net income (*EVOLA*) and the leverage ratio (*LEV*) to proxy for the risk-taking level of the firm.

Lastly, the literature suggests that firms with poor information quality, including opaque financial reporting, are likely to conceal bad news and experience stock price crashes (e.g., Jin and Myers, 2006; Hutton et al., 2009; Zhu 2016). Thus, we conjecture that the relation between competitive threats from existing rivals and firm-specific crash risk is stronger for firms with more opaque financial reporting. We employ Hutton, Marcus, and Tehranian (2009)'s *OPAQUE* to proxy for financial reporting opacity. In addition, we use the discretionary accruals based on the modified Dechow and Dichev model in McNichols (2002) (*OPAQUE2*) and the inverse value of the conservatism score in Kim, and Zhang (2016) (*OPAQUE3*) as alternative measures of financial reporting opacity.²³ A higher value of these measures reflects more information asymmetry between managers and outsiders.

In the cross-sectional analysis, we re-estimate our main regression equation (2) by including interaction-terms between *FLL* and *NON_IND_BD*, *TENURE_BD*, *CFO_INCENTIVE*, *E_INDEX*, *SIGMA*, *CVOLA*, *EVOLA*, *LEV*, *OPAQUE*, *OPAQUE2*, and *OPAQUE3*, respectively. Table 5 presents our estimation results for the key interaction terms only. We find that the sign of the coefficient estimates for all interaction terms are consistent with expectations across all the three crash risk specifications. In 30 out of 33 estimations, the coefficients are significant at the 10

²³ We also utilize a third measure of financial reporting opacity based on the modified Jones model in Dechow, Sloan, and Sweeney (1995). The results remain robust.

percent level or lower. Overall, these findings indicate that the impact of competitive threats arising from antitrust leniency programs on stock price crash risk is more pronounced for firms with more severe agency problems including weaker corporate governance, higher level of risk-taking and more opaque financial reporting.

-----Insert Table 5 -----

4.4. Benefits of Timely Bad News Disclosure

Given that the incumbents perceive the passage of FLL as a potential competitive threat (although the threat has not yet materialized as intensified product competition), they will respond strategically to deter potential entry by disclosing bad news on a timely basis to maintain their current competitive advantage in the product market. Thus, in equilibrium, if the incumbents' disclosure strategy is successful, we should expect material benefits will accrue to them after the adoption of FLL. Here, we focus on the three main aspects regarding the benefits to the incumbents: 1) competitive threats; and 2) assets pricing; and 3) U.S. imports.

4.4.1. Competitive Threats

One direct benefit arising from the timely disclosure of bad news is to reduce future competitive pressures in the product market after the adoption of FLL. That is, if timely disclosure of bad news is effective, incumbent firms should face less competitive threats after FLL.

To confirm this conjecture, we estimate an empirical model following Li, Lundholm and Minnis (2013) and measure management's perceptions of competitive threats based on textual analysis of the firm's 10-K filing. Specifically, we calculate the number of occurrences of competition-related words over total words in the 10-K filing (*PCTCOMP*) to reflect management's perceptions of the intensity of the competition they face. This measure is forward-looking and incorporates the dynamic actions of the firm's existing and potential rivals (Li et al.

2013).²⁴ In addition, we follow prior literature and use Hoberg, Phillips, and Prabhala (2014)'s measure as an alternative measure of competitive threats. In line with theories of contestable markets in industrial organization (i.e., Baumol, Panzar, and Willig 1982), they develop a new measure of product market threats, namely, fluidity (*FLUIDITY*), using firms' business descriptions provided in 10-Ks. Intuitively, this measure captures the change in a firm's product space due to moves made by competitors, including potential rivals, in the firm's product markets.

We regress each of *PCTCOMP* and *FLUIDITY* on *FLL* and control variables in Equation (2) including year and firm fixed effects. Table 6 provides the regression results. Consistent with our conjecture, we find that the *FLL* coefficients are significant and negative (*t*-statistics= -1.93 and -2.30, respectively). Overall, the evidence is in line with our conjecture that timely disclosure of bad news is effective in reducing potential competitive threats incumbent firms face after the adoption of FLL.

-----Insert Table 6 -----

4.4.2. Asset Pricing (Cost of Capital) Implications

Economic theory suggests that increased information supply reduces information asymmetries between managers and outsiders, leading to reduced costs of capital (e.g., Akerlof 1970; Glosten and Milgrom 1985; Amihud and Mendelson 1986; Merton 1987; Diamond and Verrecchia 1991; Easley, Hvidkjaer, and O'Hara 2002; Easley and O'Hara 2004). In addition, another stream of literature indicates that tail risk is a predictor of expected equity returns (e.g., Kelly and Jiang 2014; Bali Cakici, and Whitelaw 2014).

²⁴ We are indebted to Professor Michael Minnis, who graciously shared the data from 1995-2009 with us.

Because our findings indicate that *ex ante* competitive threats from potential rivals motivate managers to disclose bad news, we conjecture that antitrust leniency laws reduce costs of equity capital through their effect on stock price crash risk. To test this conjecture, we estimate a structural equation model (SEM) that decomposes the relation between *FLL* and cost of equity capital into a direct effect, and an indirect effect as mediated by stock price crash risk (Baron and Kenny 1986). Our cost of equity capital is measured based on Claus and Thomas (2001) (r_{CT}) and Gebhardt, Lee, and Swaminathan (2001) (r_{GLS}). Appendix C provides the detailed information on the estimation of the cost of equity capital.

Following prior literature using path analysis (e.g., Pevzner, Xie, and Xin 2015; DeFond, Lim, and Zang 2016; Bentley-Goode, Omer and Twedt 2019), we regress each measure of the cost of equity capital on *FLL* and on the mediating variable of stock price crash risk. In addition, we regress the mediating variable of stock price crash risk on *FLL*. All regression equations include the control variables from Section 3.3 above. Figure 2 shows the detailed paths with cost of equity capital as the ultimate outcome variable.

-----Insert Figure 2-----

Table 7, Panel A presents the results from the path analysis with the cost of equity capital measured by r_{CT} .²⁵ In Column (1), the direct path coefficient (β_1) between *FLL* and r_{CT} , controlling for stock price crash risk measured by *NCSKEW*, is insignificant and negative. The indirect path coefficient (γ_1) between *FLL* and *NCSKEW* is significantly negative (t -statistic = -2.09), reinforcing our main finding that firms facing more competitive threats from potential

²⁵ To facilitate comparison of the coefficients, path analysis standardizes all variables in the model with a zero mean and a standard deviation of one. To address saturation from the mix of observed variances and covariances that result from inclusion of firm fixed effects, the models are re-estimated with industry fixed effects. The coefficient estimates for the industry and firm fixed effects regressions are highly similar.

rivals are less likely to suppress bad news for an extended period, and thus have lower levels of future stock price crash risk. The path coefficient (β_2) between *NCSKEW* and the cost of equity capital is significant and positive (t -statistic = 4.69), consistent with bad news hoarding by insiders leading to increased costs of equity capital. The total mediated path (i.e., indirect path) for *NCSKEW* [$(\gamma_1) * (\beta_2)$] is significant and negative (t -statistic = -1.91), indicating that antitrust leniency program results in the reduced cost of equity capital via the mitigating effect on tail risk.²⁶ The results in Columns (2) and (3), Panel A using *DUVOL* and *COUNT* to measure stock price crash risk, respectively, are similar. Table 7, Panel B, provides the estimation from the path analysis with the cost of equity capital measured by r_{GLS} . The results are similar. Collectively, the evidence from the path analysis suggests that competitive threats from potential rivals reduce firms' cost of equity capital through the channel of stock price crash risk.

-----Insert Table 7-----

4.6. U.S. Import Implications

If the incumbents' strategy of timely bad news disclosure is successful in preempting potential entry, we should expect there to be no significant import growth subsequent to the adoption of FLL (after removing the pre-FLL import trend).

