POUYAN GHAZIZADEH

Empirical Studies on the Role of Financial Information in Asset and Capital Markets

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Empirische studies naar de rol van financiële informatie in markten voor activa en financiering

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To Louisa and Liam

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Pouyan Ghazizadeh

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Chapter 1

Introduction

A primary function of capital markets is to efficiently allocate capital, which entails the flow of capital to investments with the highest returns commensurate with their risk (Tobin, 1984). The market of corporate assets fulfills a similar function, where ideally productive assets are transferred to those most equipped to manage them (Manne, 1965). As in other aspects of life however, markets seldom lack frictions, preventing the full realization of the benefits that could accrue to societies. A large body of scientific inquiry identifies and examines these frictions, and finds that the asymmetric information endowment of market participants, coupled with differential incentives, leads to increased cost of capital (Jensen and Meckling, 1976), credit rationing (Stiglitz and Weiss, 1981) or even complete market breakdowns (Akerlof, 1970). To mitigate these adverse effects, firms provide financial information to parties external to the firm. For the most part, this information is generated by the firms' accounting function and its presentation and disclosure are guided by accounting standards. The studies comprising this dissertation empirically investigate several aspects of the role financial accounting information plays in asset and capital markets.

Chapter 2[1] investigates the provision of information in the market of corporate assets. A large proportion of the transfer of corporate assets across firms consists of asset sales, where firms divest part of their operations and retain others. Prior literature shows that asset sales are used to alter the scope of the firm's activities (Maksimovic and Phillips, 2001), where assets are reallocated to those who can deploy them more efficiently (Hite et al, 1987). Furthermore, asset sales serve as a primary source of financing (Lang et al, 1995; Bates, 2005, Arnold et al, 2017), enabling selling firms to focus their attention on activities where they can add most value. While the literature finds that asset sales are value increasing to both the sellers and buyers (Eckbo and Thorburn, 2013), the process of selling assets is currently not well understood (Borisova et al, 2013). What is known is that there is a lack of sufficient public information regarding the quality of an asset, as reporting requirements regarding specific parts of firms are less stringent than those required for the firm as a whole and firms may be disinclined to voluntarily provide such detailed information given concerns regarding competition (Botosan and Stanford, 2005). This implies that potential buyers may need to incur high search costs, which in turn leads to less efficient allocation of assets. A public announcement that certain assets are available for sale can serve to reduce search costs and increase the pool of potential buyers, improving the likelihood of a more efficient allocation of assets.

Our investigation of firms' supply of information during the process of selling assets yields the following. We show that in 42% of completed asset sales the selling firm pre-announces its intention to divest, and find that these

[1] For this study, the research question and design have been developed by all co-authors jointly. Pouyan Ghazizadeh is responsible for the remainder (i.e., data collection, hypothesis development, analyses and write-up).

announcements elicit economically and statistically significant positive market reactions. Our analyses further indicate that pre-announcements are used to signal the turnaround of poor prior performance and financial distress. Furthermore, our results provide some indications that non-pre-announced asset sales involve the sale of assets that are strongly sought after, potentially initiated by foreign bidders. These seemingly different incentives for the two type of asset sales are also in line with our main finding that markets react more positively to pre-announced deals than to non-pre-announced deals. In particular, pre-announced deals imply that the future of the remaining operations will improve, which constitute the majority of operations of the selling firm. The returns to the non-pre-announced deals however seem to reflect a premium paid for the sold asset, which generally constitutes a minority part of the selling firm. Importantly, our results indicate that markets deem pre-announcements credible and that most of the valuation effects of the pre-announced asset sales are incorporated into the stock price of the selling firm prior to the deal-announcement. This, coupled with our results which suggest that the decision to pre-announce is related to the motivation for the asset sale, has implications for empirical tests of asset sales.

Chapter 3[2] investigates the effect of changes in accounting standards on the asymmetric distribution of information amongst capital market participants as inferred from changes in trading costs. Facing the prospect of trading against parties that are more informed, traders either refrain from trading or price-protect themselves, both of which prevent an optimal allocation of risk and capital. The concern of being informationally

[2] For this study, Pouyan Ghazizadeh has conducted a small part of the data collection, half of the empirical analyses and three quarters of the write-up.

disadvantaged is particularly pertinent when a firm's securities trade on more than one exchange, with one exchange being more proximate to the firm's headquarters. This is due to the fact that most value relevant information is generated there, is often communicated in the firm's home-country language, and is compiled in accordance with the firm's home-country accounting standards (Halling et al, 2007), leaving the traders on the foreign exchange at an informational disadvantage. The implementation of International Financial Reporting Standards (IFRS) by more than 120 countries, which constitutes one of the largest accounting regulatory changes to date, allows us to investigate the effect of accounting standards on the international flow of capital.

Using a sample of 239 firms with level 2 or 3 ADRs from 31 countries of which 27 have adopted IFRS between 2003 and 2012, we find that IFRS adoption improves the liquidity of ADRs, in line with the reduction of the information disadvantage of investors trading on U.S. exchanges. Our results further indicate that the improvement in the liquidity of ADRs depend on the quality of the domestic legal and regulatory institutions. Tests aimed at identifying the source of the improvements do not reveal that the superior quality of IFRS relative to the pre-existing domestic GAAPs affect the liquidity improvements, but rather point towards the scale benefits that ensue from reducing the number of standards according to which cross-listed firms report. Collectively, our results imply that the adoption of IFRS in a U.S. cross-listed firm's domestic market improves access to foreign markets which have not adopted the mandate and potentially U.S. investors' capital allocation decisions, especially for those restricted to invest in securities on U.S.

exchanges. Our findings further speak to the role of accounting standards in the competition between stock exchanges.

Finally, in chapter 4[3], I investigate whether the commitment to providing conservative accounting information disciplines the allocation of capital by managers. Extant empirical accounting research mostly focusses on the role of accounting in the supply of information for valuation and monitoring purposes, but (implicitly) regards the outcome of the underlying economic activity pursued by firms as independent of the accounting method used. More recent work endogenizes the role of accounting by recognizing that the quality of information provision by firms can improve investment efficiency by mitigating both underinvestment, through reduction of information asymmetry between firms and external suppliers of capital, and overinvestment, by facilitating contracting and monitoring (Biddle and Hilary, 2006; Biddle et al, 2009). This chapter extends this line of research by focusing on the role of accounting conservatism on investment efficiency, as despite the central role of conservatism in accounting, there are both contrasting theoretical predictions and mixed empirical evidence regarding the effect of conservatism on investment efficiency.

Exploiting the staggered and unanticipated passage of state antitakeover laws in order to circumvent endogeneity concerns, I find evidence strongly in line with a disciplinary effect of conservatism on managerial investment discretion. More specifically, I find that investors react less negatively to an increase in managerial discretion for firms that report more conservatively. Using a difference-in-difference setup, I find that firms that report more conservatively do not increase their acquisition investments, while those

[3] Pouyan Ghazizadeh has conducted the entire study.

reporting less conservatively do. Furthermore, while both the operating profitability, stock performance and riskiness of less conservatively reporting firms decline after increases in managerial discretion, more conservatively reporting firms' performance is unaffected. Overall, the evidence of this chapter suggests that accounting conservatism mitigates inefficient investment that can be attributed to increased managerial discretion.

Chapter 2

Voluntary Disclosures of Corporate Asset Sales[4]

2.1 Introduction

This chapter documents the pervasive use of voluntary disclosures used to inform the market of firms' intentions to divest part of their operations. More specifically, using a novel hand-collected dataset, we document that in over 40% of corporate asset sales, selling firms inform investors with respect to their intended transactions and that these announcements give rise to significant stock market reactions. We examine the factors that affect the selling firms' decision to voluntarily disclose their intentions, as well as the capital market reactions that such disclosures bring about.

Corporate asset sales are one of the most common ways productive assets are reallocated, making up approximately half of all M&A transactions (Maksimovic and Phillips, 2001). Prior literature shows that asset sales are used to alter the scope of the firm's activities (e.g., Maksimovic and Phillips, 2001, 2002), where assets are reallocated to firms that can deploy them more

[4] This chapter is based on Ghazizadeh, P., A. de Jong, and F. P. Schlingemann. 2018. Voluntary disclosures of corporate asset sales. *Working paper*.

efficiently (Hite et al, 1987). Furthermore, asset sales serve as a primary source of financing (Lang et al, 1995; Bates, 2005, Arnold et al, 2017), enabling selling firms to focus their attention on activities where they can add most value. In line with these potential benefits, prior literature reports positive stock market reactions when firms announce asset sales (e.g., Jain, 1985). However, in order to correctly measure and interpret the market reactions to asset sales, it is crucial that the reactions include all relevant news about the deal. As our results indicate, however, in over 40% of asset sales the current literature does not take into account selling firms' announcements about intended asset sales, which leads to the underestimation of the documented market reaction to asset sales.

In addition to correctly measuring market reactions, the analysis of the decision to pre-announce that assets are put up for sale is important for several other reasons. First, the consideration that the selling firm receives is a crucial determinant of the decision to divest. In fact, selling firms' managers may require a premium over the value of the asset under their management given their reluctance to relinquish control (Jensen, 1986). As the consideration is a function of the competitive bidding process, it is important to understand the process through which potential buyers are attracted. This process, however, is currently not well understood (Borisova et al, 2013). Potential buyers may have to incur high costs of gathering information regarding the quality of the assets put up for sale, as this information is often not publicly available. A public announcement that assets are available for sale can serve to reduce search costs and increase the pool of potential buyers, leading to a higher probability of a transaction and a higher price. Conversely, a public announcement could also lead to a poor negotiation position and a lower transaction price, because the firm reveals the poor quality of the asset and weak managerial judgement in case the asset cannot be sold. As such, the analysis of pre-announcements is not only important from a measurement perspective, but also because it is an key component of the selling process and plausibly has an effect on the occurrence and pricing of the sale. Second, market reactions to asset sales are related to the information the sale reveals with regard to the remaining operations of the firm (Brown et al, 1994; Maksimovic and Phillips, 2001). As we will discuss in more detail below, managers can use pre-announcements to signal the improved prospects of the firm's remaining operations. If the decision to pre-announce asset sales is not well understood, the examination of market reactions to asset sales motivated by factors that also affect the decision to pre-announce, may lead to incorrect inferences, because the market reactions to these pre-announcements are not taken into account.

We start our analysis by documenting the prevalence of pre-announcements. Using a sample of 330 completed asset sales between 2005 and 2015 by public parent firms incorporated in the US from non-financial and non-regulated industries, we find that 42% of asset sales are preceded by a public announcement of the intention to sell. As pre-announcements are more prevalent among larger deals (transactions preceded by an announcement are 2.8 times larger than non-announced transactions), their value-weighted proportion equals 67%. We also investigate the market's reaction to pre-announcements and find that they elicit statistically and economically significant cumulative abnormal returns, which average 2.41% over a three-day event window. These abnormal returns constitute the largest market reaction to the events related to the asset sales in our sample, which further include the reactions to the deal announcements of both pre-announced and non-pre-announced asset sales. Overall, our results indicate that excluding the market reaction to pre-announcements entails an underestimation of the market reaction to asset sales of 40%.

We then investigate the determinants of the decision to pre-announce a deal using probit analyses. The probability of a pre-announcement increases when managers have incentives to signal improved prospects of the remaining assets of the firm. In particular, we find that asset sales are more likely to be pre-announced when the selling firm's stock has performed poorly in the year preceding the announcement, and when the growth opportunities of the industry of the remaining operations have received a positive shock. We find no evidence that managers use pre-announcements to signal the quality of the assets they plan to sell. We furthermore find that asset sales by larger firms and firms more dependent on external capital are more likely to be pre-announced, as well as deals that constitute a larger proportion of the selling firm.

Next, we study the market response to the announcements and its determinants. First, we find that the overall market reaction to pre-announced deals is more positive than to those which are not pre-announced. We then explore potential reasons for this finding. The results of our tests do consistently support the notion that pre-announcements are used to signal the turnaround of poor prior performance and financial distress. Furthermore, our results provide indications that non-pre-announced asset sales involve the sale of assets that are strongly sought after, potentially initiated by foreign bidders. These seemingly different incentives for the two types of asset sales are also in line with our main finding that markets react more positively to pre-announced deals than to non-pre-announced deals. In particular, pre-announced deals bring the expectation that the future of the remaining operations will improve, which constitute the majority of operations of the selling firm. The returns to the non-pre-announced deals however seem to reflect only a premium paid for the disposed asset, which generally constitutes a minority part of the selling firm. Similar to the results of the probit analyses, we find no evidence that pre-announcements are used to signal the higher quality of the assets in play.

Our results contribute to the literature that investigates asset sales. By documenting the prevalence of pre-announcements and the capital market reactions, we shed some light on the process of asset sales, which is characterized as highly opaque (Borisova et al, 2013), despite the large operational and financial effects of these transactions for the selling firms. More specifically, we show that the current literature underestimates the market reaction to asset sales by 40% when not taking into account the reactions to the pre-announcements. Furthermore, our results suggest that the decision to pre-announce is related to the motivation for the asset sale. Thus, the omission of the market reaction to pre-announcements could lead to wrong inferences. In this regard, we add to the findings of Brown et al (1994), who report that contrary to the results for healthy firms, returns to shareholders of financially distressed firms are significantly lower when asset sales proceeds are used to repay debt than when sales proceeds are retained by the firm. They ascribe this to pressure from short-term creditors who effectively expropriate wealth from shareholders; an important finding which speaks to far-reaching effects of conflicts of interest among different type of providers of capital. However, our results suggest that an alternative reason for the lower returns could be that at least a subset of firms selling to avert financial distress has pre-announced the asset sale, which would result in the lower returns recorded when only measured at the time of the deal-announcement.

Our study also contributes to the literature on voluntary disclosures. First, several studies investigate the use of voluntary disclosures in M&A transactions (e.g., Amel-Zadeh and Meeks, 2015). Whereas the focus of these studies relates to either increases or positive biases of more general types of voluntary disclosures (i.e., earnings forecasts), we contribute by investigating a type of voluntary disclosure that is a direct part of the selling process. Second, our study contributes to the more general literature that investigates the determinants of voluntary disclosures. While the literature on voluntary disclosures is extensive, we believe that several aspects of our setting contribute to this literature. First, unlike most prior literature, we study a voluntary disclosure that is not recurring or, more specifically, sticky, which allows for a much better empirical identification. Second, the private information endowment of the manager in our setting is much less known to outsiders compared to the type of voluntary disclosures studied in the bulk of prior literature (i.e., earnings guidance). As such, the unraveling principle applies far less, rendering the disclosure in our setting much more voluntary. Third, while disclosure related costs are crucial theoretically (Verrecchia, 1983), empirically identifying disclosures that carry significant proprietary costs is challenging (Beyer et al, 2010; Lang and Sul, 2014). Our study overcomes this shortcoming as the public announcement in our setting carries clear potential costs.

The remainder of the chapter is organized as follows. Section 2 reviews prior literature and develops our hypotheses. Section 3 describes the data. Section 4 and 5 present the empirical findings, and Section 6 concludes.

2.2 Prior Literature and Hypothesis Development

Corporate asset sales are driven by both strategic and financial motives, and prior research documents that their announcements typically generate positive stock reactions (e.g., Hite et al, 1987; Borisova et al, 2013). The predominant neo-classical explanation offered for this positive reaction is twofold: asset sales enable the reallocation of assets to more efficient firms, where the seller can appropriate a fraction of the ensuing synergies through the bidding process (Hite et al, 1987), while the increase in focus leads to the improvement in the management of the remaining assets (John and Ofek, 1995). Other rationales relate the reaction to the alleviation of financial constraints (Lang et al, 1994; Bates, 2005; Arnold et al, 2017) and signals of improved governance (Mitchell and Lehn, 1990; Boot, 1992). Moreover, Maksimovic and Phillips (2001, 2002) and Yang (2008) show that asset sales coincide with industry shocks and occur more often in industries with less persistent and more volatile productivity, while others provide evidence that assets sales are reactions to corporate control and shareholder activism (Berger and Ofek, 1999).

Irrespective of the rationale, asset sales alter the scope of the firm's operations and financial structure, which in turn alter both the level and volatility of the selling firm's cash flows. Given the importance of these parameters in the valuation of a firm's stock, and the volatile circumstances that are associated with asset sales, we argue that there is a strong demand for information regarding asset sales from investors. Prior theoretical work suggests that managers have an incentive to respond to the information demands of investors and reduce their estimation risk (e.g., Lambert et al, 2007). A vast body of empirical evidence indicates that managers indeed take actions to improve their firms' information environments by voluntary disclosure of private information. For example, Shroff et al (2013) find that firms increase their voluntary disclosures prior to raising capital and that these disclosures are associated with decreased information asymmetry and costs of raising capital; Anantharaman and Zhang (2011) and Balakrishnan et al (2014) find that in response to an exogenous decrease in analyst coverage managers increase the provision of voluntary disclosures; and Billings et al (2015) find that managers react to increased volatility by providing more voluntary disclosures[5].

While these findings speak to the benefits that can be reaped from voluntary disclosures, casual observation suggests that managers do not always disclose all private information, as implied by the unraveling principle (Grossman, 1981; Milgrom, 1981). This disconnect is most commonly attributed to the existence of disclosure-related costs (Verrecchia, 1983). However, although the proprietary-cost argument is intuitively appealing, empirical identification of disclosures that carry significant proprietary costs is challenging (Beyer et al, 2010; Lang and Sul, 2014). This implies for asset sales that managers with an informational advantage over investors have an

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5 Another stream of studies investigates whether information intermediaries respond to investors' information demands. E.g., DeFond and Hung (2003) report evidence that analysts provide cash flow forecasts in addition to earnings forecasts for firms in which the two forecasts are highly complementary. More related to this study, Gilson et al (2001) show that following focus increasing break-ups more and specialized analysts start providing coverage.

incentive to disclose their intention to sell part of the firm as long as the cost of doing so is lower than the benefits of providing such disclosure. In the development of our hypotheses, we distinguish between the potential effect that a public announcement may have on the deal itself, as well as how this disclosure speaks to the performance of the remaining assets of the firm.

As mentioned previously, the comparative advantage of other firms in deploying assets provides a motive for firms to sell assets, which allows the seller to appropriate a fraction of the ensuing efficiency gains through the competitive bidding process (Hite et al, 1987). In line with this, Maksimovic and Phillips (2001) show that assets are indeed more likely to be sold when they are less productive than their industry benchmarks and that most transactions result in productivity gains. Furthermore, these authors show that asset sales are more likely when the economy is undergoing positive demand shocks, as with increasing output prices more productive firms can extract more value from the assets they control, while less productive firms incur higher opportunity costs holding on to assets they are not best equipped to manage. Yang (2008) further shows that changes in firms' productivity drive asset transfers, in particular firms with rising and falling productivity buy and sell assets, which leads to greater asset reallocation in industries in which firms have less persistent productivity.

Given that both supply and demand of assets increase with changing economic conditions and volatility of firms' productivity, firm that aim to buy assets incur high search costs in their attempts to find a suitable target. This is further exacerbated by the fact that public information regarding the quality of an asset in play is often sparsely available prior to a sale, as reporting requirements regarding specific parts of firms are less stringent than those required for the firm as a whole, and firms may be disinclined to voluntarily provide such detailed information for competitive reasons (e.g., Botosan and Stanford, 2005). As such, any potential buyer runs a risk that the quality of the asset they were planning to acquire based on publicly available information differs markedly from that based on more private information. This risk is plausibly higher during changing conditions, as the already limited public information regarding the asset is more likely to be stale.

We argue that some firms looking to appropriate part of the efficiency gains from asset sales can use a public announcement that an asset is for sale as a signal of its quality, and improve the competitive bidding process. More specifically, in line with adverse selection models (Akerlof, 1970), we propose that the informationally disadvantaged buyers pool the assets available for sale, giving the sellers of high quality assets an incentive to separate themselves from the sellers of low quality assets. Note that in our setting, a low quality asset refers to an asset for which the public valuation is higher that its valuation based on private information regarding both its current and potential productivity, while this does not hold for a high quality asset. In order to be credible, the signal needs to carry a cost, to which the sellers of high quality assets are less sensitive. A public announcement meets this requirement, given that it could lead to a poor negotiation position and lower transaction price for sellers of low quality assets. Additional costs of a public announcement include the disruptive effects of knowledge of the potential sale of part of the firm – or simply uncertainty regarding its future – on the firm's customers, suppliers or key employees (Gole and Hilger, 2008). That is, while the initial search for potential assets to acquire occurs with buyers at a large information disadvantage relative to the sellers, buyers eventually get access to private information (e.g., additional rounds of due diligence, management presentations, site visits and restricted access to the seller's data rooms). A gap between publicly available information about the asset's quality and the actual quality of the asset will lower the probability of a transaction and reduce the consideration paid for the asset. A reduced consideration will elicit a negative market response as investors will revise their valuation of the asset[6]. The failure to sell an asset put up for sale is likely to be taken as a strong negative signal regarding the quality of the asset in play. As such, sellers of high quality assets can signal their type by publicly announcing their intention to sell a certain asset. Note that this signal is received both by financial markets, resulting in a positive reaction at the time of the announcement, as well as by real markets, where potential buyers should be less concerned with the risk of expending time and effort on a futile bidding process. The latter is expected to lead to more potential buyers[7], which in turn should heighten the competitive bidding process, allowing the seller to appropriate a larger portion of the efficiency gains. Furthermore, outsiders' concerns about the gap between publicly available information on the asset's quality and the actual quality of the asset will be larger for firms with less persistent productivity. We thus propose that the benefits of a signal provided by the public announcement

[6] It is in fact likely that a reduced consideration will not only affect the market's valuation of the sold asset, but that it causes investors to revisit their priors regarding the remaining operations of the selling firm as it could cast doubt regarding the quality of the information provided by the selling firm.

[7] Note that publicly announcing that an asset is in play should in and of itself increase the number of potential buyers by simply ensuring that more potential buyers are aware of the availability of the asset. This is particularly important in the setting of asset sales, as – contrary to full mergers – asset sales do not require shareholder approval (Hege et al, 2013), which allows for substantial managerial discretion over whether and which assets to sell. Furthermore, an additional benefit of a public announcement is that boards, which in case of selling part of the firm have a fiduciary duty to obtain the highest price reasonably available, can be satisfied to have met their duties (Rosenbaum and Pearl, 2009).

should be decreasing in the extent of the persistence of the asset's productivity.