Figure 3 shows aggregate U.S. import volumes three years before and three years after the adoption of FLL in year 0, with the X-axis labelled year -3, year -2, year -1, year 0, year +1, year +2, and year +3. The linear forecast trend line is shown red dots based on U.S. import volumes from the pre-FLL years (i.e., year -3, year -2, and year -1) in the red bar. This figure indicates that U.S. import volumes after the adoption of FLL in the blue bar are below the linear forecast trendline (based on the pre-FLL import information), suggesting that U.S. incumbents' disclosure

²⁶ The significance of the indirect effect is estimated using the Sobel (1982) test statistic.

strategy deters the growth of U.S. imports following the adoption of FLL. In Figure 1 of the online appendix, we show similar patterns for the majority of the top 10 U.S trading partners based on their imports into the U.S. market over the sample period (including Canada, P.R. China, Mexico, Japan, Germany, United Kingdom, China Taiwan, France, and Italy, respectively).²⁷ Overall, we do not find significant growth in U.S. imports after the adoption of FLL (after removing the pre-FLL import trend).²⁸

-----Insert Figure 3-----

5. Conclusion

This study investigates the impact of *ex ante* competitive threats from potential rivals on stock price crash risk. We document that after the adoption around the world of antitrust leniency laws, U.S. firms exhibit lower levels of stock price crash risk. This is consistent with the competitive threat hypothesis that firms in industries that face threats of future entry from potential rivals disclose bad news in a timely manner to deter entry rather than suppress it, thereby mitigating managerial bad news hoarding. In equilibrium, the impact of potential rivals on bad news disclosure offsets the well documented evidence in the literature of managerial incentives to hoard bad news due to short-term career and compensation concerns.

This study further evaluates whether this relation intensifies for firms that experience more severe agency conflicts. The cross-sectional analyses show that the negative impact of competitive threats from potential rivals on stock price crash risk varies with external market

²⁷ The deterrence effect occurs in year 1 or year 2 for most of these countries. In case of Japan and United Kingdom, it occurs in year 3. We do not observe a similar impact for Korea.

²⁸ This evidence also helps distinguish our setting from the import penetration data used in the prior literature. While tariff shocks and currency exchange rate movements used in the prior literature have a direct and immediate impact on existing competition and U.S. imports, we maintain that, as a result of bad news disclosure strategy by U.S. incumbents to counter potential entry, our FLL setting does not affect U.S. imports in a similar manner.

forces, corporate risk-taking behavior and information asymmetry. This evidence implies that competitive threats from potential rivals constrain firms' stock price crash risk, particularly for those firms subject to agency conflicts whose investors are more vulnerable to bad news hoarding. Overall, our study suggests that *ex ante* competitive threats from potential rivals play an important governance role in protecting investor interests. We especially echo Giroud and Mueller (2010)'s call for more policy efforts aimed at improving industries' competitiveness such as antitrust laws.

While our evidence on import growth after the adoption of FLL suggests that U.S. incumbents' disclosure strategy is successful in preempting entry, it is possible that such a strategy also affects other competitive aspects of the product market regarding potential foreign rivals, including investments, production and financing. These factors are worthy of academic attention, especially if they have a material positive impact on the welfare of investors in capital markets. We leave this issue for future research.

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Appendix A. Foreign Country Leniency Laws

This appendix lists the passage of leniency laws by non-U.S. countries and year. The information is based on Cartel Regulation 2013, press releases and news article. Source: Dasgupta and Žaldokas (2019).

Country	Leniency Law Adoption Year
Australia	2003
Austria	2006
Belgium	2004
Brazil	2000
Bulgaria	2003
Canada	2000
Chile	2009
China	2008
Colombia	2009
Croatia	2010
Cyprus	2011
Czech Republic	2001
Denmark	2007
Ecuador	2011
Estonia	2002
Finland	2004
France	2001
Germany	2000
Greece	2006
Hungary	2003
Iceland	2005
India	2009
Ireland	2001
Israel	2005
Italy	2007
Japan	2005
Korea	1997
Latvia	2004
Lithuania	2008
Luxembourg	2004
Malaysia	2010
Mexico	2006
Netherlands	2002
New Zealand	2004
Norway	2005
Pakistan	2007
Peru	2005
Philippines	2009
Poland	2004
Portugal	2006
Romania	2004
Russia	2007
Singapore	2006
Slovakia	2001
Slovenia	2010

South Africa	2004
Spain	2008
Sweden	2002
Switzerland	2004
Taiwan	2012
Turkey	2009
Ukraine	2012
United Kingdom	1998

Appendix B. Variable Definitions

Stock price crash risk measures:

To calculate firm-specific measures of stock price crash risk, we first estimate firm-specific weekly returns from the following expanded market and industry index model regression for each firm and year (Hutton, Marcus, and Tehranian, 2009):

$$r_{j,t} = \alpha_j + \beta_{1,j}r_{m,t-1} + \beta_{2,j}r_{i,t-1} + \beta_{3,j}r_{m,t} + \beta_{4,j}r_{i,t} + \beta_{5,j}r_{m,t+1} + \beta_{6,j}r_{i,t+1} + \varepsilon_{j,t}, \quad (\text{A.1})$$

where $r_{j,t}$ is the return on stock j in week t , $r_{m,t}$ is the return on the CRSP value-weighted market index in week t , and $r_{i,t}$ is the return on the value-weighted industry index based on the two-digit SIC code. We correct for non-synchronous trading by including the lead and lag terms for the value-weighted market and industry indices (Dimson, 1979).

We define the firm-specific weekly return, $R_{j,t}$, as the natural log of one plus the residual return from Equation (A.1). We log transform the raw residual returns to reduce the positive skew in the return distribution and help ensure symmetry (Chen, Hong, and Stein, 2001). We also estimate these measures of crash risk based on raw residual returns and obtain robust (untabulated) results.

The first firm-specific measure of stock price crash risk is the negative coefficient of skewness of firm-specific weekly returns (*NCSKEW*). It is the negative of the third moment of each stock's firm-specific weekly returns, divided by the cubed standard deviation. Thus, for any stock j over the fiscal year T ,

$$NCSKEW_{j,T} = - \left(n(n-1)^{\frac{3}{2}} \sum R_{j,t}^3 \right) / \left((n-1)(n-2) \left(\sum R_{j,t}^2 \right)^{\frac{3}{2}} \right) \quad (\text{A.2})$$

where n is the number of observations of firm-specific weekly returns during the fiscal year T . The denominator is a normalization factor (Greene, 1993). This study adopts the convention that an increase in *NCSKEW* corresponds to a stock being more "crash prone," i.e., having a more left-skewed distribution, hence, the first minus sign in Equation (A.2).

The second measure of firm-specific crash risk is *DUVOL* calculated as follows:

$$DUVOL_{j,T} = \log \left\{ (n_u - 1) \sum_{DOWN} R_{j,t}^2 / (n_d - 1) \sum_{UP} R_{j,t}^2 \right\} \quad (\text{A.3})$$

where n_u and n_d are the number of up and down weeks over the fiscal year T , respectively. For any stock j over a one-year period, we separate all of the weeks with firm-specific weekly returns above (below) the mean of the period and call this the "up" ("down") sample. We further calculate the standard deviation for the "up" and "down" samples separately. We then compute the log ratio of the standard deviation of the "down" sample to the standard deviation of the "up" sample. Similar to *NCSKEW*, a higher value of

DUVOL corresponds to a stock being more “crash prone”. This alternative measure does not involve the third moment and hence is less likely to be excessively affected by a small number of extreme returns.

The last measure, *COUNT*, is based on the number of firm-specific weekly returns exceeding 3.09 standard deviations above and below the mean firm-specific weekly return over the fiscal year, with 3.09 chosen to generate frequencies of 0.1% in the normal distribution (Hutton et al. 2009). *COUNT* is the downside frequencies minus the upside frequencies. A higher value of *COUNT* corresponds to a higher frequency of crashes. Like Hutton et al. (2009) and Kim et al. (2011a), we use the 0.1% cut-off of the normal distribution as a convenient way of obtaining reasonable benchmarks for extreme firm-specific weekly returns.

Measurement of competitive threats

FLL is a continuous variable of Foreign Leniency Law, which is constructed as the weighted average of the passage of leniency laws in all countries other than the United States:

$$FLL_{i,T} = \sum_k w_{k,i} L_{k,T},$$

where k , i and T denote a foreign country, an industry (SIC two-digit code) and a year, respectively. Here, $w_{k,i}$ is the weight of industry i 's imports from country k . $L_{k,T}$ is equal to a value of one if country k has passed a leniency law by year T , and zero otherwise.

Other variables:

KUR is the kurtosis of firm-specific weekly returns over the fiscal year.

SIGMA is the standard deviation of firm-specific weekly returns over the fiscal year.

RET is the mean of firm-specific weekly returns over the fiscal year, multiplied by 100.

ROA is earnings before extraordinary items divided by total assets at the end of the fiscal year.

MB is the ratio of market value of equity to book value of equity at the end of the fiscal year.

LEV is the book value of all liabilities divided by total assets at the end of the fiscal year.

LNSIZE is the log value of market capitalization at the end of the fiscal year.

CVOLA is the standard deviation of cash flows from operations, deflated by the lagged total assets over the current and prior 4 years;

EVOLA is the standard deviation of earnings excluding extraordinary items and discontinued operations, deflated by the lagged total assets over the current and prior 4 years;

OPAQUE is the absolute value of annual performance-adjusted discretionary accruals developed by Kothari et al. (2005).

DTURN is the average monthly share turnover over the fiscal year minus the average monthly share turnover over the previous year, where monthly share turnover is calculated as the monthly share trading volume divided by the number of shares outstanding over the month.

TENURE is the log value of the number of consecutive years in the fiscal year that the auditor has been employed by the firm (in the case of audit firm mergers, the incumbent auditor–client relationship is considered to be the continuation of the prior auditor).

BIGN is equal to 1 if a firm is audited by a Big-4 auditor (or its predecessor), and zero otherwise.

ANA is the log value of one plus the number of analysts that issue earnings forecasts for a given firm during the fiscal year.

TRA is the percentage of equity held by transient institutional investors at the end of the fiscal year.

DED is the percentage of equity held by dedicated institutional investors at the end of the fiscal year.

QIX is the percentage of equity held by quasi-indexer institutional investors at the end of the fiscal year.

HHI is Herfindahl–Hirschman industry index.

IMPORTPENE is the ratio of the value of imports to the sum of the value of imports plus shipment value minus the value of exports at the industry level.

Appendix C. Estimation of the Cost of Equity Capital

We estimate the cost of equity capital based on Claus and Thomas (2001) (r_{CT}) and Gebhardt, Lee, and Swaminathan (2001) (r_{GLS}):

r_{CT} is calculated as follows:

$$P_0 = B_0 + \sum_{i=1}^5 \frac{feps_i - r_{CT}B_{i-1}}{(1+r_{CT})^i} + \frac{(feps_5 - r_{CT}B_4) * (1+g)}{(1+r_{CT})^5(r_{CT}-g)},$$

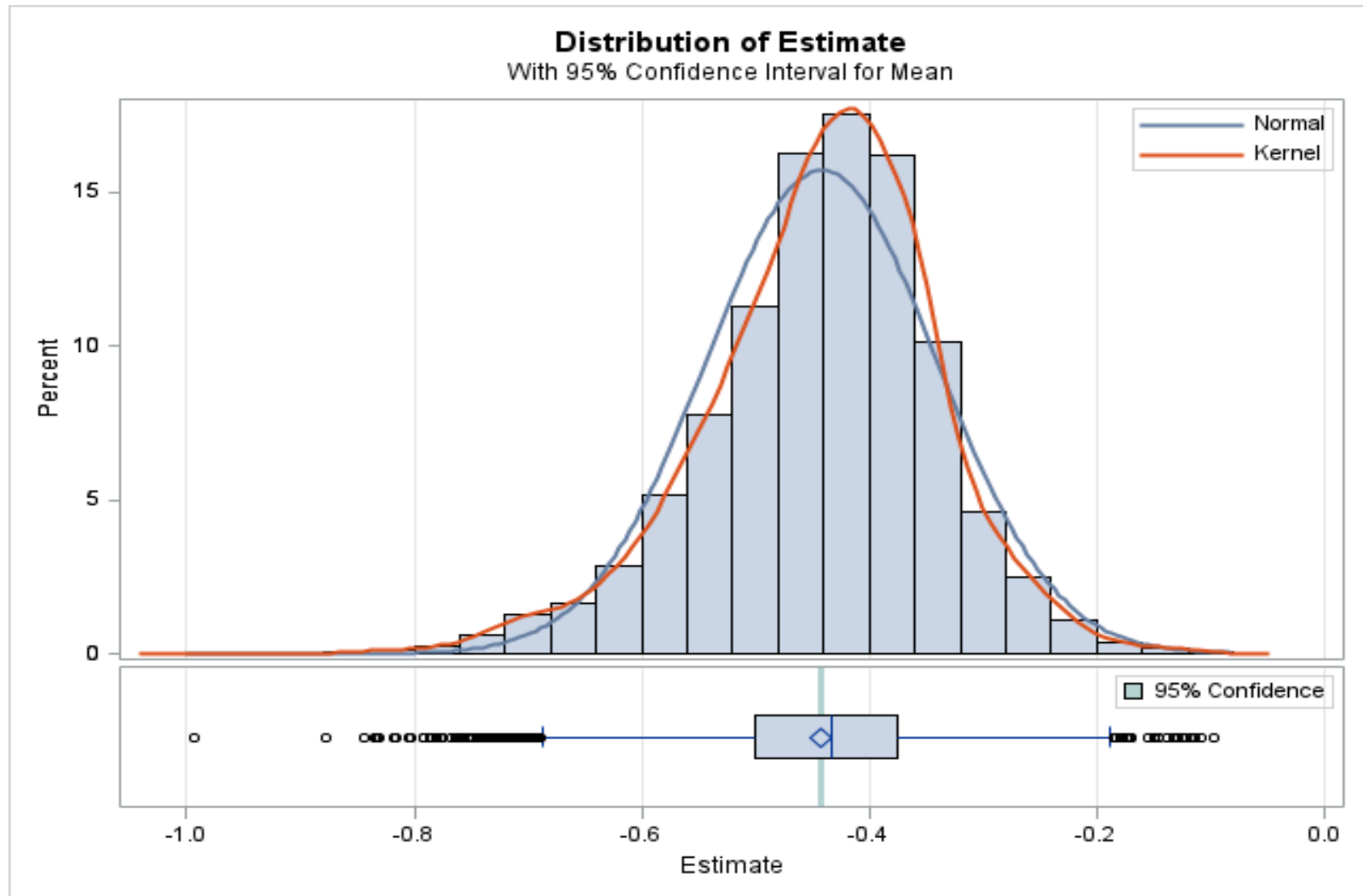
where B_0 is current book value of equity per share at the end of each year, B_t is future book value of equity per share at period t estimated using a clean surplus assumption, P_0 is current price per share at the end of each year, $feps_t$ is a t -period-ahead consensus analyst forecast of accounting earnings per share ($t=1, 2, 3, 4$ and 5) at the end of each year, and g is the growth rate of residual earnings in perpetuity equal to the expected inflation rate (10-year risk-free rate minus 3 percent) at the end of each year. Explicit forecasts of accounting earnings per share for years 4 and 5 are often unavailable. In that event, we project earnings per share for years 4 and 5 using analysts' consensus long-term growth forecast. If the long-term growth forecast is also unavailable, we assume that earnings grow at the rate of inflation beyond year 3. We use a numerical approximation program to identify r_{CT} within a .005 difference between the actual and fitted value of P_0 .

r_{GLS} is calculated as follows:

$$P_0 = B_0 + \frac{feps_1 - r_{GLS}B_0}{1+r_{GLS}} + \frac{feps_2 - r_{GLS}B_1}{(1+r_{GLS})^2} + \dots + \frac{feps_{12} - r_{GLS}B_{11}}{(1+r_{GLS})^{12}} + \frac{feps_{13} - r_{GLS}B_{12}}{r_{GLS}(1+r_{GLS})^{12}}.$$

The r_{GLS} approach uses actual book values per share and forecasted earnings per share up to three years ahead to derive future expected residual income for the first three-year period. Dividends are set equal to a constant fraction of forecasted earnings. After the first three years, we forecast $feps_4$ to $feps_{12}$ such that ROE linearly converges to industry ROE in the 12th year. Industry ROE is estimated as the five-year moving average median of past $ROEs$ of all firms in the same industry. We use a numerical approximation program to identify r_{GLS} within a 0.005 difference between the actual and fitted value of P_0 .