In addition argument on the appropriation of the efficiency gain described above, prior literature has ascribed some of the positive reactions to asset sales to the information that such a transaction reveals of the remaining operations of the firm. Firstly, the sale of underperforming assets rids the selling firm from the culprit to its poor performance, andalso facilitates improvements to its remaining assets. In particular, reducing the scope of the firm's activity enables firms to focus managerial attention on the remaining activities. Secondly, asset sales could reflect selling firms' improved outlooks. More specifically, Maksimovic and Phillips (2001) show that while firms sell assets from their peripheral and less efficient operations, they do so especially after a positive shock to their core and more efficient operations. Coupled with the reduction of financial constraints resulting from the consideration received as part of the transaction[8], asset sales facilitate the pursuit of more value enhancing projects. In line with both arguments, John and Ofek (1995) report that asset sales lead to improved operating performance of the remaining assets, while Dittmar and Shivdasani (2003) and Colak and Whited (2007) further show that sellers improve their investment efficiency after divestitures. As such, in addition to signaling the quality of the asset available for sale, a public announcement of an intended asset sale can also signal the quality of the selling firm's remaining operations[9]. In this case, the credibility

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[8] Asset sales serve as a key source of financing (Lang et al, 1995; Bates, 2005, Arnold et al, 2017). For instance, Arnold et al (2017, p. 1) report that *"the average proceeds from fixed asset sales correspond to roughly 44% of the average net amount of newly issued equity for U.S. manufacturing firms in COMPUSTAT between 1971 and 2010"*.

[9] Effectively, an asset sale itself could be considered a signal regarding the improved outlook of the remaining operations. This does not materially affect our analysis given that a public

of the signal is derived from the proprietary nature of such announcements: prior work shows that revealing good news regarding prospects of an industry may attract new entrants (Verrecchia, 1990; Dedman and Lennox, 2009), or elicit reaction by firms already operating there (Wagenhoger, 1990; Durnev and Mangen, 2009). Given the threshold these costs impose, the private information regarding the improved prospects must be sufficiently large to merit its public disclosure (Verreccia, 1983)[10].

Based on the above, we formulate the following hypotheses on the probability that firms pre-announce an intended asset sale:

H1: *The probability of an asset sale being pre-announced is negatively related to the persistence of the productivity of the sold assets.*

H2: *The probability of an asset sale being pre-announced is negatively related to the selling firm's past performance.*

H3: *The probability of an asset sale being pre-announced is positively related to the prospects of the selling firm's remaining operations.*

We then turn to our expectations of stock market reactions to asset sale information:

announcement can then be considered as the credible expedition of that signal – our expectation of a positive relationship between the stock market reactions and the selling firm's remaining operations would include those of non-announced asset sales, but still be more pronounced for pre-announced asset sales.

[10] The incentive to provide such a signal can be related to managers' career concerns, which will be elaborated below. Issuance of external capital may further incentivize the public disclosure.

H4: *The stock market reactions to pre-announced asset sales are more positive than to non-pre-announced asset sales.*

H5: *The stock market reactions to pre-announced asset sales are negatively related to the persistence of the productivity of the sold assets.*

H6: *The stock market reactions to pre-announced asset sales are negatively related to the selling firm's past performance.*

H7: *The stock market reactions to pre-announced asset sales are positively related to the prospects of the selling firm's remaining operations.*

2.3 Data

2.3.1 Sample Selection

We draw our sample from the Mergers and Acquisition database available from the Securities Data Corporation (SDC). We select all completed divestitures from January 1st, 2005 to December 31st, 2015 by public firms incorporated in the US, with a deal value of at least $50 million. Following previous studies (e.g., Schlingemann et al, 2002) we exclude deals of regulated utilities (SIC 4900-4999) and financial firms (SIC 6000-6999). This leads to a preliminary sample of 2409 deals. We match this sample with Compustat (annual and segment files) and the Center for Research in Security Prices (CRSP) and require that data necessary to construct the variables of interest (discussed below in more detail) is not missing. We further drop deals where the cusip code of the seller and target are the same, and require that the relative deal size (defined as the proportion of deal value to the market value of equity at the end of previous fiscal year) is at least 5% (unless the deal value is higher than $1 billion) and not larger than 90%. These steps reduce our sample to 770 deals.

We manually look up the deals in Factiva, most importantly to determine whether the selling firm has pre-announced the intention to sell the asset in question. Using the information retrieved from Factiva, we further clean the sample in the following ways: (1) we confirm that the date the deal-announcement was made public as reported in SDC, (2) we verify that the deal is an asset sale (we drop spinoffs, carve outs, asset swaps, sale and leaseback transactions, sale of real-estate, and drop-down acquisitions), (3) we confirm that the pre-announcement was made voluntarily, which entails that we drop deals that were preceded by rumors, were mandated by the FTC[11], or were part of a bankruptcy[12], (4) we drop deals that coincide with other major events other than quarterly earnings announcements (e.g., acquisitions by selling firm), and (5) we drop deals that were part of a general divestiture plan[13]. Finally, we manually link the sold asset to its reported segment using 10-K

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[11] In order to approve a merger, FTC often demands that a party to the proposed merger divests operations where the combination would otherwise gain too much market power. In these cases it is public knowledge which assets are to be divested, while the seller has not voluntary offered this information. Also, the information on the deal cannot be disentangled from the consequences of the merger that given the asset sale can follow.

[12] It is mandated by the Chapter 11 proceedings to publicly look for potential buyers, even for assets that are already pursued by potential buyers. The same arguments as above dictate the omission of these deals.

[13] This is the case when a firm announces plans to divest a certain dollar amount of asset sales, without specifying which assets will be sold. Generally, these plans involve the sale of multiple assets. Given the substantial dollar amount s that are involved, these plans generate large market reactions. Empirically, this poses a problem as the market's reaction to the sale of a certain asset cannot be disentangled from other assets that are sold as part of the same plan.

fillings available on EDGAR. This procedures leads to a final sample to 330 deals, of which 139 are pre-announced.

The annual distribution of the number of deals and total deal values of the asset sales in our sample, delineated by whether they were preceded by a public announcement of the intention to be sold (henceforth: deal type), is depicted in Figure 1. The results imply that the incidence of non-pre-announced deals is much more stable than pre-announced deals, which seem to be positively related to economic conditions.

[Insert Figure 1 here]

2.3.2 Stock Performance Measurement

We identify the public announcements of intended deals (pre-announcements) and add the event and the time that elapses until the public announcement regarding a definitive transaction agreement (deal announcement) to our analysis of asset sales. More specifically, we measure the cumulative abnormal returns (CAR) to the pre-announcement, and also the selling firm's stock performance during the time the market is aware of a possible transaction (so-called between period), which is potentially related to the CAR of the announcements of the intention and realization. Note however that this period is only available for pre-announced deals[14].

[14] To facilitate comparability, we take the mean duration of the between period of the pre-announced deals (i.e., 178 days) as the length of the in-between period for all non-pre-announced deals. As such, we also measure the performance of the selling firm prior to the asset sales at a different point in time than the extant literature.

To measure the market reaction to an asset sale, we construct the three-day CAR ([-1; +1] windows) for the selling firm around both announcements using Eventus (*PreAnn_CAR* and *Deal_CAR*, respectively). For pre-announced deals, we also sum the CAR of both events to capture the total market reaction (*Total_CAR*). In line with conventional event-study methodology, we use the market-model specification with the CRSP value-weighted index as the market portfolio, with market model parameters estimated over the window from 300 to 46 trading days prior to the event.

We further measure the compounded returns of the selling firms' stocks prior to and during the period between the pre-announcement and the deal announcenemt. More specifically, as argued above, past performance is a potential determinant of both the probability of a pre-announcement and the market's reaction to information pertaining to asset sales. As such, we measure the buy-and-hold abnormal returns to the selling firm in the year preceding returns and up to two trading days prior to the pre-announcement[15], and winsorize this at 1st and 99th percentiles (*Ex-ante_BHAR*). Furthermore, the market may adjust its initial reaction to the pre-announcement during the between period as more information regarding the deal is disseminated (e.g., updates regarding the deal are often provided during conference calls), while this information could also affect the market's reaction to the deal-announcement. As such, we measure the buy-and-hold abnormal returns to the selling firm over the two days after the pre-announcement and two days prior the deal-announcement for pre-announced deals (*Runup).* To facilitate a comparison we use 178 days prior and up to two days before the deal-

[15] For non-pre-announced deals, we measure the one year buy-and-hold returns up to 178 days (sample average of the time between pre-announcement and deal-announcement of pre-announced deals) prior to the deal-announcement.

announcement for the non-pre-announced deals, winsorized at 1st and 99th percentiles.

2.3.3.Other Variables and Summary Statistics

We proxy the persistence of the productivity of sold assets (*Persistence*) following Yang (2008) and estimate the coefficient of a regression of firms' productivity on their one-year lagged productivity within each target's two-digit SIC industry over the 1998-2016 period, where productivity is measured as operating income after depreciation divided by total assets[16]. Note that this measure is constant at the target's two-digit SIC industry.

To capture changes in the prospects of the selling firm's remaining operations, we construct an indicator variable *IndShock,* which takes on the value of one in case the selling firm operates in multiple industries and any of its remaining operation's industry receives a positive demand shock, and zero otherwise. A positive demand shock is an indicator variable if the growth of the Tobin's q[17] of an industry's single industry firms is in the highest quintile over the 1980-2016 period, and zero otherwise.

We further control for other observable firm characteristics, which previous literature has shown to be associated with voluntary disclosures. In particular, given that larger firms are more likely to provide voluntary disclosures (Bamber and Cheon, 1998), we include a proxy for firm size, *(ln)MVE*, constructed as the firm market value of equity at the end of the

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[16] In line with Yang (2008), we delete industries with less than 50 observations, after requiring that each firm occurs at least 5 times in the sample period to obtain a stable time series.

[17] Tobin's q is measured as ((at - ceq)+(prcc_f*csho))/at

previous fiscal year end. Based on Welker (1995) and Frankel et al (1995) we include a proxy for industry reliance on external financing (*ExtFinDep*), which we measure as the ranking of the two-digit SIC industry median need for external financing ([capx – oancf]/capx) of all firms in the industry over the 1994-2004 period following Rajan and Zingales (1998) and Acharya and Xu (2017). We also include *Leverage*, measured as the ratio of total debt to book value of total asset ((dlc+dltt)/at), to capture any differential demand for voluntary disclosures by providers of capital to the firm (Vashishtha, 2014). In line with Johnson et al (2001), we proxy a firm's exposure to litigation risk (*Litigious*) as an indicator variable that takes on the value of one in case the selling firm belongs to industries prone to litigation risk, i.e. computer hardware (SIC codes 3570–3577), computer software (SIC codes 7371–7379), or pharmaceuticals (SIC codes 2833–2836) industries, and zero otherwise. We furthermore construct a variable that captures the profitability of the selling firm and sold asset (*Profit)*, as it has been shown to affect disclosure decisions (Dedman and Lennox, 2009).

Finally, we measure several characteristics of the deals in our sample. Our main variable of interest is *PreAnn* – an indicator variable that takes on the value of one in case the deal was pre-announced, and zero otherwise. We measure the consideration paid in millions of U.S. dollars (*Deal Value*), and calculate the ratio of the deal value to the seller's market value of equity at the end of the previous fiscal year end (*Relative Size*). Given that announcements may be bundled with other news, we create the indicator variables *ConcurrentInfo-Deal* and *ConcurrentInfo-PreAnn*, which take on the value of one in case the deal announcement or pre-announcement, respectively, was within one day of the reporting date of quarterly earnings announcements, and zero otherwise. *ConcurrentInfo* is an indicator variable that takes on the value of one in case either the deal- or the pre-announcement was within one day of the reporting date of quarterly earnings announcements, and zero otherwise. For pre-announced deals, we measure the time in days between the pre-announcement and the deal-announcement (*Time-to-Completion*). *Related Asset* is an indicator variable that takes on the value of one in case the two-digit SIC industry of the asset sold equals that of the seller, and zero otherwise. We create several variables aimed to capture the characteristics of the buyers. More specifically, we create an indicator variable that takes on the value of one in case the two-digit SIC industry of the asset sold equals that of the buyer, and zero otherwise (*Intra-industry*). Similarly, *Foreign Buyer* is an indicator variable that takes on the value of one in case the buyer is not a U.S. listed firm, and zero otherwise. We further refine the possible categories by partitioning *Intra-industry* and *Foreign Buyer* into six non-overlapping subsets of binary indicator variables. More specifically, we distinguish between intra-industry buyers which share the same two-digit SIC code as the asset sold (*Inside*), financial buyers with SIC codes 6000-6999 (*Financial*), and non-financial inter-industry buyers which do not share the two-digit SIC code of the asset sold (*Outside*), for both U.S. listed firms (*Domestic*) and non-U.S. listed firms (*Foreign*). Table 1 summarizes deal and firm characteristics of the full sample.

[Insert Table 1 here]

The mean of *PreAnn* indicates that 42% of asset sales are pre-announced, which shows the pervasiveness of prior information dissemination by firms in the market of corporate asset sales and the empirical importance of taking into account these pre-announcements. The sample average of *Deal*

Value equals \$745 million, which translates into an average *Relative Size* of 21%. In line with managers having more discretion regarding the timing of the pre-announcement, the results in Table 1 indicate that 36% of pre-announcements are bundled with earnings announcement, whereas only 16% of deal announcements coincide with earnings announcements. Furthermore, the average (median) pre-announcement precedes the deal-announcement by 178 (139) days. The results in Table 1 further show that 75% of the deals in our sample involve the sale of assets from the same industry as the seller. The buyers in our sample are from the same industry as the sold asset in 48% of the time, while firms not listed in the U.S. are the buyers in 27% of the sales. Turning to selling firm characteristics, the results indicate that the distribution of *MVE* is skewed, and that in line with expectations, asset sales are preceded by negative stock performance.

2.4 Determinants of pre-announcement

2.4.1 Bivariate Analysis

As a first step in our analysis of the determinants of pre-announcement, we compare deal and firm characteristics across the two deal types and report the results in Table 2. Importantly, we find that on average pre-announced deals are 2.8 times larger than non-pre-announced deals. This is in line with attempts to increase the pool of potential buyers: a key determinant of the number of potential buyers is the financial ability of potential buyers to acquire a selling firm's assets, which is negatively related to the size of the intended deal (Shleifer and Vishny, 1992). The difference in the relative size of the deals however is statistically indistinguishable from zero, entailing that pre-announcing firms are on average larger. Furthermore, the significant difference between the deal values implies that the value-weighted proportion of pre-announced asset sales equals 67%. The results in Table 2 further indicate that pre-announced deals more often involve assets from the same industry as the seller. This finding refutes the expectation that pre-announcements are instigated by improved prospects in the selling firm's remaining operations. Furthermore, while the proportion of assets acquired by foreign and within-industry buyers does not differ significantly, significant differences across buyers emerge when refining these categories. More specifically, the results show that pre-announced deals significantly more often involve a U.S. listed financial buyer, while they end up being acquired significantly less often by domestic buyers that do not operate in the same industry.

[Insert Table 2 here]

Table 2 provides an overview of the characteristics of the selling firms. As expected based on average deal value and the relative size of the transactions, pre-announcing firms are significantly larger than their non-pre-announcing counterparts. Importantly, we find that firms that pre-announce their asset sales have worse stock performance prior to the pre-announcement than their non-pre-announcing counterparts. More specifically, the results show that pre-announcing firms underperform non-pre-announcing firms by a statistically and economically significant 7.5%. Note that in the runup period, the returns no longer differ, which entails that relative to the non-pre-announcing sample the stock returns of the pre-announcing firms have improved. Also in line with our expectations, the remaining operations of selling firms more often receive a positive shock in pre-announced deals.

Contrary to our expectations however, the difference in the means of *Persistence* across the two sample is not statistically different from zero. Furthermore, pre-announcing firms more often operate in industries that rely on external capital, which is in line with Frankel et al (1995) who report a positive association between firm's tendency to access capital markets and disclosure of information. Note however that at the firm level no difference on the use of external (debt) can be discerned. Finally, the results presented in Table 2 show that the pre-announcing firms have similar exposures to litigation risk, and, contrary to the buy-and-hold abnormal returns prior to the pre-announcement, are average more profitable than non-pre-announcing firms. Overall, we find that pre-announced asset sales involve larger deals in absolute value, conducted by firms with poorer stock performance and more improved prospects in their remaining operations.

2.4.2 Probit Regressions

As the second step in our analysis of the determinants of pre-announcement, we estimate the following probit regression:

$$Pr(PreAnn = 1) = \beta_0 + \beta_1 Persistence + \beta_2 Ex\text{-}ante\_BHAR + \beta_3 IndShock + Z_i\gamma + \varepsilon_i \qquad (1)$$

where the coefficients on *Persistence*, *Ex-ante_BHAR* and *IndShock* are aimed at testing hypotheses H1, H2 and H3 respectively, Z_i denotes publicly known and exogenous control variables as discussed in section III, γ is a vector of probit coefficients, and ε_i is orthogonal to public variables Z_i.

Table 3 shows estimation results for the specification that models the decision of a manager to disclose an intended transaction. Given that *Persistence* could not be estimated for all observations, we run our probit specification in stages: the first three models in Table 3 report the results of the probit regressions with either *Persistence*, *Ex-ante_BHAR* or *IndShock*, where all models do include the full set of the control variables discussed previously. We then run an unrestricted model, which includes all the variables. As the estimated coefficients across the models, i.e. model (4) vs models (1), (2) or (3), are effectively identical, we will discuss the results of model (4).

The estimated coefficient on *Persistence*, aimed to capture manager's attempt to signal the quality of the asset in play, is not significantly different from zero and we thus find no evidence in line with H1. Our explanation for this is that managers do not use pre-announcements to signal the quality of the *average* deal, where the cost of disclosing could outweigh the benefits. More specifically, when assets' productivity levels are volatile, the interest in those assets may vary as well, increasing the likelihood of not finding a buyer. This increases the cost of a pre-announcement, and the average deal may not involve assets of sufficiently high quality to overcome this threshold.

While the estimated coefficient on *Persistence* is inconsistent with the predictions from our hypotheses, the estimated coefficients on *Ex-ante_BHAR* and *IndShock* are consitent. More specifically, the estimated coefficient on *Ex-ante_BHAR*, our variable on interest for testing H2, is negative and significant (-0.473, *t*-value: -2.04), in line with the argument that managers that sell a part of the firm that contributes to the poor past performance have an incentive to promptly inform markets of this. The estimated coefficient on *IndShock*, our variable on interest for testing H3, is positive and significant (0.564, *t*-value: 2.77), which is in line with managers precipitously informing markets or their plans to sell assets when the prospects of their remaining operations improve. The effects of these variables are also economically significant: a one standard deviation increase in *Ex-ante_BHAR* increases the probability of pre-announcing an asset sale by 5.5%, while selling firms which have received a positive demand shock to any of their remaining operation's industry are 19% more likely to pre-announce intended asset sales.

Of the additional variables, *Size*, *Relative Size* and *ExtFinDep* have significant effects on the decision to pre-announce the sale of assets. The positive coefficient of *Size* indicates that larger firms are more likely to pre-announce intended asset sales. We offer two explanations for this. First, it is relatively less costly for large firms to provide disclosures (Bamber and Cheon, 1998). Second, due to our sample selection criteria (i.e., deal value is required to be at least 5% of the seller's market value of equity) assets sold by larger firms in our sample are larger. As there are fewer potential buyers for large assets, the benefits of a pre-announcement may be larger for larger firms. We also provide two explanations for the positive coefficient of *Relative Size*. First, the importance to inform investors in a timely fashion is positively related to the materiality of the information which is in turn increasing in the relative size of firm's operations that are discontinued. Second, and more related to our hypotheses, the relative size of the asset sale is likely to be positively related to the expected improvements in the remaining firm's operations. That is, in case the sold asset is the culprit to the negative past performance of the selling firm, the improvement post-sale should be increasing in the size of the sold asset. In case the asset is sold due to improved prospects of the remaining operations of the firm, the willingness to sell a large portion of the firm is both a stronger signal, as well as a larger influx of capital which can be used to finance growth opportunities. Finally, the positive coefficient of *ExtFinDep* is in line with Frankel et al (1995) who report a positive association between firm's tendency to access capital markets and disclosure of information.

[Insert Table 3 here]

2.5 Stock Market Reaction to Pre-announcements and Deals

In this section, we report and compare the stock market's reaction to the pre- and deal-announcements. The first column of Table 4 (Panel A) reports the average cumulative abnormal return to the deal-announcement for the entire sample (i.e., both the pre-announced and non-pre-announced deals), similar to the previous literature. The magnitude of the market's reaction (1.54%) is similar to those reported in other studies (e.g., Borisova et al, 2013), and confirms that asset sales evoke a positive reaction by shareholders. However, when we distinguish between the returns to pre-announced and non-pre-announced deals, we find that the deal-announcement returns to pre-disclosing firms are less than half of those that accrue to the non-pre-disclosing firms (0.88% vs 2.02%). The returns on the pre-announcement, however, are significantly larger than those on the deal-announcement for the pre-announced deals (2.41%), which translates into an underestimation of market reaction to the full sample of asset sales of 40%. Also, note that the pre-announced deals are much larger than their non-pre-announced counterparts, rendering the omission of this part of the market's reaction to asset sales even more economically significant. The difference between the market's reaction to the pre-announcement and deal-announcement entails that markets not only consider the disclosure of the intention to sell to be value-relevant news, but also deem the completion of the deal as very likely as they incorporate over 70% of the total effect on the pre-announcement date. Nevertheless, despite the market's positive reaction to the pre-announcement, the sum of the announcement period returns to the pre-announcement and the deal-announcement (3.29%) is not significantly larger than the market's reaction to non-pre-announced deals. The results of the statistical tests of these comparisons are reported in Panel B. The results in Panels C and D of Table 4 further show that the buy-and-hold abnormal returns in the runup period for the two type of deals do not differ significantly from zero or each other.

[Insert Table 4 here]

We next investigate whether the stock market reactions to pre-announced asset sales are more positive than non-pre-announced asset sales after controlling for other determinants of market reactions to asset sales (H4). We estimate the following OLS regression model on the full sample of asset sales, where the coefficient on *PreAnn* captures the difference in the market reaction between the two type of deals:

$$\begin{aligned} Total\_CAR = {} & \beta_0 + \beta_1 PreAnn + \beta_2 Persistence + \beta_3 Ex\text{-}ante\_BHAR + \\ & \beta_4 IndShock + \beta_5 (ln)MVE + \beta_6 Relative\ Size + \beta_7 Litigious + \\ & \beta_8 ExtFinDep + \beta_9 Leverage + \beta_{10} Profit + \beta_{11} Related\ Asset + \\ & \beta_{12} Intra\text{-}industry + \beta_{13} Foreign\ Buyer + \varepsilon \qquad (2) \end{aligned}$$

We add *Persistence*, *Ex-ante_BHAR* and *IndShock*, as these are the variables that capture managerial signaling incentives. We further control for the relative size of the sale (*Relative Size*), to capture the change in the scope of the selling firms activities, as well as size (*(ln)MVE*) and exposure to litigation risk (*Litigious*). As market reactions to asset sales have been shown to be associated with selling firms' demand for external capital (e.g., Asquith et al, 1994; Lang et al, 1995; Bates, 2005), we further add *ExtFinDep, Leverage,* and *Profit* to our specification. Given that announcements may be bundled with other news, we also include *ConcurrentInfo* to the regression model. Finally, we control for the type of asset that has been sold (*Related Asset*) and type of buyer (*Intra-industry* and *Foreign Buyer*). We further include year fixed effects, as the market reaction to asset sales may vary systematically with macroeconomic conditions[18]. The construction of variables is discussed in section III.