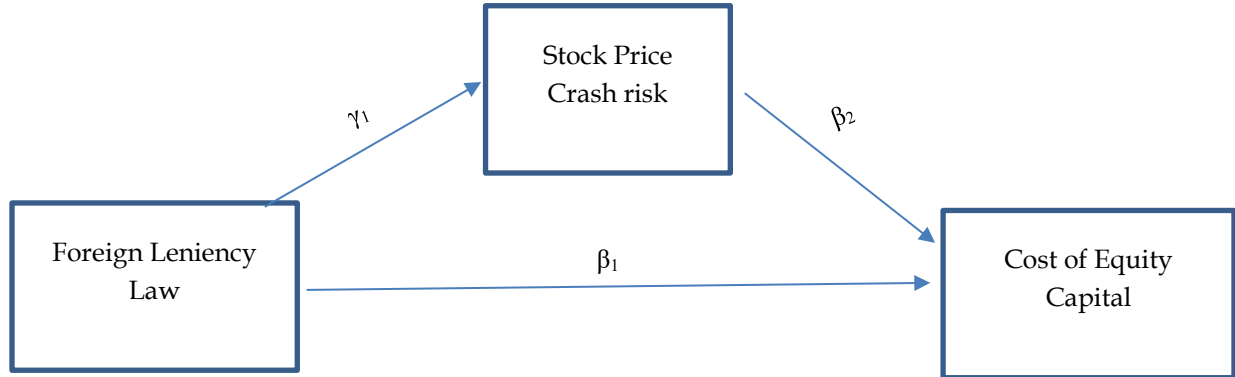
Figure 1. Distribution of *Placebo-FLL* coefficient estimates



This figure shows the distribution of placebo coefficients after using the pseudo-FLL adoption year to re-estimate regression equation (2) and repeating the same procedure for 5000 times.

Figure 2

Paths between Foreign Leniency Law and Cost of Equity Capital



This figure describes both the direct and indirect paths by which foreign leniency laws potentially affect costs of equity capital. The path analysis is conducted by utilizing a structural equation model (SEM) to estimate the system of two equations. We regress the cost of equity capital on *FLL* and on the mediating variable of stock price crash risk, and we regress the mediating variable of stock price crash risk on *FLL*. All regression equations include the control variables from Section 3.3 above. The path coefficient between foreign leniency laws and costs of equity capital (β_1) is the direct effect of foreign leniency laws on costs of equity capital, controlling for stock price crash risk. The path coefficients between foreign leniency laws and stock price crash risk (γ_1) and between stock price crash risk and costs of equity capital (β_2) represent the indirect mediating effect of stock price crash risk on the relation between foreign leniency laws and costs of equity capital. The composite coefficient $\gamma_1 \beta_2$ measures this indirect effect. A significant negative composite coefficient implies that foreign leniency laws negatively affect costs of equity capital via the mediation of stock price crash risk. All path coefficients are standardized.

Figure 3.
US Imports 3 years before and after the FLL.

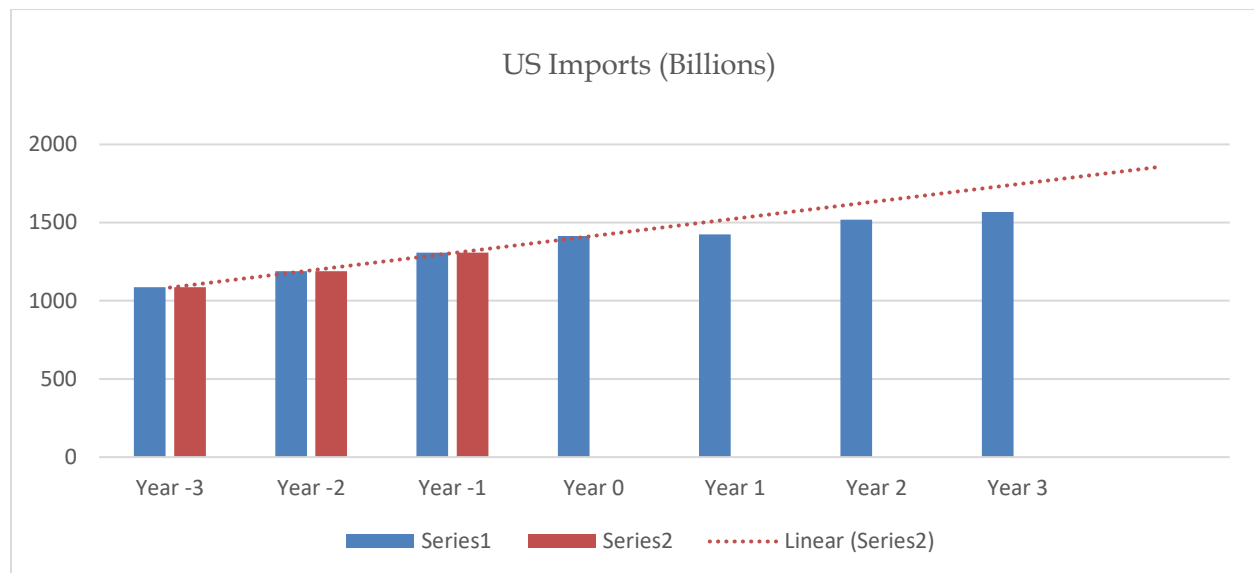


Table 1
Sample Distribution

This table presents the distribution of average FLL across sample years. The sample covers firm-year observations with non-missing values for all control variables for the period 1989 to 2013. All variables are defined in Appendix B.

Year	# observations	<i>Average FLL</i>
1989	1129	0.0000
1990	1212	0.0000
1991	1195	0.0000
1992	1284	0.0000
1993	1452	0.0000
1994	1647	0.0000
1995	1774	0.0000
1996	1904	0.0000
1997	2049	0.0123
1998	1972	0.0205
1999	1807	0.0207
2000	1652	0.0538
2001	1580	0.0567
2002	1465	0.0598
2003	1506	0.0599
2004	1542	0.0653
2005	1556	0.1045
2006	1559	0.1194
2007	1531	0.1260
2008	1404	0.1344
2009	1460	0.1353
2010	1473	0.1350
2011	1433	0.1351
2012	1418	0.1407
2013	1333	0.1397
Total	38337	

Table 2
Summary Statistics

This table presents descriptive statistics of key variables of interest for the sample of firms included in our study. The sample covers firm-year observations with non-missing values for all control variables for the period 1989 to 2013. Panel A presents descriptive statistics; Panel B presents a Pearson correlation matrix. Bold values indicate statistical significance at the 1 percent level. All variables are defined in Appendix B.

Panel A: Descriptive Statistics

	N	Mean	Standard Dev.	p5	Q1	Median	Q3	p95
<i>NCSKEW_T</i>	38337	-0.0581	0.8226	-1.2569	-0.4982	-0.0947	0.3216	1.2805
<i>DUVOL_T</i>	38337	-0.0947	0.7345	-1.2451	-0.5757	-0.1192	0.3569	1.1380
<i>COUNT_T</i>	38337	-0.0238	0.6357	-1.0000	0.0000	0.0000	0.0000	1.0000
<i>FLL_T</i>	38337	0.0591	0.1086	0.0000	0.0000	0.0271	0.0917	0.1766
<i>KUR_T</i>	38337	1.8072	2.9773	-0.4688	0.1772	0.9364	2.2929	7.0529
<i>SIGMA_T</i>	38337	0.0599	0.0336	0.0224	0.0365	0.0528	0.0755	0.1201
<i>RET_T</i>	38337	-0.0023	0.0044	-0.0071	-0.0028	-0.0014	-0.0007	-0.0002
<i>ROA_T</i>	38337	0.0690	0.2319	-0.3622	0.0503	0.1210	0.1768	0.2788
<i>MB_T</i>	38337	3.3889	4.3699	0.7031	1.3815	2.1902	3.6683	9.6784
<i>LEV_T</i>	38337	0.4399	0.2164	0.1061	0.2593	0.4443	0.6009	0.8023
<i>LNSIZE_T</i>	38337	19.9060	1.9302	16.9897	18.4847	19.7493	21.1921	23.5387
<i>CVOLA_T</i>	38337	0.1421	0.5251	0.0147	0.0338	0.0605	0.1150	0.3955
<i>EVOLA_T</i>	38337	0.1626	0.5732	0.0103	0.0268	0.0557	0.1289	0.4791
<i>OPAQUE_T</i>	38337	0.0630	0.0725	0.0035	0.0185	0.0414	0.0807	0.1920
<i>DTURN_T</i>	38337	0.0013	0.1000	-0.1519	-0.0288	0.0001	0.0303	0.1654
<i>TENURE_T</i>	38337	10.9314	8.3024	1.0000	4.0000	8.0000	16.0000	29.0000
<i>BIGN_T</i>	38337	0.8880	0.3154	0.0000	1.0000	1.0000	1.0000	1.0000
<i>ANA_T</i>	38337	1.7011	0.7805	0.6931	1.0986	1.6094	2.3026	3.0910
<i>TRA_T</i>	38337	0.1049	0.1070	0.0000	0.0089	0.0762	0.1665	0.3201
<i>DED_T</i>	38337	0.0543	0.0712	0.0000	0.0000	0.0235	0.0888	0.1956
<i>QIX_T</i>	38337	0.2667	0.2099	0.0000	0.0668	0.2502	0.4373	0.6306
<i>HHI_T</i>	38337	0.0509	0.0327	0.0215	0.0322	0.0417	0.0598	0.1037
<i>IMPORTPENE_T</i>	38337	0.3235	0.2622	0.0180	0.1299	0.2620	0.4553	1.0000

Panel B: Correlation Matrix

		1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16	17	18	19	20	21	22
<i>NCSKEW_T</i>	1																						
<i>DUVOL_T</i>	2	0.95																					
<i>COUNT_T</i>	3	0.76	0.72																				
<i>FLL_T</i>	4	0.02	0.03	0.02																			
<i>KUR_T</i>	5	0.21	0.15	0.07	0.05																		
<i>SIGMA_T</i>	6	0.01	-0.01	-0.01	-0.05	0.26																	
<i>RET_T</i>	7	0.03	0.03	0.03	0.03	-0.19	-0.60																
<i>ROA_T</i>	8	0.02	0.04	0.03	0.04	-0.11	-0.43	0.29															
<i>MB_T</i>	9	-0.04	-0.05	-0.03	-0.05	0.00	0.09	-0.06	-0.20														
<i>LEV_T</i>	10	-0.01	0.00	0.00	-0.04	-0.01	-0.16	0.08	0.11	0.16													
<i>LNSIZE_T</i>	11	0.07	0.09	0.06	0.09	-0.08	-0.49	0.26	0.30	0.16	0.19												
<i>CVOLA_T</i>	12	-0.01	-0.02	-0.02	-0.02	0.04	0.19	-0.11	-0.25	0.11	-0.09	-0.12											
<i>EVOLA_T</i>	13	-0.01	-0.02	-0.02	-0.02	0.04	0.22	-0.14	-0.31	0.14	-0.11	-0.13	0.83										
<i>OPAQUE_T</i>	14	-0.03	-0.04	-0.02	-0.05	0.02	0.23	-0.12	-0.09	0.12	-0.02	-0.19	0.11	0.10									
<i>DTURN_T</i>	15	0.02	0.02	0.02	-0.01	0.08	0.14	-0.11	0.04	0.08	0.02	0.06	0.00	0.00	0.07								
<i>TENURE_T</i>	16	0.02	0.03	0.02	0.05	-0.02	-0.31	0.16	0.18	-0.02	0.19	0.35	-0.13	-0.12	-0.15	0.02							
<i>BIGN_T</i>	17	0.03	0.03	0.02	-0.13	-0.01	-0.10	0.05	0.06	0.02	0.12	0.21	-0.07	-0.07	-0.08	0.00	0.20						
<i>ANA_T</i>	18	0.12	0.14	0.10	0.03	-0.06	-0.34	0.19	0.23	0.09	0.14	0.7	-0.10	-0.09	-0.15	0.03	0.26	0.21					
<i>TRA_T</i>	19	0.04	0.05	0.04	0.21	0.00	-0.07	0.06	0.10	0.08	-0.04	0.24	-0.02	-0.02	-0.03	0.10	0.06	0.04	0.31				
<i>DED_T</i>	20	0.04	0.05	0.03	0.06	0.00	-0.13	0.08	0.07	0.02	0.05	0.17	-0.06	-0.06	-0.08	0.01	0.1	0.07	0.23	0.27			
<i>QIX_T</i>	21	0.09	0.11	0.07	0.21	0.00	-0.33	0.20	0.22	-0.04	0.07	0.34	-0.13	-0.12	-0.18	0.01	0.29	0.10	0.40	0.53	0.36		
<i>HHI_T</i>	22	0.00	0.01	0.01	0.33	-0.03	-0.13	0.08	0.17	-0.09	0.13	0.02	-0.07	-0.05	-0.03	0.01	0.04	-0.02	-0.03	0.00	0.01	0.06	
<i>IMPORTPENE_T</i>	23	0.02	0.02	0.02	0.3	0.00	0.05	-0.02	0.00	-0.05	-0.07	0.05	0.02	0.01	0.03	0.00	-0.05	-0.05	0.08	0.12	0.03	0.05	0.08

Table 3
Impact of Foreign Leniency Law on Crash Risk

This table shows the estimated pooled cross-sectional regression of stock price crash risk on foreign leniency laws as per Equation (2). We report *t*-statistics in parentheses, based on White standard errors corrected for industry (i.e., SIC two-digits) clustering. All models include year and firm fixed-effects. *, **, and *** indicate two-tailed statistical significance at the 10, 5, and 1 percent levels, respectively. All variables are defined in Appendix B.

	<i>NCSKEW_T</i>		<i>DUVOL_T</i>		<i>COUNT_T</i>		<i>NCSKEW_T</i>		<i>DUVOL_T</i>		<i>COUNT_T</i>	
	Model 1		Model 2		Model 3		Model 4		Model 5		Model 6	
TEST VARIABLE	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
<i>FLL_T</i>	-0.3116***	(-3.79)	-0.2537***	(-4.49)	-0.2059***	(-4.77)	-0.3626***	(-4.07)	-0.3129***	(-4.94)	-0.2541***	(-4.26)
CONTROL VARIABLES												
<i>KUR_T</i>							0.0630***	(4.07)	0.0420***	(4.40)	0.0171***	(4.17)
<i>SIGMA_T</i>							0.1304	(0.20)	0.0789	(0.22)	0.0925	(0.28)
<i>RET_T</i>							0.1122***	(3.17)	0.0767***	(3.79)	0.0583**	(2.08)
<i>ROA_T</i>							-0.0440	(-1.01)	-0.0326	(-0.94)	-0.0365	(-1.57)
<i>MB_T</i>							-0.0081***	(-8.78)	-0.0070***	(-7.94)	-0.0054***	(-8.91)
<i>LEV_T</i>							-0.0711**	(-1.97)	-0.0850**	(-2.21)	-0.0364	(-1.42)
<i>LNSIZE_T</i>							-0.0851***	(-4.30)	-0.0811***	(-6.49)	-0.0392***	(-3.42)
<i>CVOLA_T</i>							0.0001	(0.01)	-0.0019	(-0.13)	-0.0004	(-0.03)
<i>EVOLA_T</i>							-0.0028	(-0.17)	-0.0039	(-0.22)	-0.0066	(-0.41)
<i>OPAQUE_T</i>							-0.0357	(-0.52)	-0.0721	(-1.31)	-0.0528	(-1.06)
<i>DTURN_T</i>							0.1356**	(2.40)	0.1349***	(2.90)	0.1404***	(4.65)
<i>TENURE_T</i>							-0.0009	(-1.00)	-0.0004	(-0.42)	-0.0011	(-1.27)
<i>BIGN_T</i>							0.0281	(1.00)	0.0285	(1.27)	0.0043	(0.23)
<i>ANA_T</i>							0.2518***	(11.80)	0.2381***	(15.97)	0.1379***	(9.88)
<i>TRA_T</i>							-0.2537***	(-2.59)	-0.2259**	(-2.53)	-0.1581*	(-1.93)
<i>DED_T</i>							0.0346	(0.45)	0.0460	(0.57)	0.0322	(0.53)
<i>QIX_T</i>							0.3829***	(9.63)	0.3844***	(10.40)	0.1980***	(6.56)
<i>HHI_T</i>							-0.0569	(-0.24)	0.0520	(0.24)	0.3120*	(1.82)
<i>IMPORTPENE_T</i>							-0.1179	(-1.40)	-0.1148	(-1.36)	0.0044	(0.07)
<i>INTERCEPT</i>	-0.0083	(-0.23)	-0.0351	(-1.18)	0.0099	(0.47)	1.1276***	(3.32)	1.0775***	(5.19)	0.5046**	(2.54)