The results of our main specification are reported in Model 3 of Table 5. The positive and significant coefficient of *PreAnn* indicates that, in line with our expectations, the market reacts more positively to pre-announced deals. In particular, keeping other determinants of market reaction to asset sales constant, pre-announced asset sales elicit an economically significant 2.1% larger CAR relative to non-pre-announced asset sales. Thus, the positive association between the CAR and pre-announcing asset sales supports the idea that managers act on their incentives to expedite the disclosure of positive news. The results reported in Table 5 further indicate that markets react significantly more positively when the selling firm's stock performance in the year preceding the runup was negative and a larger proportion of the firm is divested. Contrary to our expectations, the coefficient of *IndShock* is statistically and economically significantly negative, entailing that asset sales that coincide with a positive shock to the firm's remaining operations elicit a

[18] Our results are robust to the inclusion of industry fixed effects.

2.8% lower CAR. Our results also do not indicate that the persistence of the productivity of the sold assets affect investors' reactions. Furthermore, the negative coefficient of *ConcurrentInfo* implies that the markets reacts less positively when announcements regarding asset sales are bundled with earnings announcements. Finally, our results do not show that market reactions differ with either the type of asset which is sold (*Related Asset*), or type of buyer, regarding industry classification or nationality (*Intra-industry* and *Foreign Buyer*, respectively). Given evidence to the contrary (e.g., Borisova et al, 2013), we further refine the type of buyers (see Model 4), but our inferences remain the same.

[Insert Table 5 here]

We next proceed by investigating the determinants of the market reactions to all announcements by estimating a modified version of our main regression model for each event separately. Model 1 of Table 6 (Panel A) shows the results of the estimated coefficients for the market reactions to the pre-announcements. Consistent with earlier findings, the persistence of the productivity of the sold assets affect investors' reactions is not related to the market reactions at the time of the pre-announcement. Inconsistent with the predictions from our hypotheses, a concurrent positive shock to the remaining operation of the selling firm has a significantly negative effect on the market's reaction to the news of intended asset sales. Nevertheless, taken together the signs of the other significant coefficients in Model 1 strongly supports the idea that managers pre-announce asset sales that are aimed to turn around poor prior performance and financial distress. More specifically, the estimated coefficient of *Ex-ante_BHAR* is statistically significantly negative, implying that market reactions to pre-announcement are higher the more poorly the selling firm's stock performs in the year preceding the pre-announcement. This coefficient (-0.04) is also economically significant: a one standard deviation increase in *Ex-ante_BHAR* translates into a 1.3% higher CAR. This result is especially noteworthy given that the pre-announcement only reveals intended asset sales. This interpretation is further supported by the negative coefficient of *Profit*. In the same vein, when the pre-announcement is bundled with earnings announcement (*ConcurrentInfo_PreAnn*), the CAR are 3.3% lower – which equals the mean value of *Total_CAR* for the entire sample of pre-announced asset sales. In line with these results that suggest poor performance, we interpret the significantly positive coefficient of *Leverage* as shareholders of highly levered firms reacting positively to the possibility of an asset sale to avert bankruptcy. In particular, as Asquith et al (1994) show, asset sales are a way financially distressed firms can avoid bankruptcy, but firms in highly leveraged industries are limited in doing so. As argued in section II, the costs accompanying a pre-announcement render it credible, thus alleviating concerns shareholders may have that asset sales are not a viable option. Finally, although the coefficient of *Relative Size* is insignificant, its deviation from its coefficient in the other models and our prior that selling a larger proportion of the firm would elicit a more positive reaction, merits further discussion. A plausible reason for the near negative coefficient is that the size of asset sale offers new information to the market as to the gravity of the selling firm's underlying issues.

When turning to the results of Model 2, where the dependent variable is *Deal_CAR* and the sample consists of the pre-announced asset sales, the low explanatory power of the model is of note. More specifically, whereas the adjusted R^2 of Model 1 – a regression estimated on the market reactions to *planned* asset sales – equals 18,7%, the adjusted R^2 of Model 2 – a regression run on market reactions to *actual* asset sales that were pre-announced – equals -1.9%. In fact, in Model 2 no variable other than *Ex-ante_BHAR* is significantly related to the market reaction. The stark difference between the explanatory power of the two models, coupled with the results presented in Table 4 suggests that the market deems pre-announcements credible and that most of the implications of the asset sale are priced in at the time of the pre-announcement. Another potential reason for the low explanatory power of the Model 2 could be that in addition to pre-announcement, the transaction is highly anticipated. We therefore augment our main specification with *PreAnn_CAR*, *Runup* and *(ln)Time-to-Completion* (Model 3), and find that this substantially increases the model's explanatory power to 5.9%, although it still remains well below that of Model 1. The significant positive coefficient of *PreAnn_CAR* is in line with diminished uncertainty when the deal is announced, while the significantly negative coefficient of *Runup* indicates that indeed the deal has been anticipated during the runup period. Note that an alternative interpretation of the negative coefficient of *Runup* would be that the poor stock performance continues in the runup period. In Model 4 we further refine our proxies of the type of buyer, but the inferences remain unchanged.

[Insert Table 6 here]

The results of Model 5, where the sample consists of the asset sales that were not pre-announced, further underline the difference between the two type of transactions. In particular, the insignificant coefficients of *Persistence*, *Ex-ante_BHAR* and *IndShock* are in line with the absence of any signaling incentives. Furthermore, it is striking that none of the factors which determine the market reaction to either type of deal affect the market reaction to the other type, or in case of *Leverage* even has a significant effect in the opposite direction. Taken together, the results are in line with the idea that non-pre-announced asset sales are not aimed at turning around the performance of the firm or pursue more profitable growth opportunities available to the firm, and are potentially initiated by foreign bidders who diversify their operations in the U.S. and are willing to pay a high premium for those assets[19]. More specifically, the insignificant coefficients of *Ex-ante_BHAR*, *Profit*, and *ConcurrentInfo_Deal* run counter to the notion that the sale was prompted to turnaround prior poor performance, while the negatively significant coefficient of *Leverage* is in line with the idea that deals by selling firms not in financial distress, thus enjoying a strong bargaining position, generate higher stock market reactions. The results also do not support the notion that the sale is instigated by the intention to reallocate focus and resources towards projects in other areas that hold more promise. As such, these findings cast doubt on the argument that these assets sales are motivated by the seller. Together with the significantly positive coefficients on *Foreign Buyer* and *Relative Size*[20], the results suggest that these asset sales involve assets that were sought after and were initiated by foreign bidders. When we further refine the type of buyers in Model 6, the results indicate the foreign buyers from industries other that the target drive the positive market reactions.

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[19] Alternatively, the results in no way indicate that the positive market reactions are due to the information the deal reveals about the prospects of the remaining assets of the firm, while they are in line with the seller being able to receive a high value for the assets they sell.

[20] An unanticipated asset sale could elicit a negative market reaction, as it could be construed as negative information regarding the quality of the asset the market was not privy to. However, when the consideration is higher than the perceived market value of those assets, a positive relationship between *Relative Size* and *Deal_CAR* is likely to ensue.

We repeat our analyses of the determinants of the market reaction to the two deal types where we replace the dependent variable from the previous section (*PreAnn_CAR* in Model 1 and *Deal_CAR* in Models 2 - 4) with *Total_CAR*. The results of these tests are reported in Panel B of Table 6. Note that Model 4 and 5 are the same as Model 5 and 6 of Panel A, and are included to facilitate comparisons. The results of Model 1, 2, and 3 do not change our previously stated interpretations. The coefficient of *Foreign_Financial* in Model 3 is now significantly negative, which suggests that these are the buyers of last resort and that having pre-announced intention to sell assets can in fact hurt the bargaining position of the selling firms.

Taken together, the results reported in Table 6 provide mixed support for our stated hypotheses. Specifically, because we find a negative coefficient of *IndShok* we reject the hypothesis that pre-announcements serve to signal improvements of the remaining operations of the firm. In contrast, the negative coefficients of *Ex-ante_BHAR, Profit* and *ConcurrentInfo_PreAnn* and the positive coefficient of *Leverage* support the hypothesis that pre-announcements are used to signal the turnaround of poor prior performance and financial distress. Finally, our results provide indications that non-pre-announced asset sales involve the sale of assets in high demand, potentially initiated by foreign bidders

2.6 Conclusion

Corporate asset sales are one of the most common ways productive assets are reallocated, making up approximately half of all M&A transactions (Maksimovic and Phillips, 2001). Despite the operational and financial effects of these transactions on the selling firms, little is known about the opaque selling process (Borisova et al, 2013). This study documents and investigates the effect of the pervasive use of voluntary disclosures through which selling firms inform potential buyers and capital markets that certain assets are available for sale. More specifically, we show that in 42% of completed asset sales the selling firm pre-announces its intention to divest, and find that these announcements elicit economically and statistically significant positive market reactions. Our analyses further indicate that pre-announcements are used to signal the turnaround of poor prior performance and financial distress. Furthermore, our results provide some indications that non-pre-announced asset sales involve the sale of assets that are strongly sought after, potentially initiated by foreign bidders. These seemingly different incentives for the two type of asset sales are also in line with our main finding that markets react more positively to pre-announced deals than to non-pre-announced deals. In particular, pre-announced deals imply that the future of the remaining operations will improve, which constitute the majority of operations of the selling firm. The returns to the non-pre-announced deals however seem to reflect a premium paid for the sold asset, which generally constitutes a minority part of the selling firm. Importantly, our results indicate that markets deem pre-announcements credible and that most of the valuation effects of the pre-announced asset sales are incorporated into the stock price of the selling firm prior to the deal-announcement. This, coupled with our results which suggest that the decision to pre-announce is related to the motivation for the asset sale, has implications for empirical tests of asset sales.

Figure 1

This figure shows the distribution of asset sales delineated by deal type over years. The sample consists of asset sales by U.S. listed firms from 2005-2015, as further described in the sample selection section.

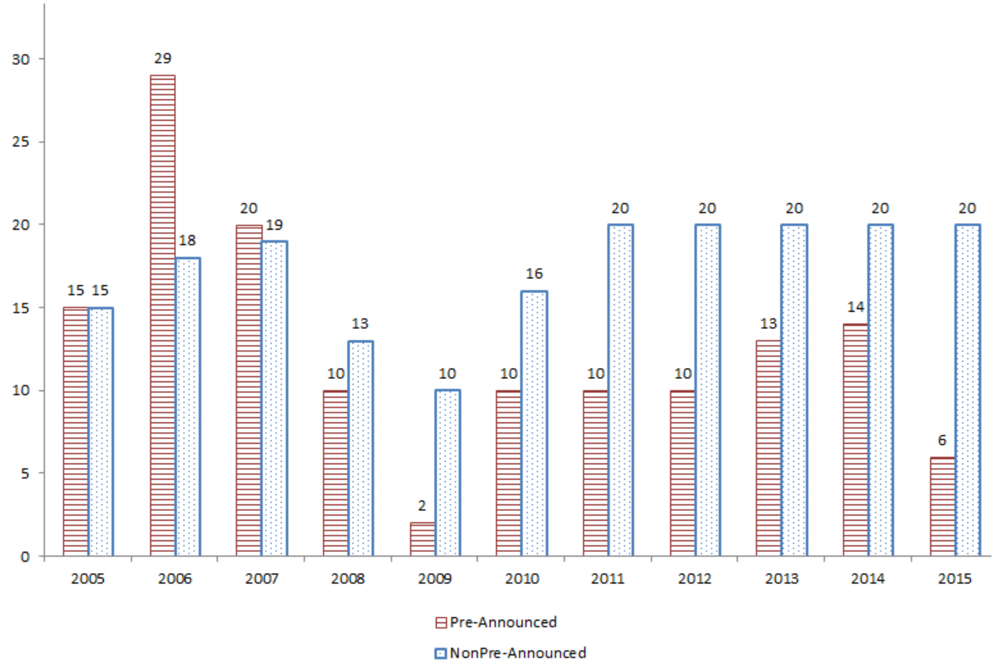

Table 1: Summary statistics

This table summarizes deal (Panel A) and firm characteristics (Panel B) of the sample, which consists of asset sales by U.S. listed firms from 2005-2015, as further described in the sample selection section. *PreAnn* is an indicator variable that takes on the value of one in case the deal was pre-announced, and zero otherwise. *Deal Value* denotes the consideration paid (mil US dollars). *Relative Size* is the ratio of the deal value to the seller's market value of equity at the end of the previous fiscal year end. *ConcurrentInfo-Deal (ConcurrentInfo-PreAnn)* is an indicator variable that takes on the value of one in case the deal- (pre-) announcement was within one day of the reporting date of quarterly earnings announcements, and zero otherwise. *ConcurrentInfo* is an indicator variable that takes on the value of one in case either the deal- or the pre-announcement was within one day of the reporting date of quarterly earnings announcements, and zero otherwise. *Time-to-Completion* denotes the time in day between the pre-announcement and the deal-announcement. *Related Asset* is an indicator variable that takes on the value of one in case the two-digit SIC industry of the asset sold equals that of the seller, and zero otherwise. *Intra-industry* is an indicator variable that takes on the value of one in case the two-digit SIC industry of the asset sold equals that of the buyer, and zero otherwise. *Foreign Buyer* is an indicator variable that takes on the value of one in case the buyer is not a U.S. listed firm, and zero otherwise. *Domestic-Inside* is an indicator variable that takes on the value of one if the buyer is a U.S. listed firm and its 2-digit SIC industry of the asset sold equals, and zero otherwise. *Domestic-Financial* is an indicator variable that takes on the value of one if the buyer is a U.S. listed financial (i.e., SIC code 6000-6999) firm, and zero otherwise. *Domestic-Outside* is an indicator variable that takes on the value of one if the buyer is a U.S. listed non-financial firm with a different two-digit SIC industry than the asset sold, and zero otherwise. *Foreign-Inside* is an indicator variable that takes on the value of one if the buyer is a non-U.S. listed firm and its two-digit SIC industry of the asset sold equals, and zero otherwise. *Foreign-Financial* is an indicator variable that takes on the value of one if the buyer is a non-U.S. listed financial firm, and zero otherwise. *Foreign -Outside* is an indicator variable that takes on the value of one if the buyer is a U.S. listed non-financial firm with a different two-digit SIC industry than the asset sold, and zero otherwise. *MVE* denotes the seller's market value of equity at the end of the previous fiscal year end. *(ln)MVE* indicates the natural logarithm of market value of equity at the end of the previous fiscal year end. *Ex-ante_BHAR* is the buy-and-hold abnormal return to the selling firm in the year preceding either the pre-announcement for pre-announced deals, or the 6 months lag of the deal-announcement for the non-announced deals, winsorized at 1st and 99th percentiles. *Runup* is the buy-and-hold abnormal return to the selling firm over the 2 days after the pre-announcement and 2 days prior the deals-announcement for pre-announced deals, or the 6 months prior and up to 2 days before the deal-announcement for the non-announced deals, winsorized at 1st and 99th percentiles. *IndShock* is an indicator variables that takes on the value of one in case any of the selling firm's remaining operation's industry receives a positive demand shock, and zero otherwise. A positive demand shock is an indicator variable that takes on the value of one if the growth of the Tobin's q (((at - ceq)+(prcc_f*csho))/at) of an industry is in the highest quintile over the 1980-2016 period, and zero otherwise. *Persistence* is the estimated coefficient of a regression of firms' productivity on their one-year lagged productivity within each target's two-digit SIC industry over the 1998-2016 period. *ExtFinDep* is a measure of firm's need for external financing, measured as the ranking of the two-digit SIC industry median need for external financing (i.e., [capx – oancf]/capx) of all firms in that industry over the 1994-2004 period. *Leverage* is the ratio of total debt to book value of total asset ((dlc+dltt)/at). *Litigious* is an indicator variable that takes on the value of one in case the selling firm belongs to industries prone to litigation risk, i.e. the computer hardware (SIC codes 3570–3577), computer software (SIC codes 7371–7379), or pharmaceuticals (SIC codes 2833–2836) industries, and zero otherwise (Johnson et al, 2001). *Profit* indicates the selling firm's profitability (oibdp/at).

	N	*Mean*	*StDev*	*p1*	*p25*	*p50*	*p75*	*p99*
Panel A: Deal Characteristics								
PreAnn	330	0.421	0.495	0	0	0	1	1
Deal Value	330	745	1555	53	110	238	756	7000
Relative Size	330	0.218	0.187	0.014	0.084	0.152	0.294	0.773
ConcurrentInfo-Deal	330	0.158	0.365	0	0	0	0	1
ConcurrentInfo-PreAnn	139	0.36	0.482	0	0	0	1	1
ConcurrentInfo	330	0.288	0.453	0	0	0	1	1
Time-to-Completion	139	178	138	9	85	139	233	634
(ln)Time-to-Completion	139	4.88	0.839	2.197	4.443	4.934	5.451	6.452
Related Asset	330	0.745	0.436	0	0	1	1	1
Intra-industry	330	0.476	0.5	0	0	0	1	1
Foreign Buyer	330	0.273	0.446	0	0	0	1	1
Domestic-Inside	330	0.336	0.473	0	0	0	1	1
Domestic-Financial	330	0.224	0.418	0	0	0	0	1
Domestic-Outside	330	0.167	0.373	0	0	0	0	1
Foreign-Inside	330	0.139	0.347	0	0	0	0	1
Foreign-Financial	330	0.045	0.209	0	0	0	0	1
Foreign-Outside	330	0.088	0.284	0	0	0	0	1
Panel B: Selling Firm Characteristics								
MVE	330	10969	37043	98	620	1727	4861	211132
(ln)MVE	330	7.595	1.651	4.588	6.429	7.454	8.489	12.26
Ex-ante_BHAR	330	-0.045	0.334	-0.82	-0.243	-0.081	0.107	1.09
Runup	330	0.018	0.231	-0.692	-0.109	0.016	0.128	0.667
IndShock	330	0.17	0.376	0	0	0	0	1
Persistence	319	0.51	0.266	0.149	0.305	0.522	0.62	1.724
ExtFinDep	330	29.782	15.706	3	16	28	47	49
Leverage	330	0.313	0.205	0	0.183	0.287	0.424	0.941
Litigious	330	0.203	0.403	0	0	0	0	1
Profit	330	0.122	0.116	-0.261	0.081	0.125	0.169	0.37

Table 2: Deal and firm characteristics by deal type

This table reports the (difference in) means of deal and firm characteristics of the sample delineated by deal type. The sample consists of asset sales by U.S. listed firms from 2005-2015, as further described in the sample selection section. The reader is referred to the caption of Table 1 for variable construction. Significance levels of the two sample mean comparison tests are denoted by ***, **, and * indicating p<0.01, p<0.05, and p<0.010 levels respectively.

	Pre-Announced		*Non-Pre-Announced*		*Difference*	
Panel A: Deal Characteristics	*N*	*Mean*	*N*	*Mean*	*Mean*	*p-*
Deal Value	139	1186	191	424	761***	0.00
Relative Size	139	0.233	191	0.206	0.027	0.19
ConcurrentInfo-Deal	139	0.108	191	0.194	-	0.03
ConcurrentInfo	139	0.417	191	0.194	0.224**	0.00
Related Asset	139	0.799	191	0.707	0.091*	0.05
Intra-industry	139	0.432	191	0.508	-0.076	0.17
Foreign Buyer	139	0.295	191	0.257	0.038	0.44
Domestic-Inside	139	0.302	191	0.361	-0.059	0.26
Domestic-Financial	139	0.317	191	0.157	0.159**	0.00
Domestic-Outside	139	0.086	191	0.225	-	0.00
Foreign-Inside	139	0.129	191	0.147	-0.017	0.65
Foreign-Financial	139	0.065	191	0.031	0.033	0.15
Foreign-Outside	139	0.101	191	0.079	0.022	0.48
Panel B: Selling Firm						
MVE	139	13037	191	9465	3572	0.38
(ln)MVE	139	8.014	191	7.290	0.724**	0.00
Ex-ante_BHAR	139	-0.089	191	-0.014	-	0.04
Runup	139	0.017	191	0.018	-0.001	0.97
IndShock	139	0.237	191	0.120	0.117**	0.00
Persistence	132	0.486	187	0.527	-0.041	0.17
ExtFinDep	139	33.849	191	26.822	7.027**	0.00
Leverage	139	0.323	191	0.306	0.017	0.46
Litigious	139	0.173	191	0.225	-0.052	0.24
Profit	139	0.134	191	0.113	0.021*	0.09

Table 3: Predicting pre-announcements

This table presents results of probit regressions for the probability of pre-announcing an asset sales. The dependent variable in all models is *PreAnn*, an indicator variable that takes on the value of one in case the deal was pre-announced, and zero otherwise. The sample consists of asset sales by U.S. listed firms from 2005-2015, as further described in the sample selection section. All models are estimated using the full. The reader is referred to the caption of Table 1 for variable construction. The symbols ***, **, and * denote statistical significance at p<0.01, p<0.05, and p<0.010 levels respectively.

	(1)	(2)	(3)	(4)
Persistence	-0.004			0.066
	(-0.01)			(0.20)
Ex-ante_BHAR		-0.473		-0.493
		(-2.04)**		(-2.03)**
IndShock			0.564	0.547
			(2.77)***	(2.54)**
(ln)MVE	0.278	0.278	0.243	0.249
	(4.76)***	(4.84)***	(4.18)***	(4.13)***
Relative Size	2.021	1.944	1.933	2.011
	(4.10)***	(4.01)***	(3.98)***	(4.08)***
ExtFinDep	0.016	0.014	0.017	0.019
	(2.96)***	(2.85)***	(3.44)***	(3.33)***
Leverage	0.421	0.249	0.311	0.400
	(1.07)	(0.65)	(0.80)	(1.00)
Litigious	-0.114	-0.141	-0.081	-0.023
	(-0.57)	(-0.72)	(-0.41)	(-0.11)
Profit	0.611	0.733	0.813	1.048
	(0.83)	(0.98)	(1.08)	(1.33)
Constant	-3.456	-2.853	-3.274	-3.045
	(-6.06)***	(-5.01)***	(-6.12)***	(-5.03)***
Pseudo_R^2	0.115	0.114	0.121	0.137
N	319	330	330	319

Table 4: Wealth effects of asset sales

This table reports and compares the cumulative abnormal returns to the events related to asset sales (i.e., pre- and deal-announcement). Additionally, this table reports and compares the buy-and-hold abnormal returns to the selling firm over the 2 days after the pre-announcement and 2 days prior the deals-announcement for pre-announced deals, or the 6 months prior and up to 2 days before the deal-announcement for the non-announced deals (*Runup*). The sample consists of asset sales by U.S. listed firms from 2005-2015, as further described in the sample selection section. The reader is referred to section 3 for variable construction. Significance levels of the two sample mean comparison tests are denoted by ***, **, and * indicating p<0.01, p<0.05, and p<0.010 levels respectively.