<i>Year Fixed Effects</i>	YES	YES	YES	YES	YES	YES
<i>Firm Fixed Effects</i>	YES	YES	YES	YES	YES	YES
N	38337	38337	38337	38337	38337	38337
Adjusted R^2	0.0413	0.0442	0.0284	0.1088	0.0926	0.0444

Table 4
Robustness Checks

This table re-estimates Equation (2), the pooled cross-sectional regression of future stock price crash risk on foreign leniency laws, for a series of robustness checks. Results are presented separately for each of our measures of stock price crash risk. In Panel A, we implement the temporal dynamics test (i.e., pre-trend analysis) to examine the dynamics of stock price crash risk surrounding the *FLL* events. In Panel B, Models (1) and (2) re-specify the measure of foreign leniency laws with the modified measures based on the top 5 and top 10 import countries in each industry (*FLL_TOP5* and *FLL_TOP10*), respectively. Model (3) provides a falsification test based on Placebo coefficient. Specifically, for each foreign country that passed the leniency law, we randomly select a pseudo adoption year of FLL and define a new variable as *Placebo-FLL*. We then re-estimate regression Equation (2) to obtain a placebo coefficient. We repeat this procedure 5,000 times. Model (4) presents the results based on *FLL* calculated from the five least important countries in terms of imports (*FLL_LEAST5*100*). Model (5) uses an export-based FLL measure (*FLL_EXP*). Model (6) provides the results after further controlling for bad news formation. Models (7) and (8) report the treatment effect results based on the stacked regression approach in Baker, Larcker and Wang (2021), without and with the inclusion of covariate controls, respectively. To economize on space, control variables (see Table 3) are suppressed. We report *t*-statistics in parentheses, based on White standard errors corrected for industry (i.e., SIC two-digits) clustering. All models include year and firm fixed-effects. *, **, and *** indicate two-tailed statistical significance at the 10, 5, and 1 percent levels, respectively. All variables are defined in Appendix B.

Panel A. Temporal Dynamics Test

<u>Dep. Variable:</u>	<i>NCSKEW</i>		<i>DUVOL</i>		<i>COUNT</i>	
	(1)		(2)		(3)	
<u>Test Variable</u>	<u>Coeff.</u>	<u>t-stat.</u>	<u>Coeff.</u>	<u>t-stat.</u>	<u>Coeff.</u>	<u>t-stat.</u>
<i>FLL</i> ⁻³	-0.0336	(-0.17)	-0.0115	(-0.06)	-0.0439	(-0.30)
<i>FLL</i> ⁻²	-0.0970	(-0.99)	-0.0975	(-1.20)	-0.1368	(-0.82)
<i>FLL</i> ⁻¹	-0.1475	(-0.78)	-0.1392	(-0.73)	-0.0476	(-0.30)
<i>FLL</i> ⁰	-0.2783**	(-2.52)	-0.1940*	(-1.94)	-0.4436***	(-4.82)
<i>FLL</i> ¹⁺	-0.2484**	(-2.52)	-0.2217***	(-2.98)	-0.1721***	(-2.82)
Controls	Yes		Yes		Yes	
Year FEs	Yes		Yes		Yes	
Firm FEs	Yes		Yes		Yes	
N	38337		38337		38337	
Adjusted <i>R</i> ²	0.1087		0.0925		0.0445	

Panel B. Additional Analyses

<u>Dep. Variable:</u>		<i>NCSKEW</i>		<i>DUVOL</i>		<i>COUNT</i>	
		<i>FLL</i>					
<u>Additional analysis</u>	Model	<u>Coeff.</u>	<u>t-stat.</u>	<u>Coeff.</u>	<u>t-stat.</u>	<u>Coeff.</u>	<u>t-stat.</u>
Alternative measures of FLL							
FLL_TOP5	(1)	-0.4424***	(-5.50)	-0.3643***	(-7.42)	-0.3494***	(-9.32)
FLL_TOP10	(2)	-0.3998***	(-4.02)	-0.3222***	(-4.40)	-0.3134***	(-7.46)
Falsification tests							
Placebo coefficient minus Actual FLL coefficient (<i>t</i> -statistics for the test of the difference equal to zero)	(3)	-0.0793***	(-55.18)	-0.0789***	(-64.36)	-0.0390***	(-34.29)
FLL_LEAST5*100	(4)	-0.8058	(-0.57)	-1.3554	(-1.20)	-0.6917	(-0.74)
Exported-based FLL							
FLL_EXP	(5)	-0.7552***	(-3.40)	-0.7170***	(-4.13)	-0.4838***	(-2.71)
Control for bad news formation							
FLL	(6)	-0.4295***	(-2.95)	-0.3775***	(-2.57)	-0.3140**	(-1.99)
DiD Bias Correction							
Without covariates	(7)	-0.2491***	(-2.67)	-0.1943***	(-2.86)	-0.2199**	(-2.45)
With covariates	(8)	-0.3330**	(-2.48)	-0.2647***	(-3.39)	-0.2552***	(-3.48)

Table 5
Differential Impact of Monitoring on Crash Risk: Agency Conflicts

This table shows the estimated pooled cross-sectional relations between foreign leniency laws, agency conflicts, and stock price crash risk as per Equation (2). We focus on three dimensions of agency conflicts: 1) corporate governance mechanisms, including the percentage of non-independent board directors (*NON_IND_BD*), the average tenure of board directors (*TENURE_BD*), CFO equity incentive, i.e., the dollar change in the value of a CFO's equity holdings corresponding to a one-percentage-point increase in the stock price (*CFO_INCENTIVE*), and the entrenchment index (*E_INDEX*) of Bebchuk, Cohen, and Ferrell (2009); 2) risk-taking behavior, including volatility of firm-specific equity returns (*SIGMA*), volatility of cash flow from operations (*CVOLA*), volatility of net income (*EVOLA*), and corporate leverage ratio (*LEV*); and 3) information asymmetry based on *OPAQUE* from Hutton, Marcus, and Tehranian (2009), the discretionary accruals based on the modified Dechow and Dichev model in McNichols (2002) (*OPAQUE2*) and the inverse value of the conservatism score in Kim, and Zhang (2016) (*OPAQUE3*), respectively. To economize on space, all of the control variables (see Table 3) are suppressed. We report only the coefficient estimates on the key interaction terms. We report *t*-statistics in parentheses, based on White standard errors corrected for industry (i.e., SIC two-digits) clustering. All models include year and firm fixed-effects. *, **, and *** indicate two-tailed statistical significance at the 10, 5, and 1 percent levels, respectively. All variables are defined in Appendix B.

<u>TEST VARIABLE</u>	<u>Model</u>	<u>NCSKEW</u>		<u>DUVOL</u>		<u>COUNT</u>	
		<u>Model 1</u>	<u>Model 2</u>	<u>Model 1</u>	<u>Model 2</u>	<u>Model 3</u>	<u>Model 4</u>
		<u>Coeff.</u>	<u>t-stat.</u>	<u>Coeff.</u>	<u>t-stat.</u>	<u>Coeff.</u>	<u>t-stat.</u>
<i>FLL*NON_IND_BD</i>	(1)	-0.2648***	(-4.21)	-0.2588***	(-3.96)	-0.2740***	(-4.10)
<i>FLL*TENURE_BD</i>	(2)	-0.0482*	(-1.70)	-0.0215	(-0.93)	-0.0400**	(-2.01)
<i>FLL*CFO_INCENTIVE</i>	(3)	-1.4695*	(-1.68)	-1.5163*	(-1.82)	-1.0660*	(-1.71)
<i>FLL*E_INDEX</i>	(4)	-0.6900***	(-3.34)	-0.6029***	(-3.08)	-0.5354***	(-3.00)
<i>FLL*SIGMA</i>	(5)	-7.4217***	(-2.93)	-4.4615**	(-2.39)	-2.7549**	(-2.27)
<i>FLL*CVOLA</i>	(6)	-0.6680**	(-2.06)	-0.5046**	(-2.22)	-0.4789**	(-2.29)
<i>FLL*EVOLA</i>	(7)	-0.7044***	(-3.03)	-0.6179***	(-3.52)	-0.4773***	(-3.08)
<i>FLL*LEV</i>	(8)	-0.5679**	(-1.97)	-0.5198***	(-2.84)	-0.3736**	(-2.03)
<i>FLL*OPAQUE</i>	(9)	-1.1282	(-1.41)	-1.1540*	(-1.82)	-0.8741	(-1.26)
<i>FLL*OPAQUE2</i>	(10)	-0.7455***	(-3.39)	-0.6306**	(-2.82)	-0.4707*	(-1.74)
<i>FLL*OPAQUE3</i>	(11)	-1.5948*	(-1.68)	-1.6162*	(-1.84)	-1.9361***	(-2.88)

Table 6
Competitive Threats

This table shows the estimated pooled cross-sectional regression of competitive threats on foreign leniency laws as per Equation (2). We follow Li, Lundholm and Minnis (2013) and calculate the number of occurrences of competition-related words over total words in the 10-K filing (*PCTCOMP*) to reflect management's perceptions of the intensity of the competition they face. Alternatively, we employ Hoberg, Phillips, and Prabhala (2014)'s measure (*FLUIDITY*) as a second measure of competitive threat. We report *t*-statistics in parentheses, based on White standard errors corrected for industry (i.e., SIC two-digits) clustering. All models include year and firm fixed-effects. *, **, and *** indicate two-tailed statistical significance at the 10, 5, and 1 percent levels, respectively. All variables are defined in Appendix B.

	<i>PCTCOMP_T</i>		<i>FLUIDITY_T</i>	
	<u>Model 1</u>		<u>Model 2</u>	
	<u>Coeff.</u>	<u>t-stat.</u>	<u>Coeff.</u>	<u>t-stat.</u>
<u>TEST VARIABLE</u>				
<i>FLL_T</i>	-0.1465*	(-1.93)	-1.7923**	(-2.30)
<u>CONTROL VARIABLES</u>				
<i>KUR_T</i>	-0.0013	(-0.84)	-0.0025	(-1.07)
<i>SIGMA_T</i>	1.2970***	(3.03)	4.9385***	(2.69)
<i>RET_T</i>	4.6034***	(2.77)	1.7070	(0.53)
<i>ROA_T</i>	0.0753*	(1.84)	-0.6205***	(-3.62)
<i>MB_T</i>	0.0013	(1.06)	-0.0152***	(-3.51)
<i>LEV_T</i>	-0.1733***	(-3.51)	0.7164***	(2.77)
<i>LNSIZE_T</i>	-0.0002	(-0.02)	0.3357***	(10.87)
<i>CVOLA_T</i>	0.0191	(0.71)	0.0373	(0.94)
<i>EVOLA_T</i>	-0.0263	(-0.92)	0.0999	(1.17)
<i>OPAQUE_T</i>	-0.0357	(-0.98)	0.1620	(0.96)
<i>DTURN_T</i>	-0.0419	(-1.41)	-0.5148***	(-8.34)
<i>TENURE_T</i>	-0.0028***	(-2.79)	0.0015	(0.42)
<i>BIGN_T</i>	0.0642***	(3.41)	0.3239***	(2.88)
<i>ANA_T</i>	0.0083	(0.97)	0.1316***	(2.72)
<i>TRA_T</i>	-0.0154	(-0.41)	-0.3739	(-1.43)
<i>DED_T</i>	0.0712	(1.29)	-0.0712	(-0.26)
<i>QIX_T</i>	-0.0236	(-0.62)	-0.2659***	(-2.92)
<i>HHI_T</i>	0.7996*	(1.88)	-0.5669	(-0.28)
<i>IMPORTPENE_T</i>	-0.1518***	(-2.97)	1.4437	(1.37)
<i>INTERCEPT</i>	0.6693***	(3.66)	-2.8126***	(-3.71)
<i>Year Fixed Effects</i>	YES		YES	
<i>Firm Fixed Effects</i>	YES		YES	
N	12932		34220	
Adjusted R ²	0.5843		0.7790	

Table 7
Path Analysis

This table reports the results from a path analysis that examines the effect of foreign leniency laws on costs of equity capital mediated by stock price crash risk. The cost of equity capital is measured using Claus and Thomas (2001) (r_{CT}) and Gebhardt, Lee, and Swaminathan (2001) (r_{GLS}), respectively. Each path coefficient is standardized. The sample covers firm-year observations with non-missing values for all variables for the period 1989 to 2013. Panels A and B present the results based on r_{CT} and r_{GLS} , respectively. The significance of the indirect effect is estimated using the Sobel (1982) test statistic. The table reports the path coefficients of interest. *, **, and *** indicate statistical significance in two-tailed tests at the 10 percent, 5 percent, and 1 percent levels, respectively. All variables are defined in Appendix B.

Panel A. r_{CT}

Mediating Variable	<i>Stock Price Crash Risk</i>					
	<i>NCSKEW</i>		<i>DUVOL</i>		<i>COUNT</i>	
	(1)		(2)		(3)	
Direct Path (β_1)	<u>Coeff.</u> -0.0027	<u>t-stat.</u> (-0.34)	<u>Coeff.</u> -0.0029	<u>t-stat.</u> (-0.36)	<u>Coeff.</u> -0.0029	<u>t-stat.</u> (-0.36)
Mediated Path for Stock Price Crash risk (γ_1)	-0.0163**	(-2.09)	-0.0116	(-1.48)	-0.0185**	(-2.26)
(β_2)	0.0322***	(4.69)	0.0290***	(4.25)	0.0168***	(2.58)
Total Mediated Path for Crash Risk (γ_1)* (β_2)	-0.0005*	(-1.91)	-0.0003	(-1.40)	-0.0003*	(-1.70)
Controls	YES		YES		YES	
N	22168		22168		22168	

Panel B. r_{GLS}

Mediating Variable	<i>Stock Price Crash Risk</i>					
	<i>NCSKEW</i>		<i>DUVOL</i>		<i>COUNT</i>	
	(1)		(2)		(3)	
Direct Path	<u>Coeff.</u>	<u>t-stat.</u>	<u>Coeff.</u>	<u>t-stat.</u>	<u>Coeff.</u>	<u>t-stat.</u>
(β_1)	-0.0644***	(-9.18)	-0.0645***	(-9.20)	-0.0645***	(-9.20)
Mediated Path for Stock Price Crash risk						
(γ_1)	-0.0164**	(-2.23)	-0.0131*	(-1.76)	-0.0179**	(-2.36)
(β_2)	0.0276***	(4.80)	0.0269***	(4.71)	0.0201***	(3.59)
Total Mediated Path for Crash Risk						
(γ_1)* (β_2)	-0.0005**	(-2.06)	-0.0004*	(-1.65)	-0.0004**	(-1.97)
Controls	YES		YES		YES	
N	27280		27280		27280	

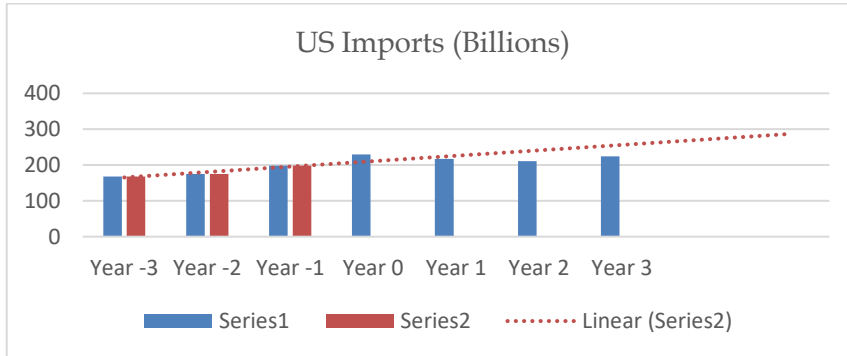
Online Appendix

***“Ex ante* Competitive Threats and Stock Price Crash Risk: Evidence from Antitrust Leniency Programs”**

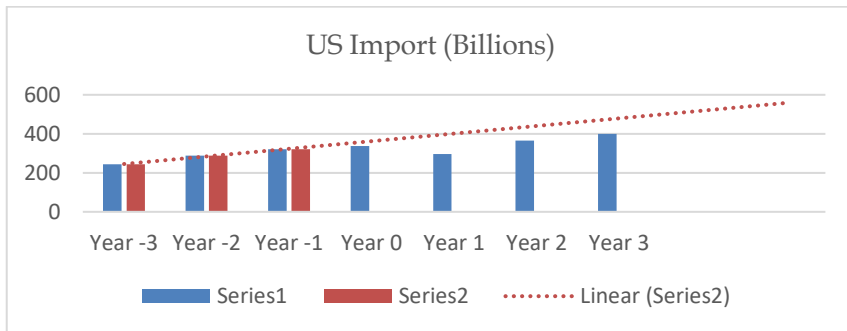
Figure OA1.