Panel A: Cumulative Abnormal Returns

	(1)	(2)	(3)	(4)	(5)
Sample	Full	Non-Pre-Announced	Pre-Announced	Pre-Announced	Pre-Announced
N	(n=330)	(n=191)	(n=139)	(n=139)	(n=139)
Variable	*Deal_CAR*	*Deal_CAR*	*Deal_CAR*	*PreAnn_CAR*	*Total_CAR*
Mean	1.54%	2.02%	0.88%	2.41%	3.29%
(*p*-value)	0.000	0.000	0.049	0.000	0.000

Panel B: Differences CAR

	(2) - (3)	(2) - (5)
Difference in Means	1.13%	-1.27%
(*p*-value)	0.083	0.112

Panel C: BAHR

	(1)	(2)
Sample	Non-Pre-Announced	Pre-Announced
N	(n=191)	(n=139)
Variable	*Runup*	*Runup*
Mean	1.83%	1.74%
(*p*-value)	0.148	0.172

Panel D: Differences BHAR

	(2) - (1)
Difference in Means	-0.10%
(*p*-value)	0.485

Table 5: Multivariate analysis of stock market reactions to asset sales

This table reports the OLS regressions of cumulative abnormal returns to asset sales. The sample consists of asset sales by U.S. listed firms from 2005-2015, as further described in the sample selection section. The dependent variable in models 1 and 2 is *Deal_CAR* and *Total_CAR* in models 3 and 4. The reader is referred to section 3 for variable construction. All regressions include year fixed effects. The symbols ***, **, and * denote statistical significance at $p<0.01$, $p<0.05$, and $p<0.010$ levels respectively.

	(1)	(2)
	Total_CAR	*Total_CAR*
PreAnn	0.021	0.021
	(1.69)*	(1.67)*
Persistence	0.002	0.001
	(0.07)	(0.03)
Ex-ante_BHAR	-0.064	-0.065
	(-3.86)***	(-3.88)***
IndShock	-0.028	-0.025
	(-1.73)*	(-1.55)
(ln)MVE	-0.005	-0.005
	(-1.15)	(-1.14)
Relative Size	0.063	0.067
	(1.78)*	(1.88)*
Litigious	-0.004	-0.003
	(-0.32)	(-0.21)
ExtFinDep	0.000	0.000
	(0.18)	(0.30)
Leverage	-0.033	-0.032
	(-1.18)	(-1.13)
Profit	-0.068	-0.065
	(-1.36)	(-1.30)
ConcurrentInfo	-0.026	-0.026
	(-2.14)**	(-2.09)**
Related Asset	-0.009	-0.010
	(-0.70)	(-0.75)
Intra-industry	-0.006	
	(-0.53)	
Foreign Buyer	0.007	
	(0.58)	
Domestic_Financial		0.008
		(0.58)
Domestic_Outside		0.009
		(0.55)
Foreign_Inside		0.013
		(0.78)
Foreign_Financial		-0.029
		(-1.12)
Foreign_Outside		0.032
		(1.55)
Constant	0.138	0.128
	(3.26)***	(2.99)***
*Adj_* R^2	0.083	0.087
N	319	319

Table 6: Multivariate analysis of stock market reactions to asset sales by deal type

This table reports the OLS regressions of cumulative abnormal returns to the events related to asset sales. Regressions are run on sub-samples of the sample of asset sales by U.S. listed firms from 2005-2015, as further described in the sample selection section. The sample in Models 1, 2, 3 and 4 of Panel A, and Models 1, 2 and 3 in Panel B consists of pre-announced deals, and the remaining models are run on the sample of non-announced deals. The dependent variable in each model is indicated below the model number. The reader is referred to section 3 for variable construction. All regressions include year fixed effects. The symbols ***, **, and * denote statistical significance at $p<0.01$, $p<0.05$, and $p<0.010$ levels respectively.

Panel A	(1)	(2)	(3)	(4)	(5)	(6)
Sample	*Pre-Announced*				*NonPre-Announced*	
Variable	*PreAnn_CAR*	*Deal_CAR*	*Deal_CAR*	*Deal_CAR*	*Deal_CAR*	*Deal_CAR*
Persistence	0.006	-0.026	-0.022	-0.027	0.013	0.010
	(0.18)	(-0.80)	(-0.70)	(-0.83)	(0.57)	(0.43)
Ex-ante_BHAR	-0.040	-0.063	-0.046	-0.046	-0.028	-0.024
	(-1.78)*	(-3.07)***	(-2.25)**	(-2.23)**	(-1.58)	(-1.35)
IndShock	-0.047	-0.007	0.009	0.009	-0.015	-0.017
	(-2.62)***	(-0.39)	(0.54)	(0.50)	(-0.74)	(-0.82)
(ln)MVE	-0.007	-0.000	-0.000	0.000	-0.000	-0.000
	(-1.21)	(-0.01)	(-0.03)	(0.09)	(-0.06)	(-0.08)
Relative Size	-0.071	0.043	0.061	0.064	0.140	0.144
	(-1.64)	(1.06)	(1.53)	(1.58)	(3.40)***	(3.49)***
Litigious	-0.016	0.011	0.008	0.010	0.002	0.003
	(-0.89)	(0.64)	(0.50)	(0.58)	(0.14)	(0.23)
ExtFinDep	0.001	-0.000	-0.000	-0.000	0.000	0.000
	(1.43)	(-0.96)	(-0.94)	(-0.96)	(0.14)	(0.19)
Leverage	0.080	-0.020	-0.036	-0.030	-0.051	-0.049
	(1.88)*	(-0.49)	(-0.92)	(-0.74)	(-1.82)*	(-1.75)*
Profit	-0.170	0.044	0.121	0.126	-0.051	-0.054
	(-2.18)**	(0.60)	(1.61)	(1.64)	(-1.03)	(-1.08)

(continued on next page)

Panel A	(1)	(2)	(3)	(4)	(5)	(6)
Sample		*Pre-Announced*			*NonPre-Announced*	
Variable	*PreAnn_CAR*	*Deal_CAR*	*Deal_CAR*	*Deal_CAR*	*Deal_CAR*	*Deal_CAR*
ConcurrentInfo_PreAnn	-0.033		-0.018	-0.018		
	(-2.14)**		(-1.23)	(-1.24)		
ConcurrentInfo-Deal		-0.015	-0.004	-0.006	-0.016	-0.015
		(-0.71)	(-0.20)	(-0.31)	(-1.12)	(-1.06)
PreAnn_CAR			0.169	0.156		
			(1.95)*	(1.76)*		
Runup			-0.065	-0.065		
			(-2.24)**	(-2.20)**		
(ln)Time-to-Completion			0.004	0.005		
			(0.53)	(0.62)		
Related Asset	-0.010	-0.017	-0.011	-0.012	-0.004	-0.003
	(-0.61)	(-1.07)	(-0.68)	(-0.74)	(-0.28)	(-0.23)
Intra-industry		0.005	0.007		-0.011	
		(0.36)	(0.57)		(-0.96)	
Foreign Buyer		-0.007	-0.009		0.029	
		(-0.51)	(-0.63)		(2.18)**	
Domestic_Financial				-0.010		-0.004
				(-0.61)		(-0.26)
Domestic_Outside				-0.022		0.016
				(-0.94)		(1.01)
Foreign_Inside				-0.020		0.024
				(-1.04)		(1.37)
Foreign_Financial				-0.030		-0.001
				(-1.20)		(-0.04)
Foreign_Outside				0.002		0.063
				(0.07)		(2.76)***
Constant	0.131	0.100	0.047	0.050	0.039	0.026
	(2.29)**	(1.90)*	(0.71)	(0.73)	(0.79)	(0.53)
Adj R^2	0.187	-0.019	0.059	0.050	0.139	0.146
N	132	132	132	132	187	187

Table 6: Multivariate analysis of stock market reactions to asset sales by deal type (*continued*)

Panel B	(1)	(2)	(3)	(4)	(5)
Sample		*Pre-Announced*		*NonPre-Announced*	
Variable	*Total_CAR*	*Total_CAR*	*Total_CAR*	*Deal_CAR*	*Deal_CAR*
Persistence	-0.026	-0.026	-0.049	0.013	0.010
	(-0.49)	(-0.51)	(-0.94)	(0.57)	(0.43)
Ex-ante_BHAR	-0.091	-0.084	-0.082	-0.028	-0.024
	(-2.66)***	(-2.50)**	(-2.44)**	(-1.58)	(-1.35)
IndShock	-0.053	-0.045	-0.046	-0.015	-0.017
	(-1.97)*	(-1.66)*	(-1.67)*	(-0.74)	(-0.82)
(ln)MVE	-0.008	-0.007	-0.005	-0.000	-0.000
	(-0.96)	(-0.83)	(-0.61)	(-0.06)	(-0.08)
Relative Size	-0.043	-0.025	-0.016	0.140	0.144
	(-0.66)	(-0.38)	(-0.25)	(3.40)***	(3.49)***
Litigious	-0.008	-0.011	-0.006	0.002	0.003
	(-0.28)	(-0.41)	(-0.21)	(0.14)	(0.23)
ExtFinDep	0.000	0.000	0.000	0.000	0.000
	(0.52)	(0.58)	(0.45)	(0.14)	(0.19)
Leverage	0.049	0.047	0.068	-0.051	-0.049
	(0.76)	(0.75)	(1.05)	(-1.82)*	(-1.75)*
Profit	-0.152	-0.053	-0.051	-0.051	-0.054
	(-1.28)	(-0.44)	(-0.42)	(-1.03)	(-1.08)
ConcurrentInfo_PreAnn	-0.058	-0.061	-0.062		
	(-2.45)**	(-2.59)**	(-2.65)***		
ConcurrentInfo-Deal	-0.038	-0.029	-0.038	-0.016	-0.015
	(-1.13)	(-0.88)	(-1.15)	(-1.12)	(-1.06)
Runup		-0.111	-0.106		
		(-2.32)**	(-2.23)**		

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Panel B	(1)	(2)	(3)	(4)	(5)
Sample		*Pre-Announced*		*NonPre-Announced*	
Variable	*Total_CAR*	*Total_CAR*	*Total_CAR*	*Deal_CAR*	*Deal_CAR*
(ln)Time-to-Completion		0.015	0.013		
		(1.15)	(0.99)		
Related Asset	-0.030	-0.022	-0.030	-0.004	-0.003
	(-1.17)	(-0.87)	(-1.16)	(-0.28)	(-0.23)
Intra-industry	0.003	0.009		-0.011	
	(0.14)	(0.44)		(-0.96)	
Foreign Buyer	-0.033	-0.036		0.029	
	(-1.46)	(-1.61)		(2.18)**	
Domestic_Financial			0.005		-0.004
			(0.18)		(-0.26)
Domestic_Outside			-0.061		0.016
			(-1.61)		(1.01)
Foreign_Inside			-0.042		0.024
			(-1.34)		(1.37)
Foreign_Financial			-0.076		-0.001
			(-1.89)*		(-0.04)
Foreign_Outside			-0.014		0.063
			(-0.37)		(2.76)***
Constant	0.261	0.148	0.162	0.039	0.026
	(3.01)***	(1.35)	(1.48)	(0.79)	(0.53)
Adj_R^2	0.138	0.175	0.187	0.139	0.146
N	132	132	132	187	187

Chapter 3

The effect of IFRS on ADR liquidity[21]

3.1 Introduction

This chapter examines the effect of the mandatory adoption of International Financial Reporting Standards (IFRS) in a firm's domestic market on the liquidity of its American Depository Receipts (ADRs) traded on a U.S. exchange. Whereas prior research has examined the immediate effects of IFRS adoption on investor informedness, liquidity, and trading in domestic markets (Aharony et al. 2010; Jao et al. 2012; Byard et al. 2011; DeFond et al. 2011; Hong 2013; Hong et al. 2014; Landsman et al. 2012; Tan et al. 2011; Yu and Wahid 2014), little is known about whether IFRS adoption has similar or distinctively different effects on cross-listing markets outside the IFRS environment. In this study, we explicitly compare the effect of IFRS on the liquidity of ADR securities with that on the liquidity of domestic securities. Examining whether and how these effects differ is important for the following reasons. First, IFRS adoption and, more generally, moves towards convergence of international accounting standards aim at improving the international flow of capital (Chen et al 2015; IOSCO 1998). Prior research

[21] This chapter is based on Ghazizadeh, P., E. Peek, and D. Roesch. 2018. The effect of IFRS on ADR liquidity. *Working paper.*

shows that IFRS reporting indeed facilitates foreign direct investment (DeFond et al. 2011; Florou and Pope 2012; Yu and Wahid 2014). However, in addition to receiving foreign direct investments, firms access foreign capital by listing on foreign exchanges, either directly or through the use of depository receipts such as ADRs. Because firms trade off the costs and benefits of being traded on more than one exchange when deciding on such cross-listings, exchanges compete on benefits in the effort to attract listings (Pagano et al. 2002). Huddart et al. (1999) show that imposing strict disclosure requirements is one way to attract trading volume. Therefore, an important yet unanswered question is whether improvements in the accounting standards introduced by some exchanges affect the relative attractiveness of others and, in turn, cause trading to gravitate towards a few exchanges rather than catalyze global trading. Second, other factors than the quality of accounting standards, including the presence of knowledgeable investors and intermediaries, also play an important role in cross-listing decisions. The presence of these other factors causes firms to have heterogeneous cross-listing motives, which in turn determine whether information frictions between foreign and domestic investors matter for foreign investors' willingness to trade. In the presence of such heterogeneity, changes in the domestic information environment may not affect all cross-listed stocks equally and could therefore cause a shift in the type of firms that cross-list on U.S. exchanges.

We provide several reasons for why the adoption of IFRS in the domestic market could affect liquidity differences between the domestic and foreign trading venues. First, the adoption of IFRS in a cross-listed firm's domestic market could affect the trading frictions or trading cost differences that motivated the firm's initial choice to trade on multiple exchanges. More specifically, IFRS adoption mitigates the risk of insider trading or market manipulation in a company's domestic market (Brochet et al. 2013), which in turn would lower investors' benefits of receiving better protection on a U.S. exchange. Furthermore, an important factor creating a trading cost advantage for U.S. exchanges and drawing trading activity away from domestic exchanges is the greater informativeness of order flows under a U.S. GAAP reporting system than under a local GAAP reporting system. We expect that this trading cost advantage decreases after the adoption of IFRS in a firm's domestic market, when domestic order flows of peer stocks become more informative about cross-listed stocks' value changes. In line with models predicting the concentration of trading in one exchange (Pagano, 1989), these effects could lead to a pull of trading activity towards the domestic market, improving (deteriorating) liquidity of the securities on the cross-listed firm's domestic (U.S.) exchange. Alternatively, IFRS adoption could improve trading liquidity on the U.S. exchange by reducing the information processing costs to U.S. investors, amongst other reasons, because IFRS is arguably more similar to U.S. GAAP than most local GAAPs and the increased use of IFRS worldwide leads to greater accounting comparability across ADRs.

To test the effect of IFRS adoption on ADR liquidity, relative to domestic liquidity, we collect a sample of 239 firms with level 2 or 3 ADRs from 31 countries of which 27 have adopted IFRS between 2003 and 2012. We find that, in line with a reduction in the information advantage of domestic investors over U.S. investors, IFRS adoption improves the liquidity of ADRs. We find no evidence that the adoption of IFRS is associated with a significant increase in liquidity in the domestic exchange, which refutes the idea that a shift in trading towards the domestic exchange ensues following improvements in the domestic information environment. Our results further indicate that the improvement in the liquidity of ADRs depends on the quality of the domestic legal and regulatory institutions. Specifically, we find that mandatory IFRS adoption only affects the liquidity of ADRs of which the issuer's primary securities trade in one of the developed markets, which are argued to have higher quality institutions (Christensen et al. 2013). We also find that post-IFRS improvements in ADR liquidity increase as a function of direct measures of legal and regulatory quality. Tests aimed at identifying the source of the improvements do not support the idea that the superior quality of IFRS relative to the pre-existing domestic GAAPs drive the liquidity improvements, but rather point towards the scale benefits that follow from reducing the number of standards according to which cross-listed firms report. In an additional analysis, we find that the mandatory adoption of IFRS affects the liquidity of level 1 ADRs in a similar way as that of level 2 or 3 ADRs. We rule out that systematic differences in firm-specific characteristics between cross-listings from developed and emerging markets affect our results.

Our findings contribute to the literature in several ways. First, our findings add to the literature that investigates the consequences of mandatory IFRS adoption, primarily on domestic exchanges, by showing that improvements brought about by IFRS adoption can spillover to exchanges outside the IFRS environment. To the best of our knowledge, our study is the first to directly document that IFRS reporting improves liquidity of foreign securities traded on U.S. exchanges. As such, our results imply that IFRS adoption not only improves access to foreign capital through its positive effects on foreign direct investment (DeFond et al. 2011; Florou and Pope

2012; Yu and Wahid 2014) and access to other foreign markets within the IFRS environment (Chen et al. 2015), but also by improving access to a foreign market that has not adopted IFRS.

Second, the finding that changes in domestic accounting standards affect the liquidity of securities traded on a foreign exchange also adds to the literature that examines cross-listings. While it is clear that the relative liquidity an exchange can offer is a crucial determinant of its competitiveness, we show that improvements in accounting standards do not necessarily improve an exchange's liquidity advantage over competing exchanges. In fact, contrary to the predictions of gravitational pull models (Pagano, 1989), our results imply that the liquidity in foreign exchanges may improve (more) following domestic accounting improvements, which has important implications for 'race for the top' models (Huddart et al. 1999).

Third and finally, our results are of importance to the literature that investigates the effect of informational frictions on investors' capital allocation decisions. In particular, despite the diversification benefits that can arise from holding internationally balanced portfolios, prior research documents a preference of investors towards domestic securities (French and Poterba 1991; Ivkovic and Weisbenner 2005; Massa and Simonov 2006). Informational frictions, including a lack of familiarity with local GAAPs increasing information processing costs, are cited as factors that contribute to this so-called home bias (Kang and Stulz 1997; Brennan and Cao 1997). While we do not measure investor holdings, and therefore do not directly measure imbalances in holdings, our results indicate that IFRS adoption reduces the information disadvantage of U.S. investors relative to domestic investors. As such, our results add to studies that document a positive association between firms' accounting choices and (U.S.) institutional investor ownership in (non-U.S.) foreign firms (Bradshaw et al. 2004; Covrig et al. 2007; Khurana and Michas, 2011).

The remainder of this chapter is organized as follows. Section 2 reviews prior literature and develops our hypotheses. In Section 3 we describe our sample selection procedure, variable construction, and research design. Section 4 presents our empirical findings, and Section 5 concludes.

3.2 Hypothesis Development

3.2.1 Background

The past decade has seen an unparalleled move towards global standardization of accounting regulation, in particular for publicly listed companies. Following the introduction of International Financial Reporting Standards (IFRS) in Australia and the European Union in 2005, more and more countries worldwide have started to mandate the use of IFRS for public reporting, possibly stimulated by the coincidentally increasing network benefits of IFRS (Ramanna and Sletten 2014).[22] Prior literature broadly distinguishes two types of immediate benefits arising from such a move towards IFRS reporting.[23] A first benefit is that standardization of accounting regulation helps to increase the comparability of financial reports across countries. For example, Yip and Young (2012) show that following the

22 Large economies that have adopted IFRS include Brazil, Canada, Korea, Mexico, and Russia.

23 For a detailed review of academic research on the accounting and capital market effects of IFRS adoption see De George et al. (2016).

adoption of IFRS the mapping of economic events to accounting earnings has become more similar across peer firms from different European countries. Furthermore, using a similar approach towards measuring comparability Barth et al. (2012) find that non-US firms' accounting systems have become more comparable to those of U.S. peers after switching from local GAAP to IFRS reporting. This finding is not unexpected as, despite being different, IFRS and U.S. GAAP have similar conceptual bases and have exhibited substantial convergence over time.

A second benefit of IFRS reporting is that it improves corporate transparency. More than most local GAAPs around the world IFRS has as primary objective to ensure that financial statements provide decision-useful information to a firm's capital suppliers. Consequently, from the perspective of (equity) investors a switch from local GAAP to IFRS must help to make financial statements more investor-focused and, in turn, improve the information environment of the firm. Consistent with this idea, prior research finds that after a change to IFRS firms report earnings that elicit a greater market response upon announcement (Landsman et al. 2012) and are more value relevant (Aharony et al. 2010). Also, IFRS adopters see their analysts become more accurate in forecasting earnings (Byard et al. 2011; Horton et al. 2013; Tan et al. 2011). The evidence is, however, not unequivocal as some studies show that IFRS has not helped to improve earnings' predictive value for future earnings or cash flows (Atwood et al. 2011) and has increased earnings smoothing and accrual aggressiveness while decreasing loss recognition timeliness (Ahmed et al. 2013). Further, some academics have expressed concerns about implementation and enforcement differences across countries (Ball 2006; Soderstrom and Sun 2007).

The documented effects of IFRS on accounting quality and comparability ultimately have capital market consequences. That is, theory predicts that increased transparency and comparability helps to reduce problems of adverse selection and estimation risk in capital markets and, consequently, improves market liquidity and lowers firms' cost of capital (e.g., Diamond and Verrecchia 1991; Glosten and Milgrom 1985; Lambert et al. 2007). Confirming theoretical predictions, Daske et al. (2008) show that the mandatory adoption of IFRS led to an average increase in market liquidity in the European Union. IFRS adoption affects market liquidity especially if firms operate in an investor-oriented regulatory environment (Daske et al. 2008), have stronger incentives for transparent reporting, or have greater scrutiny from analysts (Daske et al. 2013). Likewise, prior research shows that both voluntary and mandatory adoption of IFRS has helped firms to reduce their cost of equity (Leuz and Verrecchia 2000; Li 2010). Further, the adoption of IFRS also promotes cross-border investing and listing activities. In accordance with the idea that IFRS increases comparability and hence reduces informational barriers between countries (Bae et al. 2008), research shows that the mandatory adoption of IFRS has increased cross-border investments (Beneish et al. 2015; Yu and Wahid 2014) as well as increased firms' propensity to issue shares in foreign markets, especially in other IFRS adoption countries (Chen et al. 2015), while reducing underpricing and increasing the proceeds of such issues (Hong et al. 2014).

In sum, there is a substantial body of evidence indicating that the adoption of IFRS has increased accounting transparency and comparability, which in turn has improved investors' willingness to trade and firms' access to capital. There is also some evidence that the adoption of IFRS has improved firms' access to foreign equity markets by removing the burden of preparing a second set of financial statements under foreign GAAP and by reducing foreign investors' information disadvantage over local investors. Research focusing on this effect has, however, not explicitly distinguished between equity cross-listings in the U.S. versus those in other countries.[24] Below we will argue that unique characteristics of the U.S. regulatory environment cause the adoption of IFRS to have distinctively different effects on U.S. and other cross-listings.