**US Imports 3 years before and after the FLL
for the TOP 10 U.S Trading Partners based on their Imports into the U.S. Market.**

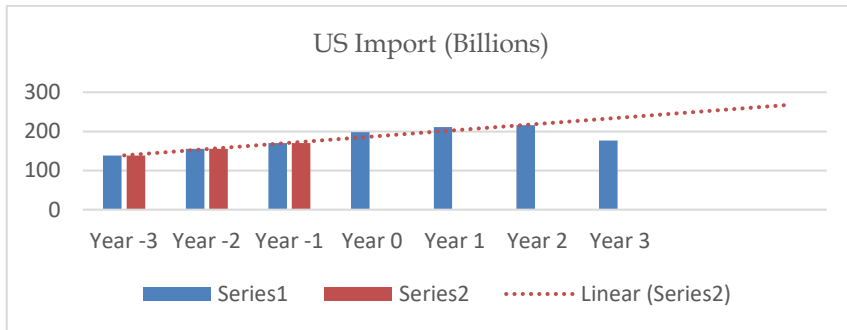
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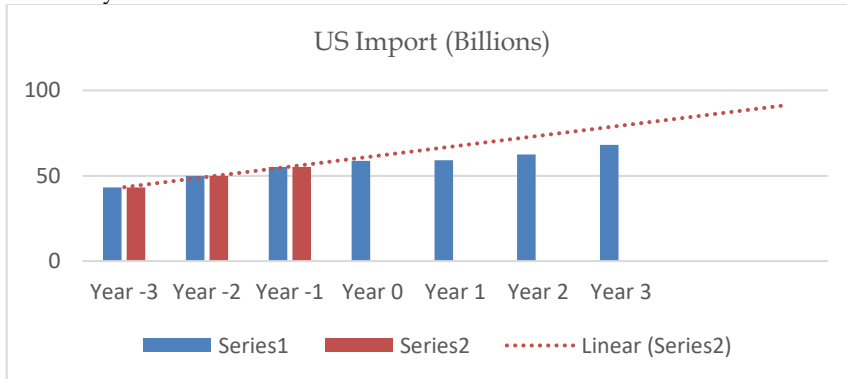
P.R. China



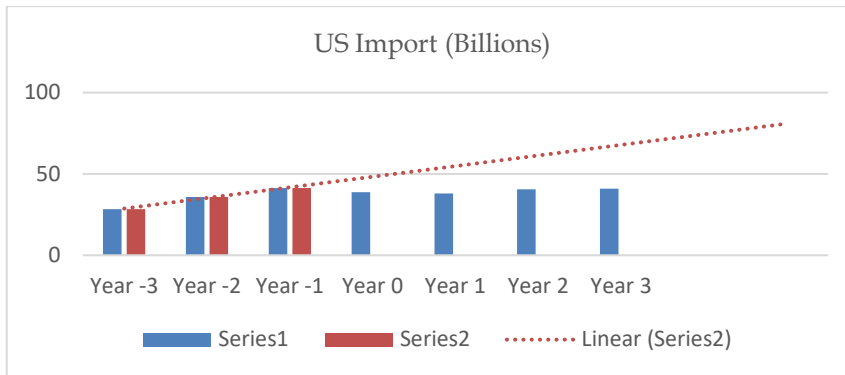
Mexico



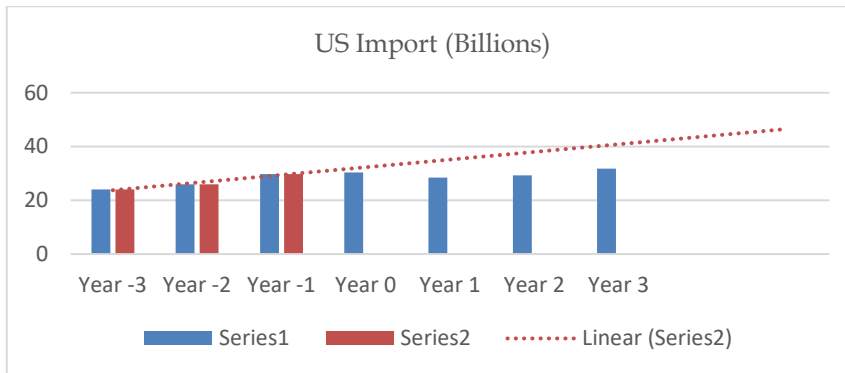
Germany



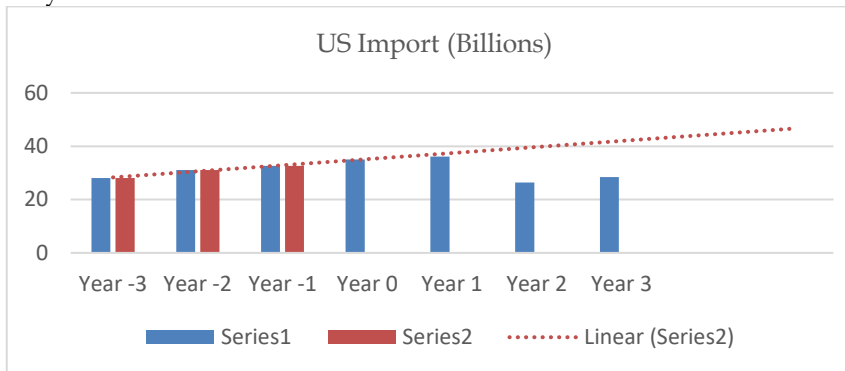
China Taiwan



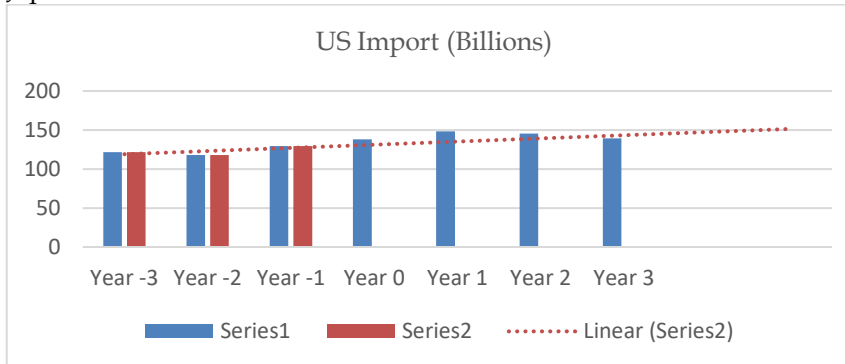
France



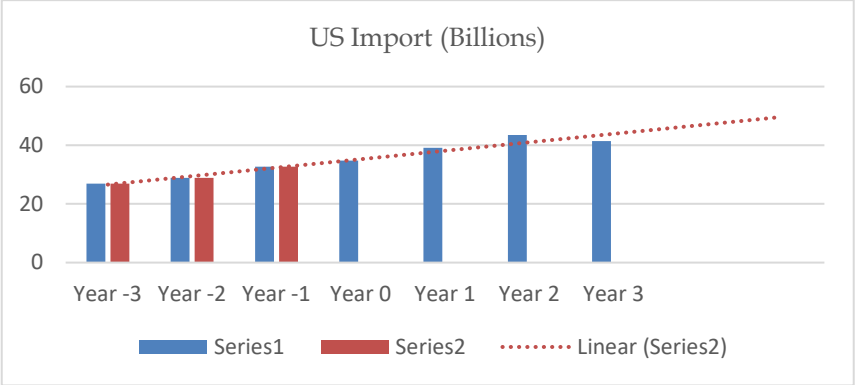
Italy



Japan



United Kingdom



Korea