3.2.2 Main Hypothesis

Theory predicts that without the existence of trading frictions or trading cost differences across exchanges and with domestic investors being informationally advantaged over foreign investors, concentrating all trading on a firm's domestic exchange would achieve the greatest possible liquidity (Pagano 1989). Foreign exchanges can, however, attract trading activity by providing better investor protection against, for example, insider trading or market manipulation (Chowdhry and Nanda 1991). In fact, having embraced insider trading and market manipulation rules that are considered among the strictest in the world (Cumming et al. 2011), the U.S. stock exchanges (NYSE, NASDAQ, and AMEX) attract comparatively much trading in stocks originating from countries with underdeveloped markets and weak investor protection (Halling et al. 2008). This distinctive characteristic of the U.S. exchanges potentially has consequences for the effect of countries' adoption of IFRS on the benefits of U.S. cross-listings. That is, following Brochet et al.

[24] Univariate statistics reported by Chen et al. (2015) suggest that the number of U.S. cross-listings has not increased.

(2013), who find that transparency improvements following the mandatory adoption of IFRS has reduced insider trading profitability, we expect that the adoption of IFRS limits the accumulation of private information and, consequently, mitigates the risk of insider trading or market manipulation in a company's domestic market. This, in turn, would lower investors' benefits of receiving better protection on the U.S. exchanges and pull more trading activity towards the domestic market.

An alternative mechanism through which the adoption of IFRS in a firm's domestic market may affect the distribution of trading in the firm's stock across domestic and foreign exchanges is the changed informativeness of domestic order flows. As shown by Baruch et al. (2007), order flows of peer stocks can inform market makers about potential changes in a stock's value, thus helping to mitigate problems of adverse selection and stimulate trading (within a market). Basic logic predicts that the order flows of peer stocks are especially informative if trade orders are driven by timely and accurate information. Based on a broad range of research it is reasonable to assume that both U.S. GAAP and IFRS are superior, as compared to local sets of accounting standards, in providing such timely and accurate information publicly as well as in stimulating the production of private information (e.g., Byard et al. 2011). Under this assumption, an important factor creating a trading cost advantage for U.S. exchanges and drawing trading activity away from domestic exchanges is the greater informativeness of order flows under a U.S. GAAP reporting system than under a local GAAP reporting system. We expect that this trading cost advantage decreases after the adoption of IFRS in a firm's domestic market, when domestic order flows of peer stocks become more informative about cross-listed stocks' value changes.

While the above mechanisms imply that the adoption of IFRS would shift the distribution of trading in the firm's stock towards the domestic exchange, IFRS adoption could also increase trading volume on the U.S. exchange. Specifically, prior literature shows that investor portfolios are disproportionately tilted towards the stocks of firms from their own countries (French and Poterba, 1991), and partially ascribe this to informational advantages of domestic over foreign investors (Halling et al., 2008). Relatedly, Ivkovic and Weisbenner (2005) and Massa and Simonov (2006) show that investors earn excess returns when they invest locally, which implies that the bias towards more approximate stocks reflects an informational advantage rather than a behavioral bias. As accounting information is a primary source of value-relevant information, the various accounting standards that cross-listed firms use to prepare their financial statements pose a cost to trading on the U.S. exchange. That is, a U.S. investor looking to diversity his portfolio internationally by investing in ADRs must incur a cost when learning an additional set of accounting standards, which in turn should render him more reluctant to invest. In line with this argument, Yu and Wahid (2014) show that differences in accounting standards affect investor demand by imposing greater information-processing costs on those less familiar with the reporting standards. Further, Lundholm et al (2014) show that firms cross-listed on U.S. exchanges in fact respond to a perceived reluctance on the part of U.S. investors and attempt to lower U.S. investors' information disadvantage by providing clearer and more concrete disclosures. The adoption of IFRS by cross-listed firms likely reduces U.S. investors' information-processing costs because (a) the marginal costs of learning one broadly adopted set of accounting standards such as IFRS are much lower than those of learning various domestic GAAPs and (b) IFRS and U.S. GAAP exhibit strong similarities. A potentially moderating factor is that, unlike when reporting under domestic GAAP, cross-listed firms that report under IFRS are not required to provide additional U.S. GAAP reconciliation disclosures. Nonetheless, on a net basis, we expect that the adoption of IFRS in a firm's domestic market reduces U.S. investors' information disadvantage, which in turn may increase trading activity on the U.S. exchange.

In summary, as there are arguments for both a gravitational pull of trading activity to the domestic exchange and increased trading activity on the U.S. exchange, we do not take an a priori stance on the direction of the effects of IFRS adoption on a firm's domestic and ADR liquidity and state our hypothesis in null form:

Hypothesis: *IFRS adoption in a firm's domestic market is not associated with liquidity differences between the firm's domestic securities and its U.S. ADR securities.*

3.3 Methodology

3.3.1 Sample Selection and Liquidity Measurement

Our main sample consists of 239 level 2 or 3 ADR securities traded on one of the U.S. exchanges during the years 1998 and 2015, each matched with the primary listing of the issuer of the ADR's underlying shares. In an additional analysis, we analyze 572 level 1 ADR securities and their matched primary listings. To construct a comprehensive sample of ADRs we combine data from various sources. Following prior studies, we obtain data from the August 2017 versions of the Bank of New York, Deutsche Bank, and JP

Morgan ADR websites. Because these websites primarily focus on currently listed ADRs and thus create a risk of survivorship bias, we combine these data with ADR data from CRSP, Eikon, Datastream, and the U.S. SEC's 1996 – 2015 lists of International Registered and Reporting Companies, which all include delisted ADRs. We distinguish level 1 from level 2 or 3 ADRs, for each sample year separately, based on whether the issuer of the ADR's underlying shares (hereafter referred to as the issuer or the firm) is classified as an 'International Registered and Reporting Company' by the U.S. SEC.[25] We identify each issuer's primary listing using Datastream and require that the issuer has its primary listing in its country of domicile.

Accounting data come from Worldscope and price and volume data, both for the ADRs and the issuers' primary listings, come from Datastream. We establish the accounting standards used by each issuer, prior to the adoption of IFRS, using Worldscope's classification, where we categorize U.S. GAAP and IFRS as "International GAAP" and all other accounting standards as "local GAAP".[26] Firms that report under International GAAP prior to their domestic market's adoption of IFRS are considered voluntary adopters. We measure liquidity by firm-quarter using the following three measures: the natural logarithm of median trading volume (in U.S. dollar) during the quarter; the natural logarithm of the median of Amihud's (2002) illiquidity measure during the quarter times minus one, and the proportion of zero-return trading days during the quarter times minus one. Whereas each of these variables has

[25] We classify years in which ADRs change from level 1 to level 2/3 or vice versa as transition years and exclude these years from the analysis. Further, because the SEC's list of International Registered and Reporting Companies does not contain data for 1999, we classify 1999 as a transition year if ADRs change their registration from 1998 to 2000.

[26] To correct a small number of potential irregularities Worldscope's accounting standards classification, we assume that issuers do not switch back from "International GAAP" to local "GAAP".

been widely used in prior research, it is reasonable to assume that they measure liquidity with noise. In our empirical tests, we therefore us a factor score of the three variables, which we label liquidity factor, as our main measure of liquidity. The factor analysis yields one factor with an eigenvalue greater than 1.0, both in the domestic market (2.00) and the ADR market (1.82).

Table 1 provides an overview of the sample composition by country. The sample consists of 12,143 firm-quarters from 239 unique firms from 31 countries of which 27 have adopted IFRS between 2003 (Singapore) and 2012 (Argentina and Taiwan). While the number of countries that adopted IFRS during our sample period far exceed those that have not, the distribution of observations before and after mandatory adoption is more balanced (8,003 and 4,140 respectively). Approximately 10 percent (i.e., 802 out of 8,003) of the firm-quarter observations before mandatory adoption are observations from firms that voluntary adopted International GAAP (IFRS or U.S.GAAP). Table 1 further indicates that 13 domestic markets are part of the European Union and 18 domestic markets are located in a developed country. The last three columns of Table 1 provide information on institutional characteristics of issuers' countries of domicile. Specifically, the columns report a measure of the magnitude of accounting differences between IFRS and pre-existing GAAP (Bae et al., 2008) as well as percentile-ranked measures of the strength of the judicial system (*Rule of law*) and regulatory environment (*Regulatory quality*), taken from the Worldbank's Worldwide Governance Indicators. Two institutional measures are significantly associated with the indicator variables for E.U. membership and market development. In particular, the (untabulated) correlation coefficients between the E.U. indicator and, respectively, *Rule of law* and *Regulatory quality* are 0.55 and 0.50. Further, the correlation coefficients between the developed market indicator and, respectively, *Rule of law* and *Regulatory quality* are 0.80 and 0.76.

[Insert Table 1 here]

Panel A of Table 2 displays descriptive statistics of our liquidity measures, issuers' market capitalization, and issuers' domestic daily return volatility for the full sample. The means and medians of dollar trading volume indicate that most trading in sample firms' securities occurs in the home market. In fact, the sample mean (median) of average daily volume during the quarter amounts to \$115 million (\$18 million) in issuers' home market, compared to \$20 million (\$3 million) in the ADR market. Trading also generates a smaller price impact in the home market. The mean Amihud illiquidity measure is 0.155 in the home market versus 0.189 in the ADR market. Surprisingly, home markets appear less liquid, on average, when measured as the proportion of trading days in a quarter with zero returns. This finding confirms the notion that the three measures reflect different dimensions of liquidity and provides further support for using the measures' factor score as a comprehensive measure of liquidity. The descriptive statistics also show that all variables are positively skewed, which motivates the use of log-transformed measures in the empirical tests. Panel B of Table 2 reports the correlations among the firm characteristics. The Pearson correlation coefficients indicate that the trading volume and liquidity on the U.S. and home exchanges are positively correlated. Furthermore, the Amihud and zero-return-based measures of illiquidity are positively correlated with each other and negatively correlated with the trading volume measure. The correlations further indicate that larger firms are more liquid and have less volatile returns.

[Insert Table 2 here]

3.3.2 Research Design

Our empirical analysis focuses on assessing the effect of IFRS adoption on liquidity differences between primary securities traded on a domestic exchange and ADR securities traded on a U.S. exchange. To do so, we first measure the effects of IFRS adoption on the liquidity of securities traded on the domestic and U.S. exchanges separately and then determine the effect on liquidity differences by comparing the observed effects between the two exchanges. While many studies investigate the capital market consequences of IFRS adoption in domestic markets, their focus is not on cross-listed firms. In prior studies, cross-listed firms are either excluded from the sample (e.g., Christensen et al., 2013) or the reported tests cannot be used to clearly isolate the effect of IFRS adoption for these firms (e.g., Daske et al., 2008; Li, 2010; Daske et al., 2013)[27]. However, as described in the previous section, the consequences of IFRS adoption on the liquidity of the individual securities of firms cross-listed in the U.S. are *ex-ante* not obvious. In particular, depending on which effect dominates, we could find an increase in ADRs' relative liquidity because of a reduction in domestic investors' information advantage or a decrease in ADRs' relative liquidity as a result of domestic markets' increased gravitational pull. Alternatively, if cross-listed firms

[27] For instance, in testing the effect of IFRS on cost of equity capital, Li (2010) controls for the effect of firms being cross-listed and the results indicate that they have – *on average* – a lower cost of equity capital (Table 4, p. 621). However, our focus is on the *change,* rather than the difference in level across groups, of a dependent variable, which would have required the cross-listing variables to be interacted in the regression specification (Eq (1), p. 614). One exception is Daske et al (2008) who report untabulated results that suggest IFRS adopters that are cross-listed on U.S. exchanges experience lower, if any, liquidity improvements (p. 1120).

committed to high-quality disclosure and enforcement prior to the adoption of IFRS, we may not find a material incremental effect of adoption.

In order to isolate the effects of IFRS adoption, we employ a difference-in-difference design in which cross-listed firms from countries that have not (yet) adopted IFRS serve as a control group for cross-listed firms from countries that have. We prefer this approach over comparing cross-listed firms to their home market peers because under our approach treatment firms and control firms share the partially unobservable motivation to cross-list[28]. We estimate the following model for domestic and ADR securities separately:

$$\begin{aligned} \text{Liquidity factor}_{ijt} = {} & \beta_1(\text{Post mandatory IFRS}_{jt}) + \beta_2\text{Voluntary adopter}_i + \\ & \beta_3(\text{Post mandatory IFRS}_{jt} \times \text{Voluntary adopter}_i) + \\ & \textstyle\sum \beta_k \text{Controls}_k + a_i + \gamma_t + \varepsilon_{ijt} \end{aligned} \qquad (1)$$

where *Liquidity factor*$_{ijt}$ is the liquidity factor score, as defined earlier, for security i, from country j, in year-quarter t; *Post mandatory IFRS*$_{jt}$ is an indicator variable that equals 1 if country j has adopted IFRS by year-quarter t, and *Voluntary adopter*$_i$ is an indicator variable that equals 1 if firm i has adopted International GAAP before the mandatory adoption of IFRS by its country of domicile. *Controls*$_j$ is a vector of two control variables: the natural logarithm of a firm's market capitalization at the end of the quarter (*ln[Market capitalization]*) and the natural logarithm of one plus the volatility of a firm's domestic daily returns during the quarter (*ln[1 + Return volatility]*). Further, a_i are firm fixed effects, and γ_t are year-quarter fixed effects. The firm fixed effects control for

[28] Note that our implicit assumption is that the unobservable motivation is homogeneous across cross-listed firms from different countries – which probably is not the case. However, we believe this poses a smaller problem than the alternative. We further address this issue in the tests reported in Table 6.

unobserved, time-invariant differences across firms, including any capital market differences across cross-listed firms from adopting and non-adopting countries prior to IFRS adoption, while year-quarter fixed effects control for capital market changes unrelated to the adoption of IFRS (e.g., general trends). The coefficient of interest is β_1, which captures the *incremental* change in liquidity of the home market (ADR) securities of cross-listed mandatory adopters, relative to the change in liquidity of the home market (ADR) securities of cross-listed non-adopters. In all regressions, we cluster standard errors at the country and calendar quarter level.

3.4 Results

3.4.1 Regression analysis

Table 3 presents the results of the regression analysis of the effect of IFRS adoption in a firm's domestic market on domestic and ADR liquidity. Models 1 and 2 display the results of the main regression specification, where the dependent variable is the liquidity factor score and the sample consist of, respectively, ADR securities (Model 1) and domestic securities (Model 2). Column 3 reports differences between the coefficients of the ADR securities sample and those of the domestic securities sample.[29] In line with the argument that IFRS reduces the information disadvantage of investors trading on the U.S. exchange, we find that the coefficient on *Post mandatory IFRS x Mandatory adopter* is significantly positive (0.140, $p<0.05$). Because an increase

[29] Note that, because all explanatory variables in columns 1 and 2 are the same, the coefficient differences and t-statistics that we report in column 3 are equal to the coefficient estimates that we obtain if we regress ADR versus home market differences in the liquidity factor on all explanatory variables.

in ADR liquidity in the U.S. market does not automatically preclude a shift of trading towards the domestic market we focus our hypothesis test on the effect differences between the ADR market and the home market, reported in column 3. Model 2 shows that mandatory IFRS adoption is not associated with a significant increase in liquidity in the domestic exchange (0.054, p=0.26). Nevertheless, we find that the difference in liquidity across the two exchanges, as reported in column 3, is not significantly different from zero (0.086, p=0.21). This finding leads us to conclude that, on average, mandatory adoption of IFRS in the domestic market is not associated with a global shift of trading from the ADR exchange towards the domestic exchange or vice versa. For voluntary adopters we draw the same conclusion. That is, the directions and magnitudes of the coefficients capturing the effect of mandatory IFRS adoption on voluntary adopters' ADR and home market liquidity are similar. Further, the coefficients are not statistically significant in either of the two models and they are not statistically different from each other (0.132, p=0.31).

[Insert Table 3 here]

The finding that on a global scale mandatory IFRS adoption is not associated with changes in ADR versus domestic liquidity differences motivates us to refine the analysis. An important reason for doing so is that the effect of the IFRS mandate is arguably conditional on the prevailing regulatory quality of the country in which it is adopted (Christensen et al., 2016). In particular, the effect of IFRS, which is deemed of higher quality than most sets of local accounting standards, could be larger in countries with lower-quality regulatory institutions before its implementation. Alternatively, it could be argued that the presence of high-quality regulatory institutions is a necessary condition for the effects of IFRS adoption to materialize. Therefore, pooling all IFRS adoptions, without distinguishing among regulatory environments, could mask important cross-section differences. To address this issue, we separately examine the effect of IFRS adoption for cross-listed firms from European Union (E.U.) countries and firms from non-E.U. countries. By comparing E.U. with non-E.U. firms we build on prior research arguing that the legal and regulatory institutions in E.U. countries are of higher quality than those of other IFRS adopting countries (Christensen et al., 2013). In Models 4 and 5, we interact the explanatory variables of interest with two binary indicators: *EU* and *non-EU*. We further refine the *non-EU* group into *non-EU developed* and *non-EU emerging*, and report these results in Models 7 and 8.

A second potential concern is that other regulatory or enforcement changes enacted around the mandatory adoption of IFRS drive our results (Christensen et al., 2013, 2016). To alleviate this concern, we exploit the the fact that voluntary adoption of international GAAP by some of the firms in our sample is less likely to systematically coincide with country level regulatory and enforcement changes. More specifically, we include an additional indicator variable, labeled *Post voluntary International GAAP*, that reflects the timing of the voluntary switch to IFRS or USGAAP (i.e., takes on the value of one in and after the first quarter with international GAAP reporting) and interact this variable with the region indicator variables discussed above.

The results of the additional tests indicate that the effect of IFRS on ADR versus domestic liquidity differences indeed depends on the quality of the prevailing legal and regulatory institutions. More specifically, we find that the positive effect of IFRS adoption in E.U. countries on the liquidity of

ADRs is both economically and statistically more significant than in the pooled sample (Model 4 vs Model 1), while the implementation of IFRS has no effect in the combined group of developed and emerging non-E.U. countries. As before, we find that the adoption of IFRS affects liquidity only in the U.S. market; its effect on domestic liquidity remains insignificant. The IFRS-induced increase in ADR liquidity for E.U. issuers is of such magnitude that the difference in the liquidity effect of IFRS between the two trading venues is statistically significant (0.261, $p<0.10$). The economic interpretation of this difference is that the mandatory adoption of IFRS reduces the liquidity gap between the ADR and domestic securities for E.U. firm by one-fourth of the standard deviation in *Liquidity factor*, which underlines the economic significance of the observed IFRS effect. The results further dismiss concerns that other contemporaneously implemented regulatory or enforcement changes drive our findings. In particular, the coefficient on *Post voluntary International GAAP x EU* indicates that cross-listed firms from the E.U. that voluntarily adopted IFRS also experience an increase in ADR liquidity that is significantly greater than the concomitant change in domestic liquidity (0.240, $p<0.10$). In line with the notion that the scale benefits that accrue to U.S. investors from learning IFRS materialize when more firms report according to the new standards – that is, at the time of the mandatory implementation of IFRS – we find an additional increase in the ADR liquidity of voluntary adopters at the time of mandatory adoption (0.262, $p<0.05$). The incremental change in ADR versus domestic liquidity differences following the mandatory adoption of IFRS is positive but not significantly different from zero for voluntary adopters (0.266, $p=0.22$).

When we further split up the group of non-E.U. firms based on their domestic markets' level of development in Models 7 and 8, the results we obtain by and large echo the above findings. In particular, we find that mandatory IFRS adoption has a similar effect on the ADR versus domestic liquidity gap for E.U. firms and non-E.U. firms from developed markets (0.272, $p<0.10$ and 0.337, $p<0.10$, respectively). Perhaps most strikingly, we find a significant decrease in the liquidity of the ADR securities of firms from emerging non-EU markets (-0.119, $p<0.10$). Considering that the coefficient on *Post mandatory IFRS × Mandatory adopter × non-EU emerging* in Model 8 is not significantly positive, which would have reflected a shift in trading, we interpret the decrease in ADR liquidity as confirming the argument of Shleifer (2005), Bhattacharya and Daouk (2009) and Ball (2006), that in the absence of effective regulatory and legal institutions, new regulation could facilitate abuse. Overall, these results reinforce the conclusion that IFRS reporting reduces the information disadvantage of investors trading on a U.S. exchange, but only when the institutions that must safeguard its correct implementation are of sufficiently high quality. We fail to find evidence that is consistent with the idea that improvements in firms' information environment result in a shift of trading to the more liquid domestic markets.

3.4.2 Institutional Characteristics

The results reported in Table 3 show that the effect of IFRS adoption on liquidity is restricted to ADR securities and conditional on an issuer's country of domicile. In this section, we further explore potential sources of these findings. More specifically, the improved liquidity of ADR securities could arise from two, not mutually exclusive, sources: (1) a reduction in the informational disadvantage of U.S. investors, even in the absence of any improvements in the quality of reporting, and (2) the higher quality of IFRS in comparison with domestic GAAPs. In an additional analysis, we interact the main explanatory variables of interest with the Bae et al. (2008) measure of the magnitude of accounting differences between IFRS and local GAAP. If the documented effects are associated with improvements brought about by IFRS, we expect that liquidity improves especially for firms from countries with a larger pre-existing distance to IFRS. Furthermore, whereas in the previous analysis we attributed regional differences in the effect of IFRS to differences in the quality of legal and regulatory institutions, we now complement this analysis by interacting the main explanatory variables of interest with two direct measures of institutional quality, *Rule of law* and *Regulatory quality*.

We report the results of these additional tests in Table 4. For reasons of brevity we only report the differences between the coefficients of the ADR liquidity regression and those of the domestic liquidity regression. In column 1 the coefficient difference of *Post mandatory IFRS × Mandatory adopter × Rule of Law,* which reflects whether the mandatory IFRS effect on (mandatory adopters') ADR versus domestic liquidity differences varies with the strength of the domestic judicial system, has the expected sign but is not statistically significant (0.533, p=0.23). Similarly, the coefficient difference of *Post mandatory IFRS × Mandatory adopter × Regulatory quality* is positive but not statistically significant (0.538, p=0.25). We find similar results for the voluntary adopters at the time of mandatory adoption. However, this is less surprising given that the results reported in Table 3 suggest that the liquidity improvement for these firms mainly occurs at the time of their voluntary switch to International GAAP. In line with this idea, we find that the coefficient difference of *Post voluntary International GAAP x Regulatory quality* is significantly positive (0.683, p<0.05). Focusing on ADR liquidity in untabulated regressions, we do find evidence that post-IFRS improvements in ADR liquidity positively depend on the strength of legal and regulatory institutions in an issuer's home country. More specifically, both for *Rule of Law* and *Regulatory quality*, we find that the coefficients on *Post mandatory IFRS x Mandatory adopter x Institutional characteristic* are positively significant (0.699, p<0.01 and 0.748, p<0.05, respectively). We also find that, the effects of *Rule of Law* and *Regulatory quality* on the association between IFRS and voluntary adopters' ADR liquidity are only significantly positive at the time of the voluntary switch. This suggests that voluntary switches by firms from countries with strong legal and regulatory institutions are considered more credible than those by firms from countries with weak institutions, which is *ex-ante* not obvious (Daske et al., 2013).

[Insert Table 4 here]

In column 3 the coefficient difference of *Post mandatory IFRS x Mandatory adopter x Institutional characteristic,* which measures whether the effect of mandatory IFRS adoption on ADR versus domestic liquidity differences increases with the distance between local GAAP and IFRS, is not significantly different from zero. However, the coefficient difference of *Post mandatory IFRS x Voluntary adopter x Institutional characteristic* is significantly positive, implying that, as opposed to mandatory adopters, the effect of mandatory IFRS adoption on voluntary adopters' liquidity gap increases in local GAAP's distance from IFRS. With respect to the effects of the voluntary adoption of International GAAP, we find that distance to IFRS reduces the (positive) effect of voluntary IFRS adoption on differences between ADR and domestic liquidity (-0.155, $p<0.10$). When we use ADR liquidity as the dependent variable, we find similar results (not tabulated), with the exception that the coefficient of *Post voluntary International GAAP × Institutional characteristic* is not significant. Overall, our results indicate that IFRS adoption has beneficial capital market consequences in the ADR market, but that these benefits depend on the quality of the firms' domestic legal and regulatory institutions. Furthermore, our test do not confirm that these improvements arise from the higher quality of IFRS compared to local GAAP, rendering reductions in information processing costs as the most likely source.

3.4.3 Firm Level Characteristics

A potential concern is that the previously documented differences in the liquidity effects of IFRS between firms from developed countries and firms from emerging countries is not (only) driven by differences in domestic institutional quality but (also) by systematic differences in the type of firms that cross-list on U.S. exchanges. That is, while seeking a U.S. listing to reap bonding benefits that arise from a commitment to higher regulatory standards is more likely to be a primary motive for firms from emerging markets, prior literature documents several other motives that vary in importance across firms and industries[30] and, importantly, could be differentially correlated with the consequences of IFRS implementation. Because the identification of our

[30] Examples include raising capital, seeking expertise of knowledgeable analysts and investors, and product market benefits. We refer the reader to Pagano et al. (2002) for an overview (Table 1, p. 2654).

previous tests relies on variation at the region of origin level, these tests could suffer from measurement error problems.

As a first step to address this issue, we investigate whether the firms in our sample indeed differ systematically across regions of origin. More specifically, we estimate a probit model to discern which factors distinguish cross-listings from developed countries from those from emerging countries. We include the following variables as regressors: issuers' pre-IFRS average of the foreign sales to sales ratio, the pre-IFRS average of U.S. peers' share in the issuers' ICB Supersector (labeled *Pre-IFRS average of U.S. industry importance*), the pre-IFRS average of the explanatory value of U.S. market index returns for the issuers' domestic stock returns (*Pre-IFRS average of RSQUS*), and a pre-IFRS high-tech industry indicator. We further include the pre-IFRS average of *Liquidity factor* in the issuer's domestic market, the pre-IFRS average of daily return volatility in the issuer's domestic market, and the pre-IFRS average of the natural log of the issuer's market capitalization as control variables.

The results of our probit regression, reported in Table 5 (Model 2), show that the probability that a cross-listed firm comes from a developed country is positively related to the proportion of its foreign sales to total sales and the importance of U.S. peers in its industry. This finding suggests that cross-listed firms from developed markets have an international focus and come from industries that are familiar to U.S. investors. The coefficient on the high-tech indicator is not significant, which refutes the idea that U.S. exchanges attract high-tech companies from European countries (Pagano et al. 2002). We find that the co-movement of firms' domestic stock returns with U.S. market returns is negatively related to the likelihood that a listing is from a developed market. This finding is in line with the idea that firms from developed countries that cross-list on U.S. exchanges seek a broadening of their investor base by offering attractive diversification opportunities to foreign investors. We further find that firms cross-listed on U.S. exchanges are more likely to be from developed markets when domestic securities are more liquid. We run the same regression using as the dependent variable an indicator for whether the issuer comes from an E.U. member state or a non-E.U. country (Model 1). Because the results of Models 1 and 2 are very similar, and probability estimates of these models are highly correlated, we focus on Model 2 in the remainder of our analysis.

[Insert Table 5 here]

The probit regression results suggest that investor base optimization is more likely to motivate a U.S. cross-listing for firms from developed countries than for those from emerging markets, whereas the pursuit of a more liquid market is more likely to drive U.S. cross-listing for firms from emerging markets. To the extent that the implementation of IFRS is related to these differences, interpreting the results of Table 3 to reflect (solely) institutional characteristics would be erroneous[31]. For instance, it is possible that the comparability benefits arising from the implementation of IFRS have a larger impact on firms that operate more internationally. As in our sample these firms are more likely to be from developed countries, this, rather than superior legal and regulatory institutions, could (partially) account for the positive effect of IFRS on the liquidity of ADRs. Furthermore, explicitly accounting for these other motives should improve the estimation of the coefficients of our regressions.

[31] Note that the test reported in Table 4 mitigate this concern.

Given the above, we rerun our main regression analyses after controlling for the estimated propensity of a firm being from a developed country and interactions of this propensity and the main explanatory variables of interest. This robustness analysis, which we report in Table 6, does not does not market change our main findings, i.e., those that relate to the effect of the mandatory adoption of IFRS. That is, we still find that mandatory IFRS adoption positively affects the liquidity of ADRs from developed countries, such that the difference in liquidity between the ADR market and the domestic market significantly increases (0.530, $p<0.01$ for E.U. markets and 0.549, $p<0.01$ for non-E.U. developed markets). We no longer find a significant negative effect of IFRS adoption on the liquidity of ADRs from emerging markets. We do find that after controlling for the estimated propensities, the results for voluntary adopters change. That is, we no longer find any significant effects for these firms at the time of the adoption, while in all regions the liquidity in their domestic exchanges improve at the time of their voluntary adoption.

[Insert Table 6 here]

3.4.4 Level 1 ADRs

The previous analyses suggest that IFRS adoption has a comparatively stronger effect on the information environment of U.S. investors than on that of domestic investors, thus creating a comparatively larger liquidity improvement in the ADR market. If the observed improvement in ADR liquidity indeed results from a reduction in U.S. investors' information processing costs, as we argued earlier, rather than, for example, an increase in the effectiveness of U.S. enforcement, we would expect to observe a similar improvement in the liquidity of level 1 ADR securities. Issuers of level 1 ADR securities are exempt from most SEC reporting requirements and thus subject to a significantly weaker enforcement regime in the U.S. If our prior findings are primarily driven by a change in U.S. enforcement, we would expect that the effect of IFRS adoption on ADR liquidity is smaller for level 1 ADRs than for level 2 or 3 ADRs.

To examine this issue, we re-run our main empirical tests using a sample of 572 level 1 ADR securities and their matched primary (domestic) listings. We refer to section 3 for a discussion of how we construct this sample. Column 1 of Table 7 displays the results of a replication of the analysis presented in column 9 of Table 3. When analyzing level 1 ADR liquidity, we find that the mandatory adoption of IFRS has a significantly greater influence on ADR liquidity than on domestic liquidity, both for issuers from E.U. countries (0.497, $p<0.01$) and for issuers from non-E.U. developed countries (0.312, $p<0.05$). Mandatory IFRS adoption also improves ADR liquidity more than domestic liquidity for E.U.-domiciled voluntary adopters of International GAAP (1.093, $p<0.01$). Furthermore, voluntary adoption of International GAAP improves ADR liquidity more than domestic liquidity for issuers from non-E.U. countries, both developed (0.199, $p<0.05$) and emerging (0.421, $p<0.01$).

[Insert Table 7 here]

Replications of the three analyses reported in Table 4 show that *Rule of law* and *Regulatory quality* are positively associated with the extent to which mandatory IFRS adoption improves level 1 ADR liquidity (more than domestic liquidity) (1.157, $p<0.01$ for *Rule of law* and 1.128, $p<0.05$ for *regulatory quality*). This finding is consistent with earlier findings in a sample of level 2 or 3 ADRs and confirms that mandatory IFRS adoption improves U.S. investors' information environment especially if the domestic market's legal and regulatory institutions are of high quality. We find that voluntary adoption of International GAAP improves ADR liquidity (more than domestic liquidity) especially if the judicial system in the voluntary adopter's domestic market is weak (-1.005, $p<0.05$) or if the distance between the voluntary adopter's local GAAP and IFRS is large (-0.241, $p<0.01$). This finding suggests that better information can help U.S. investors overcome their resistance to investing in level 1 ADR securities of issuers from countries with weak information or regulatory environments. All in all, the results of our analysis of level 1 ADR securities confirms that the information effects rather than the enforcement effects of IFRS adoption drive the observed improvement in (relative) ADR liquidity.

3.5 Conclusion

The implementation of IFRS by more than 120 countries constitutes one of the largest accounting regulation changes to date. While various effects of IFRS on the adopting markets have been examined, little is known about its effect on markets that have not adopted these standards. This chapter examines the effect of the mandatory adoption of IFRS in the domestic market of a U.S. cross-listed firm on the liquidity of its securities traded on the U.S. and domestic exchange. Using a sample of 239 firms with level 2 or 3 ADRs from 31 countries, of which 27 have adopted IFRS between 2003 and

2012, we find that IFRS adoption improves the liquidity of ADRs more than that of domestic securities, which is in line with the notion that IFRS reduces the information disadvantage of investors trading on U.S. exchanges. Our results further indicate that the improvement in the liquidity of ADRs depends on the quality of the domestic legal and regulatory institutions. Tests aimed at identifying the source of the improvements do not confirm that the presumably superior quality of IFRS relative to the pre-existing local GAAPs affect the liquidity improvements, but rather point towards a reduction in U.S. investors' information processing costs that results from the positive effect of IFRS adoption on accounting comparability. Collectively, our results imply that the adoption of IFRS in a U.S. cross-listed firm's domestic market improves access to foreign markets that have not adopted the mandate and potentially U.S. investors' capital allocation decisions, especially for those restricted to invest in securities on U.S. exchanges. Our findings further speak to the role of accounting standards in the competition between stock exchanges.

Table 1. Sample

				Number of ADRs		Number of observations			Institutional characteristics		
Country	Mand. IFRS date	EU	Developed	Total	Of which: voluntary adopters	Before mand. IFRS adoption	Of which: voluntary adopters	After mand. IFRS adoption	Rule of law (percentile)	Regulatory quality (percentile)	Distance from IFRS
Argentina	Dec-12	0	0	12	1	568	16	137	0.23	0.22	14
Australia	Dec-05	0	1	12	0	285	0	180	0.91	0.86	4
Brazil	Dec-09	0	0	30	1	998	8	644	0.33	0.48	11
Chile	Dec-09	0	0	11	0	474	0	251	0.82	0.83	13
China		0	0	2	2	72	48	0	0.25	0.35	9
Denmark	Dec-05	1	1	1	1	33	4	40	0.93	0.90	11
Finland	Dec-05	1	1	3	2	84	32	48	0.94	0.90	15
France	Dec-05	1	1	17	0	421	0	274	0.84	0.74	12
Germany	Dec-05	1	1	5	3	133	60	68	0.89	0.83	11
Greece	Dec-05	1	1	1	1	30	20	16	0.67	0.67	17
Hong Kong	Dec-05	0	1	7	0	159	0	168	0.83	0.92	3
India		0	0	16	4	543	65	0	0.49	0.27	8
Indonesia		0	0	2	0	138	0	0	0.16	0.22	4
Ireland	Dec-05	1	1	5	0	117	0	134	0.88	0.88	1
Israel	Dec-08	0	0	4	3	164	88	112	0.70	0.70	6
Italy	Dec-05	1	1	6	1	152	12	168	0.62	0.70	12
Japan		0	1	12	8	527	349	0	0.82	0.71	9
Mexico	Dec-11	0	0	17	2	767	8	226	0.29	0.54	1

Netherlands	Dec-05	1	1	7	0	196	0	76	0.91	0.91	3
Norway	Dec-05	1	1	3	2	43	12	48	0.93	0.79	7
Philippines	Dec-05	0	0	1	0	33	0	40	0.30	0.38	10
Portugal	Dec-05	1	1	2	1	66	8	40	0.79	0.74	13
Singapore	Dec-03	0	1	1	1	18	12	20	0.86	0.92	1
South Africa	Dec-05	0	0	6	1	156	20	174	0.50	0.62	0
South Korea	Dec-11	0	0	7	0	287	0	112	0.70	0.64	6
Spain	Dec-05	1	1	5	1	165	20	140	0.80	0.77	16
Sweden	Dec-05	1	1	2	0	66	0	44	0.92	0.84	10
Switzerland	Dec-05	0	1	1	1	16	4	40	0.92	0.87	12
Taiwan	Dec-12	0	0	6	1	286	4	72	0.68	0.71	6
Turkey	Dec-07	0	0	1	1	31	4	32	0.48	0.51	14
United Kingdom	Dec-05	1	1	34	1	975	8	836	0.90	0.90	1
Total				239	39	8,003	802	4,140			

This table provides an overview of the sample composition and the country-level variables used in the analyses. The main sample consists of 12,143 firm-quarters from 239 level 2 or 3 ADR securities traded on one of the U.S. exchanges during the years 1998 and 2015 with available price and volume data in Datastream and accounting data in Worldscope. The table reports the number of unique firms and voluntary adopters per country, and the distribution of firm-quarter observations before and after IFRS adoption. For each country, the table presents information regarding the date of mandatory adoption of IFRS (blank when IFRS has not been adopted until 2016), membership of the European Union (1 indicates membership, 0 otherwise), whether the country is classified as a developed country (1 indicates the country being from developed markets, 0 otherwise), the average percentile score of the rule of law index taken from the Worldbank's Worldwide Governance Indicators, the average percentile score of the regulatory quality index taken from the Worldbank's Worldwide Governance Indicators, and the magnitude of accounting differences between IFRS and the country's pre-existing GAAP (Bae et al., 2008).

Table 2. Descriptive Statistics and Correlations

Panel A: Descriptive statistics

Variable	Mean	StDev.	P1	Q1	Median	Q3	P99
Volume in USD mln (ADR)	20.237	45.420	0.003	0.430	2.820	15.523	274.788
Volume in USD mln (Home market)	115.058	378.092	0.002	1.272	17.511	81.924	3,282.762
Amihud illiquidity (ADR)	0.189	0.832	0.000	0.001	0.004	0.030	6.768
Amihud illiquidity (Home market)	0.155	0.718	0.000	0.000	0.001	0.009	5.842
Zero returns (ADR)	0.030	0.061	0.000	0.000	0.000	0.045	0.385
Zero returns (Home market)	0.095	0.122	0.000	0.030	0.062	0.108	0.769
Total market capitalization in USD mln	22,342.4	33,619.3	45.8	1,898.1	8,257.5	26,293.5	173,138.6
Return volatility (Home market)	0.022	0.012	0.007	0.014	0.019	0.027	0.072

Panel B: Correlations (Pearson)

Variable	(2)	(3)	(4)	(5)	(6)	(7)	(8)
(1) ln[Volume] (ADR)	0.452	-0.897	-0.424	-0.205	-0.238	0.621	-0.076
(2) ln[Volume] (Home market)		-0.472	-0.931	-0.157	-0.351	0.712	-0.134
(3) ln[Amihud illiquidity] (ADR)			0.500	0.357	0.305	-0.715	0.228
(4) ln[Amihud illiquidity] (Home market)				0.207	0.461	-0.725	0.265
(5) Zero returns (ADR)					0.309	-0.315	0.177
(6) Zero returns (Home market)						-0.381	0.148
(7) ln[Market capitalization]							-0.339
(8) Return volatility (Home market)							

This table presents descriptive statistics (Panel A) and Pearson's correlation coefficients (Panel B) of the domestic and ADR volume and liquidity proxies and firm-level control variables for the sample of 12,143 firm-quarters of 239 level 2 or 3 ADR issuers during the years 1998 and 2015 with available price and volume data, both for ADR and domestic securities, in Datastream and accounting data in Worldscope. Volume is the median trading volume (in USD million) during the quarter. Amihud illiquidity is natural logarithm of the median of Amihud's (2002) illiquidity measure during the quarter. Zero returns is the proportion of zero-return trading days during the quarter times. Total market capitalization equals the stock price times the number of shares outstanding (in USD million) measured at the end of the quarter. Return volatility is the standard deviation of daily stock returns over a quarter. All correlation coefficients reported in Panel B are significant at the 1% level.

Table 3. Regression Analysis of the Effect of IFRS Adoption on ADR and Home Market Liquidity

Independent variable	(1) ADR	(2) Home	(1) - (2)	(4) ADR	(5) Home	(4) - (5)	(7) ADR	(8) Home	(7) - (8)
ln[Market capitalization]	0.436	0.325	0.111	0.438	0.325	0.113	0.442	0.327	0.115
	(18.58)***	(19.94)***	(3.47)***	(17.13)***	(20.37)***	(3.41)***	(18.05)***	(19.73)***	(3.55)***
ln[1 + Return volatility]	0.233	-0.016	0.250	0.235	-0.015	0.249	0.219	-0.018	0.237
	(4.08)***	(-0.38)	(3.70)***	(4.14)***	(-0.34)	(3.67)***	(4.23)***	(-0.43)	(3.86)***
Effect of mandatory IFRS adoption									
Post mandatory IFRS x Mandatory adopter	0.140	0.054	0.086						
	(2.49)**	(1.14)	(1.29)						
Post mandatory IFRS x Mandatory adopter x EU				0.371	0.110	0.261	0.398	0.125	0.272
				(4.76)***	(1.11)	(1.80)*	(5.03)***	(1.29)	(1.82)*
Post mandatory IFRS x Mandatory adopter x non-EU				-0.028	0.014	-0.042			
				(-0.38)	(0.23)	(-0.43)			
Post mandatory IFRS x Mandatory adopter x non-EU developed							0.441	0.104	0.337
							(3.69)***	(1.34)	(1.94)*
Post mandatory IFRS x Mandatory adopter x non-EU emerging							-0.119	0.013	-0.131
							(-1.79)*	(0.18)	(-1.11)

(continued on next page)

Independent variable	(1) ADR	(2) Home	(1) - (2)	(4) ADR	(5) Home	(4) - (5)	(7) ADR	(8) Home	(7) - (8)
Post mandatory IFRS x Voluntary adopter	0.139	0.007	0.132						
	(1.49)	(0.08)	(1.03)						
Post mandatory IFRS x Voluntary adopter x EU				0.262	-0.004	0.266	0.289	0.014	0.276
				(2.56)**	(-0.02)	(1.24)	(2.89)***	(0.10)	(1.29)
Post mandatory IFRS x Voluntary adopter x non-EU				-0.062	0.024	-0.086			
				(-0.53)	(0.44)	(-0.84)			
Post mandatory IFRS x Voluntary adopter x non-EU developed							-0.227	-0.159	-0.068
							(-0.49)	(-4.04)***	(-0.15)
Post mandatory IFRS x Voluntary adopter x non-EU emerging							0.001	0.056	-0.055
							(0.01)	(1.10)	(-0.48)

(continued on next page)

Independent variable	(1) ADR	(2) Home	(1) - (2)	(4) ADR	(5) Home	(4) - (5)	(7) ADR	(8) Home	(7) - (8)
Effect of voluntary International GAAP adoption									
Post voluntary International GAAP x EU				0.325	0.085	0.240	0.320	0.084	0.236
				(2.74)**	(0.95)	(1.93)*	(2.75)***	(0.94)	(1.93)*
Post voluntary International GAAP x non-EU				0.177	0.119	0.057			
				(1.52)	(1.46)	(0.45)			
Post voluntary International GAAP x non-EU developed							0.405	0.336	0.069
							(6.50)***	(4.11)***	(0.59)
Post voluntary International GAAP x non-EU emerging							0.053	0.050	0.003
							(0.42)	(0.61)	(0.02)
Adj. R^2	70.1%	82.4%		70.9%	82.5%		71.4%	82.6%	

This table reports the OLS regressions of the effect of IFRS adoption in a firm's domestic market on domestic and ADR liquidity. The sample consists of 12,143 firm-quarters of 239 issuers of level 2 or 3 ADR and domestic securities during the years 1998 and 2015. The dependent variable in models 1, 4 and 7 is the *Liquidity factor* in the ADR market, which is equal to the one factor with an eigenvalue greater than 1.0 extracted from three liquidity variables reported in Table 2 (i.e., *Volume*, -1 times *Amihud illiquidity*, and -1 times *Zero returns*). The dependent variable in models 2, 5 and 8 is the *Liquidity factor* in the domestic market, constructed in the same fashion. Columns 3, 6 and 9 report differences between the coefficients of the ADR and domestic samples. *Market capitalization* is calculated as the stock price times the number of shares outstanding (in USD million) at the end of the quarter. *Return volatility* is the standard deviation of daily stock returns over a quarter. *Post mandatory IFRS* is an indicator variable that takes on the value of one for firm-quarters with IFRS reporting beginning in the calendar quarter following the first fiscal-year end after IFRS became mandatory (Christensen et al. 2013), and zero otherwise. *Mandatory adopter* is an indicator variable that takes on the value of one for firm-quarters of firms that report under local GAAP until the mandatory adoption of IFRS in their domestic market, and zero otherwise. *EU* (*non-EU*) is an indicator variable that takes on the value of one for firm-quarters of firms that are (not) from a European Union country, and zero otherwise. *non-EU developed* (*non-EU emerging*) is an indicator variable that takes on the value of one for firm-quarters of firms that are not from a European Union member country and from a developed (emerging) market, and zero otherwise. *Voluntary adopter* is an indicator variable that takes on the value of one for firm-quarters of firms which report according to "International GAAP" (either U.S. GAAP or IFRS) prior to the mandatory adoption of IFRS in their domestic market, and zero otherwise. *Post voluntary International GAAP* is an indicator variable that takes on the value of one for firm-quarters of firms reporting according to "International GAAP" prior to the mandatory adoption of IFRS in their domestic market, and zero otherwise. All regressions include country, industry (Campbell 1996) and quarter-year fixed effects, with *t*-statistics (in parentheses) based on robust standard errors that are clustered by country and calendar quarter (Christensen et al. 2013). The symbols ***, **, and * denote statistical significance at $p<0.01$, $p<0.05$, and $p<0.10$ levels respectively.

Table 4: Regression Analysis of the Influence of Institutional Characteristics on the Effect of IFRS Adoption on ADR and Home Market Liquidity

	ADR liquidity versus domestic liquidity		
Independent variable	(1) IC = Rule of Law (percentile)	(2) IC = Regulatory Quality (percentile)	(3) IC = ln[1 + Distance from IFRS]
ln[Market capitalization]	0.115	0.112	0.112
	(3.59)***	(3.50)***	(3.48)***
ln[1 + Return volatility]	0.246	0.241	0.253
	(3.75)***	(3.58)***	(3.80)***
Voluntary adopter	-0.539	-0.423	-0.513
	(-1.12)	(-1.77)*	(-1.16)
Institutional characteristic	-0.860	-0.914	0.000
	(-1.87)*	(-1.72)*	(0.00)
Voluntary adopter x Institutional characteristic	0.766	0.550	0.223
	(1.15)	(1.67)	(1.15)
Post mandatory IFRS x Mandatory adopter	-0.271	-0.288	0.079
	(-0.96)	(-0.90)	(0.31)
Post mandatory IFRS x Voluntary adopter	-0.257	-0.350	-0.270
	(-1.12)	(-0.72)	(-1.63)
Post voluntary International GAAP	-0.233	-0.332	0.419
	(-0.91)	(-1.58)	(2.20)**
Post mandatory IFRS x Mandatory adopter x Institutional characteristic	0.533	0.538	0.000
	(1.22)	(1.17)	(0.00)
Post mandatory IFRS x Voluntary adopter x Institutional characteristic	0.565	0.581	0.206
	(1.31)	(0.78)	(2.36)**
Post voluntary International GAAP x Institutional characteristic	0.454	0.683	-0.155
	(1.33)	(2.45)**	(-1.92)*

This table reports differences between the coefficients of OLS regressions of the influence of institutional characteristics on the effect of IFRS adoption on liquidity in two different samples: a sample of ADR securities and a sample of (matched) domestic securities. The sample consists of 12,143 firm-quarters of 239 issuers of level 2 or 3 ADR and domestic securities during the years 1998 and 2015. The dependent variable in all regressions is *Liquidity factor* (equal to the one factor with an eigenvalue greater than 1.0 extracted from three liquidity variables reported in Table 2 – *Volume*, -1 times *Amihud illiquidity*, and -1 times *Zero returns*). All columns report differences between the coefficients of the ADR and domestic samples. *Institutional characteristic* is the percentile score of the rule of law index (taken from the Worldbank's Worldwide Governance Indicators) in column (1), the percentile score of the regulatory quality index (Worldbank's Worldwide Governance Indicators) in column (2), and the magnitude of accounting differences between IFRS and the country's pre-existing GAAP (Bae et al., 2008) in column (3). The remaining variables are as defined in Table 3. All regressions include country, industry (Campbell 1996) and quarter-year fixed effects, with *t*-statistics (in parentheses) based on robust standard errors that are clustered by country and calendar quarter (Christensen et al. 2013). The symbols ***, **, and * denote statistical significance at $p<0.01$, $p<0.05$, and $p<0.10$ levels respectively.

Table 5: Probit Regression Analysis of the Relationship between ADR Characteristics and Region of Origin

Independent variable	(1) EU vs. non-EU	(2) Developed vs. emerging
Intercept	0.703	0.879
	(2.11)**	(2.51)**
Pre-IFRS avg. of ln(Market capitalization)	0.190	0.238
	(1.04)	(1.59)
Pre-IFRS avg.of ln(1 + Return volatility)	0.857	0.624
	(1.88)*	(1.59)
Pre-IFRS avg.of home market liquidity factor	1.280	0.917
	(3.61)***	(3.29)***
Pre-IFRS avg.of foreign sales-to-sales	3.171	2.052
	(5.73)***	(4.03)***
Pre-IFRS avg.of RSQUS	-20.446	-21.023
	(-5.90)***	(-7.34)***
Pre-IFRS hightech indicator	-0.309	-0.286
	(-1.11)	(-1.05)
Pre-IFRS avg.of U.S. industry importance	2.389	2.356
	(1.70)*	(1.84)*
N	239	239
Pseudo-R^2	47.7%	48.9%

This table presents results of probit regressions testing the relationship between ADR issuer characteristics, prior to the adoption of IFRS, and region of origin. The sample consists of 12,143 firm-quarters of 239 issuers of level 2 or 3 ADR and domestic securities during the years 1998 and 2015. The dependent variable in model (1) is *EU*, an indicator variable equal to one for firms that are from a European Union country, and zero otherwise. The dependent variable in model (2) is *Developed*, an indicator variable that takes on the value of one for firms that are from a developed country, and zero otherwise. All independent variables are the pre-IFRS averages of quarterly values. *Market capitalization* is stock price times the number of shares outstanding (in USD million) during the quarter. *Return volatility* is the standard deviation of daily stock returns over a quarter. *Liquidity factor* equals the factor with an eigenvalue greater than 1.0 extracted from three liquidity variables reported in Table 2 (i.e., *Volume*, -1 times *Amihud illiquidity*, and -1 times *Zero returns*). *Foreign sales-to-sales* is the ratio of sales derived outside the firm's domestic market to total sales. *RSQUS* is the adjusted R-square of a regression of domestic daily stock returns on the daily returns of the value-weighted U.S. market index, estimated by firm-quarter. *Hightech indicator* is an indicator variable that takes on the value of one for firms in high-tech sectors (as defined by Pagano et al. 2002), and zero otherwise. *U.S. industry importance* is the fraction of an ICB supersector's global market capitalization that comes from U.S. firms. The symbols ***, **, and * denote statistical significance at $p<0.01$, $p<0.05$, and $p<0.10$ levels respectively.

Table 6: Regression Analysis of the Effect of IFRS Adoption on ADR and Home Market Liquidity while Controlling for ADR Differences across Regions of Origin

Independent variable	(1) ADR	(2) Home	(1) - (2)
ln[Market capitalization]	0.436	0.328	0.108
	(17.07)***	(20.96)***	(3.56)***
ln[1 + Return volatility]	0.223	-0.017	0.240
	(4.53)***	(-0.41)	(4.12)***
ADR characteristics control variables			
Post mandatory IFRS x Mandatory adopter x Developed propensity	-0.131	0.202	-0.333
	(-0.93)	(1.50)	(-2.03)*
Post mandatory IFRS x Voluntary adopter x Developed propensity	0.098	0.285	-0.187
	(0.29)	(1.73)*	(-0.49)
Post voluntary International GAAP x Developed propensity	0.115	-0.408	0.523
	(0.40)	(-2.51)**	(1.76)*

(continued on next page)

Independent variable	(1) ADR	(2) Home	(1) - (2)
Effect of mandatory IFRS adoption			
Post mandatory IFRS x Mandatory adopter x EU	0.493	-0.037	0.530
	(3.92)***	(-0.32)	(4.06)***
Post mandatory IFRS x Voluntary adopter x EU	0.179	-0.190	0.369
	(0.70)	(-1.09)	(1.25)
Post mandatory IFRS x Mandatory adopter x non-EU developed	0.519	-0.030	0.549
	(4.06)***	(-0.29)	(4.34)***
Post mandatory IFRS x Voluntary adopter x non-EU developed	-0.338	-0.337	-0.001
	(-0.67)	(-3.30)***	(-0.00)
Post mandatory IFRS x Mandatory adopter x non-EU emerging	-0.098	-0.028	-0.070
	(-1.36)	(-0.33)	(-0.50)
Post mandatory IFRS x Voluntary adopter x non-EU emerging	0.007	-0.093	0.099
	(0.03)	(-1.29)	(0.51)
Effect of voluntary International GAAP adoption			
Post voluntary International GAAP x EU	0.255	0.394	-0.139
	(1.01)	(2.08)**	(-0.51)
Post voluntary International GAAP x non-EU developed	0.361	0.606	-0.245
	(1.77)*	(3.56)***	(-0.89)
Post voluntary International GAAP x non-EU emerging	-0.065	0.261	-0.325
	(-0.39)	(2.16)**	(-1.85)*
Adj. R^2	72.1%	82.7%	

This table reports the OLS regressions of the effect of IFRS adoption in a firm's domestic market on domestic and ADR liquidity while controlling for ADR differences across regions of origin. The sample consists of 12,143 firm-quarters of 239 issuers of level 2 or 3 ADR and domestic securities during the years 1998 and 2015. The dependent variable in model (1) is the *Liquidity factor* in the ADR market, which is equal to the one factor with an eigenvalue greater than 1.0 extracted from three liquidity variables reported in Table 2 (i.e., *Volume*, -1 times *Amihud illiquidity*, and -1 times *Zero returns*). The dependent variable in model 2 is the *Liquidity factor* in the domestic market, constructed in the same fashion. Column 3 reports differences between the coefficients of the ADR and domestic samples. *Developed propensity* is the predicted propensity of a firm being from a developed country, estimated using Model 2 in Table 5. The remaining variables are as defined in Table 3. All regressions include country, industry (Campbell 1996) and quarter-year fixed effects, with *t*-statistics (in parentheses) based on robust standard errors that are clustered by country and calendar quarter (Christensen et al. 2013). The symbols ***, **, and * denote statistical significance at $p<0.01$, $p<0.05$, and $p<0.10$ levels respectively.

Table 7: Regression Analysis of the Effect of IFRS Adoption on level I ADR Liquidity

	ADR liquidity versus domestic liquidity			
Independent variable	Table 3 (9) replication	Table 4 (1) replication - IC = Rule of Law (percentile)	Table 4 (2) replication - IC = Reg. Quality (percentile)	Table 4 (3) replication - IC = ln[1 + Distance from IFRS]
ln[Market capitalization]	-0.100	-0.101	-0.100	-0.102
	(-2.39)**	(-2.50)**	(-2.49)**	(-2.55)**
ln[1 + Return volatility]	-0.029	-0.028	-0.023	-0.018
	(-0.30)	(-0.30)	(-0.25)	(-0.19)
Effect of voluntary International GAAP adoption				
Post voluntary International GAAP x EU	-0.269			
	(-0.98)			
Post voluntary International GAAP x non-EU developed	0.199			
	(2.31)**			
Post voluntary International GAAP x non-EU emerging	0.421			
	(4.22)***			
Post voluntary International GAAP x Inst. characteristic		-1.005	-0.934	-0.241
		(-2.05)**	(-0.96)	(-2.94)***

(continued on next page)

Independent variable	ADR liquidity versus domestic liquidity			
	Table 3 (9) replication	Table 4 (1) replication - IC = Rule of Law (percentile)	Table 4 (2) replication - IC = Reg. Quality (percentile)	Table 4 (3) replication - IC = ln[1 + Distance from IFRS]
Effect of mandatory IFRS adoption				
Post mandatory IFRS x Mandatory adopter x EU	0.497 (6.13)***			
Post mandatory IFRS x Mandatory adopter x non-EU developed	0.312 (2.45)**			
Post mandatory IFRS x Mandatory adopter x non-EU emerging	-0.140 (-0.89)			
Post mandatory IFRS x Mandatory adopter x Inst. characteristic		1.157 (3.78)***	1.218 (2.35)**	-0.113 (-0.96)
Post mandatory IFRS x Voluntary adopter x EU	1.093 (5.90)***			
Post mandatory IFRS x Voluntary adopter x non-EU developed	-0.002 (-0.02)			
Post mandatory IFRS x Voluntary adopter x non-EU emerging	-0.267 (-1.35)			
Post mandatory IFRS x Voluntary adopter x Inst. characteristic		3.043 (3.21)***	4.788 (3.47)***	0.355 (2.39)**

This table reports differences between the coefficients of OLS regressions of the influence of institutional characteristics on the effect of IFRS adoption on liquidity in two different samples: a sample of ADR securities and a sample of (matched) domestic securities. The sample consists of 18,925 firm-quarters of 239 issuers of level 1 ADR and domestic securities during the years 1998 and 2015. The dependent variable in all regressions is *Liquidity factor* (equal to the one factor with an eigenvalue greater than 1.0 extracted from three liquidity variables reported in Table 2 – *Volume*, -1 times *Amihud illiquidity*, and -1 times *Zero returns*). All columns report differences between the coefficients of the ADR and domestic samples. *Institutional characteristic* is the percentile score of the rule of law index (taken from the Worldbank's Worldwide Governance Indicators) in column (2), the percentile score of the regulatory quality index (Worldbank's Worldwide Governance Indicators) in column (3), and the magnitude of accounting differences between IFRS and the country's pre-existing GAAP (Bae et al., 2008) in column (4). The remaining variables are as defined in Table 3. All regressions include country, industry (Campbell 1996) and quarter-year fixed effects, with *t*-statistics (in parentheses) based on robust standard errors that are clustered by country and calendar quarter (Christensen et al. 2013). The symbols ***, **, and * denote statistical significance at $p<0.01$, $p<0.05$, and $p<0.10$ levels respectively.

Chapter 4

The Disciplinary Role of Accounting Conservatism: Evidence from State Antitakeover Laws[32]

4.1 Introduction

Accounting conservatism is defined as the asymmetric verification threshold for gains versus losses, where the threshold is higher for losses (Basu, 1997; Khan and Watts, 2009), and is one of the oldest and most influential principles in accounting (Sterling, 1970; Watts, 2003). While the benefits of conservatism in debt contracting have received widespread empirical support (*e.g.*, Ahmed et al., 2002; Wittenberg-Moerman, 2008; Zhang, 2008; Aier et al., 2014), its effect on investment decisions and especially its value to equity holders are far less clear[33]. As such, I analyze the effects of conservatism on managerial investment decisions, focussing on corporate acquisitions. More specifically, I exploit the staggered and

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[32] This chapter is based on Ghazizadeh, P. 2018. The disciplinary role of accounting conservatism: Evidence from state antitakeover laws. *Working paper*.

[33] More specifically, Francis and Martin (2010) and García Lara *et al.* (2016) provide evidence that conservatism is associated with improved investment efficiency, whereas Kravet (2014) shows that conservatism is associated with less risk taking. While Kravet (2014) is silent on the efficiency of the investment decisions, a large literature is concerned with the adverse effects of managerial risk aversion for *equity* holders (*e.g.*, Coles *et al.*, 2006).

unanticipated passage of state antitakeover laws (henceforth SAL) as an exogenous shock to managerial investment discretion (*e.g.*, Bertrand and Mullainathan, 2003; Gormley and Matsa, 2016). I further argue that the pre-SAL level of conservatism of firms incorporated in a state is unrelated to the passage of SAL, as firms have built up their level of conservatism over years prior to these *unexpected* changes. As the passage of SAL does not differentially affect the investment opportunity set for firms conditional on their level of conservatism this setting allows me to circumvent endogeneity concerns and test whether accounting conservatism disciplines managers when there is a marked increase in their investment discretion.

Extant theoretical work does not provide unambiguous predictions as to whether conservatism disciplines managers' investment decisions and thereby affects equity values. On the one hand, conservatism is considered an efficient contracting mechanism (Watts, 2003), as the more timely recognition of losses expedites the signal investors receive regarding managers' prior inefficient investment decisions, which reduces their willingness to make negative NPV investments (Ball, 2001; Ball and Shivakumar, 2005). As such, conservatism can constrain managers' opportunistic behavior (*i.e.*, overinvestment), especially in settings prone to moral hazard problems. On the other hand, it is also argued that conservatism can have dysfunctional consequences by inducing underinvestment in positive, but risky, NPV projects (Roychowdhury, 2010; Kravet, 2014). More specifically, risky investments expose firms to larger potential losses than less risky investments with *in expectation* the same NPV. As large losses are disproportionately more costly than smaller losses (*e.g.*, they are much more likely to trigger debt covenants), even risk neutral managers would rationally avoid the riskier investments when under conservative accounting the large loss is recognized timely, but the reporting of a corresponding large gain from the risky investment is deferred (Kravet, 2014). Thus, for conservatism to truly act as a disciplinary mechanism, it should not only reduce overinvestment (*i.e.*, negative NPV investments), but it should also not impede efficient risk taking (*i.e.*, risky and positive NPV investments).

To ascertain whether conservatism disciplines managerial behavior, I analyze corporate acquisitions in relation to the passage of state antitakeover laws, as this context provides several advantages. First, corporate acquisitions have been widely recognized as a prime event in which managers can pursue both overinvestment and inefficient risk reduction – which are exactly the manifestations of managerial discretion that conservatism may affect. For instance, Jensen (1986) describes managers' incentives to engage in empire building via acquisitions at the expense of equity holders, while Amihud and Lev (1981) and Gormley and Matsa (2016) show that managers actively pursue the reduction of their firms' riskiness via value-destroying acquisitions[34]. Second, corporate acquisitions are discrete investments with readily identifiable dates, which typically have a large impact on the firm's value. This facilitates 'clean' empirical analysis (Roosenboom et al, 2014), especially in a difference-in-difference setting. Finally, prior work shows that the market for corporate control is one of the institutional arrangements that can discipline managers' acquisition decisions (Mitchell and Lehn, 1990; Berger and Ofek, 1996). State antitakeover laws however reduce the disciplinary threat of acquisitions by increasing the cost of hostile takeovers (*e.g.*, Karpoff and

[34] The findings of Gormley and Matsa (2016) are especially relevant to the present study, as they show that inefficient risk reduction occurs when managers are shielded from the market for corporate control as a results of SAL.

Malatesta, 1989), and thus provide an exogenous shock to managerial discretion. By focusing on the *change* in the acquisition activity of firms before and after the passage of SAL conditional upon pre-SAL levels of conservatism, I avoid most of the endogeneity concerns, which – as the authors acknowledge – make it difficult to infer a causal link between conservatism and investment efficiency in previous studies (*i.e.*, Francis and Martin, 2010; Kravet, 2014; García Lara et al, 2016).

In order to test whether conservatism reporting disciplines managers' acquisition decisions, I exploit the staggered passage of SAL as an exogenous shock to managerial discretion. I construct a measure for accounting conservatism similar to the *c_score* (Khan and Watts, 2009), and use its pre-SAL level to classify the firm's conservatism. I then examine the relation between conservatism and acquisition activity for publicly listed US firms incorporated in states that pass SAL with those incorporated elsewhere. More specifically, using a difference-in-difference setup, I test whether the *ex-ante* classification of firms by their level of conservatism explains the *ex post* change in acquisition activity, firm riskiness, and performance. These tests are similar to Balakrishnan et al (2016), who also use an *ex-ante* measure of conservatism based on conservatism levels prior to a shock (*i.e.*, the 2007-2008 global financial crisis), which is then interacted with an indicator variable designating post-shock observations, in a specification including firm fixed effects. Using state-year and industry-year fixed effects, I further control for unobserved, time-varying differences across industries as well as unobserved, state-level economic conditions that may coincide with the passage of SAL (Gormley and Matsa, 2016). Thus, in essence, I extend the findings of Gormley and Matsa (2016) by testing whether the sensitivity of acquisition activity to an exogenous shock to managerial discretion is moderated by the firm's level of conservative accounting.

I start my investigation by testing whether stock market participants react differently to the news of the passage of SAL conditional on the *ex-ante* measure of conservatism of affected firms. More specifically, I conduct event studies on three dates when the first news reports related to the passage of SAL in Delaware – the state in which approximately 60% of the sample firms are incorporated – were disseminated. I then regress the abnormal returns on the *ex-ante* conservative measure, and find that equity holders react less negatively to an increase in managerial discretion for conservative firms. This result is in line with the hypothesis that equity holders *anticipate* that the increased managerial discretion will adversely affect the *value* of their stocks, and – more importantly for this study – that they expect conservative accounting to mitigate this.

I then turn to the examination of the changes in corporate acquisitions by firms. I find strong evidence that the *ex-ante* level of conservatism is negatively related to the change in the acquisitions activity. That is, whereas firms that did not report conservatively prior to the adoption of the SAL increase their acquisition activity, there is no change in the behavior of conservatively reporting firms. The results also reveal that impact of SAL and conservatism on acquisition activity is highly economically significant: the estimated increase in acquisition activity by 1.04% is almost half of the overall sample average of annual acquisitions, but it is fully offset by a two standard deviation increase (0.40) in the level of conservatism. As the adoption of SAL did not affect firm's investment opportunities, these results indicate that the conservatism strongly affects acquisition activity.

The main results are robust to alternative specifications and subsamples. For instance, I obtain similar results when I include different time fixed effects (*i.e.*, state-year and industry year fixed effects or year fixed effects), or different sets of control variables. The results are also robust to the exclusion of firms incorporated in Delaware, or the restriction of the sample period to only the 11 years surrounding the adoption of the SAL. I also find that the observed relationship is stronger for firms incorporated in states where the strength of the SAL was high (Armstrong et al, 2012), providing further assurance regarding the causal effect of the conservatism on managerial discretion. The findings also survive tests designed to parse out any concerns regarding changes in acquisition activity due to anticipation of the SAL. Finally, the same results are obtained when the tests are run separately for each year in which SAL were adopted, which allows the inclusion of firms from states in which no SAL were ever adopted, rendering the results robust to different benchmark firms.

A concern voiced in prior literature regarding the effects of conservatism is that they may be spuriously driven by management ability, when better managers would also report more conservatively. I address this issue by controlling for two different proxies for managerial ability; the effect of conservatism remains. More specifically, using pre-SAL industry adjusted profitability as a proxy for managerial ability, I find that both managerial ability and conservatism are negatively related to post SAL acquisition activity, indicating that even taking into account the lower proclivity of good managers to overinvest, conservatism still continues to exhibit a disciplinary effect. Using an alternative specification, we find that the mitigating effect of conservatism on acquisition activity is stronger for firms with lower ability. I find very similar results when using the *MA-Score* (Demerjian et al, 2012) as a proxy for managerial ability. Thus, while I concur with Roychowdhury's (2010) view that an interpretation of better managers embracing conservatism is valuable, my results seem to suggest that conservatism has an independent disciplinary effect on acquisition investments[35].

I next test the differential changes in firm performance and riskiness after the adoption of SAL conditional on firms' pre-SAL level of conservative reporting, as the differential change in acquisition activity cannot unambiguously be interpreted as inefficient. More specifically, whereas an increase in acquisition activity would be *inefficient* if it is the results of the pursuit of managerial self-interest (*i.e.*, empire-building or risk reduction), an increase would be *efficient* when the insulation from takeovers provided by state antitakeover laws allows managers to pursue more risky, yet (in the long run) more profitable acquisitions they previously would have forgone. As the proposed potential dysfunctional effects of conservatism on efficient risk taking would *negatively* affect an increase in acquisition activity of the latter kind (*i.e.*, efficient), whereas the proposed disciplinary effects of conservatism would *negatively* affect an increase in acquisition activity of the former kind (*i.e.*, inefficient), testing the effect of conservatism on firm performance and riskiness after increased acquisition activity allows me to measure the *net effect* of conservatism on investment efficiency. The results of my analyses strongly indicate that conservatism disciplines managers (more than it provides them disincentives to pursue risky projects); while firm performance and riskiness decreases after the adoption of SAL, the level of pre-SAL conservatism is

[35] I still urge the reader to remain cautious in interpreting these results as conclusive, as managerial ability (and more generally governance quality) is notoriously hard to measure.

positively related to both' Together with the differential acquisition activity, this finding indicates that conservative reporting indeed curbs managerial overinvestment, while being highly inconsistent with the notion that conservatism prevents efficient risk taking.

The contribution of this study is twofold. First, there is no consensus in the literature as to the impact of conservatism on investment efficiency. More specifically, concerns have been voiced regarding potentially adverse effects of conservatism on investment decisions (*e.g.*, Penman and Zhang, 2002; Roychowdhury, 2010; Kravet, 2014). The results of this chapter however point in the opposite direction, as they strongly suggest that conservatism improves managerial (acquisition) investment decisions, and that equity holder value this. Furthermore, as argued by Roychowdhury (2010), the implications of conservative accounting for managerial decision making and firm value should be taken into account, when one considers the recent move towards fair-value based standards supported by influential bodies such as the FASB.

Second, the evidence that suggests conservatism improves firm value is either based on debt market advantages, or wanting when it relates to equity holders. Speaking to value implications for equity holder, the literature has mostly focussed on managerial investment decisions (*i.e.*, overinvestment and risk avoidance). However, the evidence provided by these studies (*i.e.*, Francis and Martin, 2010; García Lara et al., 2016; Kravet, 2014) is insufficient to infer a causal effect of conservatism (Roychowdhury, 2010; Francis and Martin, 2010; García Lara et al., 2016; Kravet, 2014). The setting used in this chapter allows me to avoid the endogeneity concerns of prior studies, and especially the event study results – to the best of my knowledge the first in evaluating equity holders' reaction conditional on the level of conservatism[36] – probably provide the most salient and direct evidence that equity holders value conservatism. The contribution of the results of this study to this literature are perhaps best manifested as follows; while my findings lend support to those of Francis and Martin (2010) and García Lara et al (2016), they cast doubt on those of Kravet's (2014) seminal[37] work. The reader should however be made aware of the concurrent and the closely related study by Cedergren et al (2015), which arrives at exactly the opposite conclusion[38].

The remainder of this study is organized as follows. The next section reviews the extant literature and develops the hypotheses. In section 3 the research design, the selection of the sample, and variable construction are discussed. Section 4 reports and discusses the results, and section 5 concludes.

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[36] Khan and Watts (2009) also conduct several event studies, but they investigate changes to conservatism in response to several events, rather than gauge the reaction of equity holders to those events.

[37] While Roychowdhury (2010) initiates the discussion of the potential dysfunctional effects of conservatism on managerial investment decisions, Kravet (2014) further develops this hypothesis and provides the first empirical evidence.

[38] Another concurrent and closely related study that merits further discussion is Balakrishnan et al (2016), which investigates the mitigating effect of conservatism on *under*investment due to the amelioration of financial market frictions (and not managerial risk incentives, as is part of the focus of this chapter). Put differently, whereas my results indicate that conservatism mitigates managerial misuse of investment discretion (*i.e.*, *over*investment), their results suggest that conservatism facilitates external financing that allows managers to pursue value adding projects.

4.2 Background and Hypothesis Development

4.2.1 Managerial Discretion and Investment Efficiency

Managers enjoy a large degree of discretion over the investment policy of their firms, as shareholders have delegated the decision-making authority to them (Jensen and Meckling, 1976). The superior information managers possess further impedes the ability of shareholders to evaluate their actions, *de facto* further increasing managers' discretion. Given the potential divergence of incentives, it has been widely recognized that managers are inclined to use corporate resources to pursue their own self-interests, (e.g., Berle and Means, 1923; Jensen and Meckling, 1976). Two manifestations of such moral hazard problems have been the focus of much of the prior literature: empire building (e.g., Baumol, 1959; Marris, 1964; Jensen, 1986) and inefficient risk reduction (*e.g.*, Amihud and Lev, 1981; Morck, Shleifer and Vishny, 1990, Gormley and Matsa, 2016)[39]. The former refers to the managerial departure from an optimal level of investment by investing in negative NPV projects, where managers only bare part of the cost, while enjoying most of its (private) benefits. The inefficient risk reduction builds on the premise that managers' wealth is not properly diversified, as the risk associated with managers' income is closely related to the firm's risk (Amihud and Lev, 1981). By diversifying the operations of their firms, they are able to reduce the risk to their human capital (Morck et al., 1990), even if this comes at the expense of shareholders.

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[39] Besides these moral hazard problems, inefficient investment could also arise due to *adverse selection* problems. For instance, firms could either overinvest when they have successfully sold overprices securities (*e.g.*, Baker et al, 2003), or underinvest due to credit rationing when suppliers of capital recognize the firms' incentive to sell overpriced securities (*e.g.*, Myers and Majluf, 1984). Similar credit rationing arguments can ensue in *ex post* underinvestment due to *ex-ante* recognition of moral hazard problems (*e.g.,* Stiglitz and Weiss, 1981).

Corporate acquisitions have been widely recognized as a prime vehicle through which managers can pursue both overinvestment and inefficient risk reduction[40]. By increasing the resources under their span of control through acquisitions, managers increase the non- pecuniary benefits (*e.g.*, prestige) they enjoy (Stulz, 1990). Managers could further be motivated by pecuniary benefits (*i.e.*, higher compensation) that arise from increases in firm size through acquisitions (Lambert, Larcker and Weigelt, 1991)[41]. Corporate acquisitions, especially of assets in (unrelated) businesses with imperfectly correlated cash flows, can also lead to reductions in the firm's riskiness – *i.e.,* the coinsurance effect (Lewellen, 1971). Amihud and Lev (1981) show that such unrelated acquisitions are prompted by manager's desire to reduce their undiversifiable employment risk. In line with this, Gormley and Matsa (2011) show that managers react to increased liability risk by acquiring large, unrelated businesses with relatively high operating cash, especially when they have a higher personal exposure to their firms' risk.

[40] Note that corporate acquisitions, in addition to allowing managers to reap theses private benefits, are also characterized by (1) high degrees of information asymmetry, and (2) are notoriously hard to evaluate, even absent any information frictions (*e.g.*, estimating the counterfactual). As such, corporate acquisitions are highly prone to moral hazard problems.

[41] Note that the positive relation Lambert et al. (1991) show between managerial compensation and firm size is stronger for levels than for changes (*e.g.*, through acquisitions). Moreover, Lambert and Larcker (1987) show that changes to managerial compensation and wealth only increase after acquisitions which increase shareholder wealth. Avery, Chevalier and Schaefer (1998) show that while the compensation of managers does not increase after completing acquisitions, the likelihood of gaining outside directorships does increase, further attesting to the increase in prestige post acquisitions.

4.2.2 Market for Corporate Control and State Antitakeover Laws

One possible disciplinary mechanism which deals with such managerial misconduct is the market for corporate control, where firms headed by inefficient managers are – through being acquired – replaced by those managers that apply their assets more efficiently (Manne, 1965; Marris, 1964; Jensen, 1986). Put differently, managers compete for the privilege to manage the firm's resources (Jensen and Ruback, 1983). Mitchell and Lehn (1990) provide empirical support for the disciplinary role of the market for corporate control and show that firms that make poor acquisitions, themselves become targets, and the bad acquisitions are undone subsequent to being bought. Similarly, Berger and Ofek (1996) show that firms with higher levels of diversification discount are more likely to be targeted and broken up.

Given the evidence on the disciplinary role of takeover threats, any restriction to the takeover market arguably increases managers' investment discretion. State antitakeover laws are such a restriction. A number of states passed a number of antitakeover laws in the mid- to late 1980s (the so-called 'second generation') and late 1980s and early 1990s (so-called 'third-generation') (Armstrong et al., 2012). These antitakeover laws effectively increase the cost of a takeover, and subsequently reduce the likelihood of a takeover[42]. Empirical evidence supports this notion; Comment and Schwert (1995) show that takeover premiums increased after the adoption of state antitakeover laws, and Schwert (2000) shows a decrease in the incidence of

[42] *E.g.*, business combination laws impose a moratorium on transactions – including mergers – between the acquirer and targeted firm for a period of three to five years, once the stake of the acquirer has reached a certain threshold (John, Li and Pang, 2017). Several studies provide more elaborate descriptions of state antitakeover laws (*e.g.*, Bertrand and Mullainathan, 2003; Armstrong et al., 2012; Gormley and Matsa, 2016).

takeovers. Prior studies further show that insulating managers from takeovers results in overinvestment and risk reduction[43]. For instance, Cheng, Nagar and Rajan (2004) find that managers significantly reduced their (risky) stockholdings, thus reducing their wealth exposure to firm performance, while retaining their prior level of control. Garvey and Hanka (1999) show that firms in states that adopted antitakeover laws reduced their leverage. Gormley and Matsa (2016) show that managers pursue value-destroying acquisitions to reduce their firms' risk of distress when they are insulated from disciplinary takeovers.

Despite the abovementioned evidence that suggests managers abuse the protection provided by alleviating the threat of hostile takeovers, it is nevertheless possible that state antitakeover laws have a positive effect on investment decisions of managers – without affecting the investment opportunity set. In particular, it is argued that when takeover threat is high, managers may be less inclined to tie their human capital to the firm which would otherwise foster the creation of innovative products (Shleifer and Summers, 1988). In line with this argument, Chemmanur and Tian (2018) show that protection against hostile takeovers spurs innovation[44].

4.2.3 Conservatism and Investment Efficiency

As managers have different preferences in allocating the firm's resources, it is crucial to the firm's governance structures that the consequences (*i.e.*, output) of managerial decisions are observable. One of the

43 Bertrand and Mullainathan (1999, 2003) also provides evidence of other increases in agency costs, such as increased managerial compensation and reduced productivity.
44 See Atanassov (2013) for evidence of the opposite.

key roles of accounting is to generate information regarding the firm's operations, such that it is used to monitor managers. However, managers themselves are responsible for, or at least can affect, the preparation of financial reports. The rules governing the preparation of financial reports (*i.e.*, GAAP) have therefor evolved to contain certain attributes to safeguard their usefulness, despite the managers' proclivity to favorably skew reported performance (Kothari, Rammana and Skinner, 2010). Conservatism, a lower verifiability threshold for adverse information, is such an attribute. Kothari et al. (2010) vividly illustrate this; when the benefits that could arise from expenditures are sufficiently uncertain, these costs are expensed (in violation of the matching principle), as managers with limited tenures have an incentive to (indefinitely) postpone their recognition as expenses and provide a favorably biased picture of the firm's performance. As conservatism impedes the ability of managers to misrepresent financial reports (*i.e.*, improves the credibility and arguably the accuracy of information), the more timely recognition of losses also expedites the signal monitors receive regarding managers' prior inefficient business decisions and facilitates timely intervention. In line with these arguments, Roychowdhury (2010, p. 180) states that the primary hypothesized purpose of conservatism is to facilitate accounting's role in firm monitoring and governance by external parties.

In addition to enhancing *ex post* monitoring leading to more efficient *ex-ante* decision making, conservatism can also directly improve investment efficiency by increasing the sensitivity of managerial wealth to inefficient investment. Managers' compensation and the firm's debt covenants are often based on accounting earnings measures (Watts, 2003). Timely recognition of adverse outcomes reduces the earnings-based compensation of managers and accelerates covenant violations, increasing the cost born by the manager by investing inefficiently (Ball and Shivakumar, 2005). Similarly, conservatism can also precipitate the discontinuation of loss-making projects[45] (Francis and Martin, 2010). That is, as managers incur a cost when abandoning a loss-making project in the absence of conservatism (*e.g.*, through write-offs that reduce earnings), they have an incentive not to – especially given their limited tenure. As under conservatism the losses would be recognized early on, there is no additional cost for the manager to discontinue the loss-giving project. Taken together, conservatism can enhance investment efficiency directly by increasing the managerial cost of inefficient investment, and indirectly by facilitating more timely *ex post* monitoring.

Prior studies provide evidence in support of the notion that conservatism enhances investment efficiency. More specifically, Francis and Martin (2010) find that more conservatively reporting firms make more profitable acquisitions and divest more timely. García Lara et al. (2016) provide further evidence by showing an association between conservative reporting and reduced overinvestment[46]. In addition, several studies investigate whether conservatism improves monitoring. For instance, Ahmed and Duellman (2007) find evidence consistent with accounting conservatism assisting directors in reducing agency costs, while LaFond and Watts (2008)

45 Note that while the quality of the project here is low, the *ex-ante* estimation of its profitability is not considered. That is, this could have been – *in expectation* – a positive NPV project, which did not turn out to be profitable.

46 García Lara et al. (2016) also show that conservatism improves investment efficiency by limiting underinvestment that arises from debt-equity conflicts. Balakrishnan et al. (2016), studying the investment behavior of firms during the 2007-2008 financial crisis, report similar findings, further bolstering the potential of conservatism to curb underinvestment in the presence of information frictions.

indicate that there is a higher demand for conservatism in firms with high *ex-ante* agency conflicts (*i.e.*, high information asymmetry).

4.2.4 Conservatism and Underinvestment in Risky Projects

Alternatively, it is also argued that conservatism can have dysfunctional consequences by inducing underinvestment in positive, but risky, NPV projects (Roychowdhury, 2010; Kravet, 2014). More specifically, risky investments expose firms to larger potential losses than less risky investments with in expectation the same NPV. As large losses are disproportionately more costly than smaller losses (*e.g.*, they are much more likely to trigger debt covenants), even risk neutral managers would rationally avoid the riskier investments when under conservatism accounting the large loss is recognized timely, but the reporting of a corresponding large gain from the risky investment is deferred (Kravet, 2014). In line with this hypothesis, Kravet (2014) finds that managers of more conservatively reporting firms make less risk acquisitions.

4.2.5 Hypothesis Development

Based on the above, I argue that the adoption of antitakeover laws in the state in which a manager's firm is headquartered increases acquisition activity. More specifically, the insulation from takeover threats should both allow managers to pursue their own self-interest through acquisitions, as well as allow them to pursue more risky, yet (in the long run) more profitable acquisitions they previously would have forgone. The goal of this study however lies in uncovering the role of conservative reporting in managerial investment decisions. The proposed disciplinary effect of conservatism is expected to mitigate the increase in acquisition activity when this arises from manager's pursuit of self-interest in allocating the firm's resources. The effect of the potential dysfunctional role of conservatism is less clear-cut: if conservatism deters managers from taking efficient risks, it could either *mitigate efficient* investments that are undertaken under the protection of state antitakeover laws, or *exacerbate inefficient* acquisitions aimed at reducing the riskiness of the firm. Overall, I expect the mitigating effects to weigh in more heavily, leading to the following hypothesis:

> *H1: The increase in acquisitions activity following the adoption of state antitakeover laws will be less pronounced for firms that reported more conservatively prior to the adoption of state antitakeover laws.*

As evidence in line with the above hypothesis would still be inadequate in allowing me to make inferences, I exploit the opposite implications of the two proposed effects of conservatism on investment efficiency. More specifically, whereas an increase in acquisition activity would be *inefficient* if it is the results of the pursuit of managerial self-interest (*i.e.*, empire-building or risk reduction), an increase would be *efficient* when the insulation from takeovers provided by state antitakeover laws allows managers to pursue more risky, yet (in the long run) more profitable acquisitions they previously would have forgone. As the proposed potential dysfunctional effects of conservatism on efficient risk taking would *negatively* affect an increase in acquisition activity of the latter kind (*i.e.*, efficient), whereas the proposed disciplinary effects of conservatism would *negatively* affect an increase in acquisition activity of the former kind (*i.e.*, inefficient), testing the effect of conservatism on firm performance and riskiness after increased acquisition activity allows me to measure the *net effect* of conservatism on investment efficiency. To be able to state my hypotheses directionally, I (randomly) assume that conservatism has a disciplinary effect on managerial investment decisions, leading to the following hypotheses:

> *H2: The decrease in firm performance following the adoption of state antitakeover laws will be less pronounced for firm who reported more conservatively prior to the adoption of state antitakeover laws.*

> *H3: The decrease in firm riskiness following the adoption of state antitakeover laws will be less pronounced for firm who reported more conservatively prior to the adoption of state antitakeover laws.*

4.3 Methodology

4.3.1 Research Design

The goal of this study is to analyze the disciplinary effect of accounting conservatism on managerial investment decisions. However, given the potential joint determination of conservatism in reporting and investment decisions[47], and unobservable determinants of conservatism, a static analysis could yield biased results. In order to circumvent these issues, I exploit the staggered and unanticipated passage of state antitakeover laws. The adoption of these laws provides an exogenous shock to managerial discretion, which allows me to better isolate the effect of conservatism on investments arising from increased managerial discretion. That is, I argue that if conservatism

[47] For instance, it is possible that the level of conservatism in the firm's reporting is set in anticipation of future investment decisions.

affects investment decisions by mitigating managers' abuse of their increased discretion, we should observe a less pronounced *change* in acquisition activity, the more conservatively they report. Furthermore, the unanticipated passage of these laws ensures that other factors' effect on conservatism (*e.g.*, anticipated investments) does not affect my results.

More specifically, I argue that firms' level of conservatism in the year preceding the adoption is unlikely to be related to the passage of state antitakeover laws, as these were adopted unexpectedly (Bertrand and Mullainathan, 2003; Armstrong et al, 2012), whereas firms build up their level of conservatism over the course of years[48]. Moreover, I argue that firm-level *changes* in acquisition activity can be ascribed to the increased managerial discretion – *i.e.*, there is no change to firms' investment opportunity set. It then follows that the passage of state antitakeover laws did *not* differentially affect the investment opportunity set for firms conditional on their level of conservatism. I also argue that firms' pre-adoption level of conservatism is a good indicator for future conservatism, or more specifically, that firms that report relatively more conservatively prior to the adoption remain relatively more conservative after the passage of state antitakeover laws[49]. The latter is in line with the reported stickiness of conservatism by the extant literature (*e.g.*, Khan and Watts, 2009; Kim et al., 2013). Taken together, the above allows me to effectively use firms' pre-adoption level of conservatism as an instrument for their cross-sectional ranking after the passage of state antitakeover laws. The use of an *ex-ante* measure of conservatism, *i.e.* measuring conservatism prior to the shock and holding it constant after, then ensures that concerns

[48] As Balakrishnan et al. (2016, p1) state; *"accounting conservatism is a long-run equilibrium response to various institutional factors and firm characteristics"*.
[49] This line of argumentation is adopted from Balakrishnan et al. (2016, p 20).

regarding anticipation of investment or any other factors affecting the level of conservatism, are alleviated.

Using the above *ex-ante* measure of conservatism, I employ a difference-in-difference design to test whether the sensitivity of acquisition activity to increased managerial discretion is mitigated by the level of conservative reporting. More specifically, similar to Gormley and Matsa (2016) I estimate:

$$Y_{ijlst} = \beta_1 SAL_{st} + \beta_2 SAL_{st} * Conservatism_i + a_i + \gamma_{lt} + \lambda_{jt} + + \varepsilon_{ijlst} \qquad (1)$$

where Y is the outcome of interest for firm i, in industry j, located in state l, incorporated in state s, in year t; SAL in an indicator variable that equals 1 if state s has passed a state antitakeover law by year t; a_i are firm fixed effects; γ_{lt} are state-by-year fixed effects; and λ_{jt} are industry-by-year fixed effects. The firm fixed effects control for unobserved, time-invariant differences across firms; state-by-year fixed effects control for unobserved, time-varying difference across states (*e.g.*, local business cycles); and industry-by-year fixed effects control for unobserved, time-varying difference across industries (*e.g.*, differential trends across industries). As Gormley and Matsa (2016, p. 437) state, *the inclusion of state-by-year and industry-by-year fixed effects ensure that the difference-in-difference estimates are robust to many types of unobservable omitted variables that might otherwise confound the analysis*[50]. The coefficients of interest are β_1 and especially β_2: the former captures the differential response of two firms that operate in the same state, but where only one of these firms is incorporated in a state that passes an antitakeover law; the latter captures the moderating effect of conservatism on the differential response. Furthermore note that the

[50] For the additional advantages of this estimation strategy, the interested reader is referred to the original article by Gormley and Matsa (2016, p.437).

pre-adoption level of conservatism itself is subsumed by the firm fixed effects. Finally, the standard errors are clustered at the state-of-incorporation level.

4.3.2 Sample Selection and Variable Construction

I restrict my sample to the period from 1980 to 1996, where the sample period begins (ends) 5 years before (after) the adoption of the first (last) state antitakeover laws. The data on the SAL are adopted from Armstrong et al (2012)[51]. Following Custodio (2014) I delete observations from the financial industry (SIC 6000-6999), agriculture (SIC lower than 1000), government services (SIC 9000), other noneconomic activities (SIC 8600 and 8800), and unclassified services (SIC 8900). I further require that firms have non-missing data on the variables used in the analysis (see below), and that firms have total assets of at least \$10 million.

In order to measure the conservativeness of the financial reporting of firms I construct a measure for accounting conservatism similar to the *c_score* (Khan and Watts, 2009), which has been widely adopted in the accounting literature (*e.g.*, Kim et al, 2013; Balakrishnan et al, 2016). More specifically, it is estimated as the Basu (1997) cross-sectional regression

$$X_i = \beta_0 + \beta_1 D_i + \beta_2 R_i + \beta_3 D_i R_i + \varepsilon_i \qquad (2)$$

where i indexes the firm, X is earnings, R is returns, D is a dummy variable equal to 1 when $R < 0$ and equal to 0 otherwise, and ε is the residual. β_2 and β_3 respectively measure the timeliness of good and the *incremental* timeliness of bad news (*i.e.*, conservatism). To estimate the timeliness of both good and bad

51 Armstrong et al. (2012) obtain the year of enactment of state antitakeover laws from Bertrand and Mullainathan (2003), but add the strength of said laws.

news at the *firm-year* level, Khan and Watts (2009) further specify that both β_2 and β_3 are linear functions of *Size*, *Market-to-Book* and *Leverage* each year:

$$g\text{-}score = \beta_2 = \mu_0 + \mu_1 Size_i + \mu_2 M/B_i + \mu_3 Lev_i \qquad (3)$$

$$c\text{-}score = \beta_3 = \lambda_0 + \lambda_1 Size_i + \lambda_2 M/B_i + \lambda_3 Lev_i \qquad (4)$$

Note that the estimate of μ_i *and* λ_i ($i = 0 - 3$) are constant across firm, but vary over time. Substitution of equation *(4)* and equation *(5)* into equation *(1)* yields

$$\begin{aligned} X_i = {} & \beta_0 + \beta_1 D_i + \\ & R_i(\mu_0 + \mu_1 Size_i + \mu_2 M/B_i + \mu_3 Lev_i) + \\ & D_i R_i(\lambda_0 + \lambda_1 Size_i + \lambda_2 M/B_i + \lambda_3 Lev_i) + \\ & (\delta_0 Size_i + \delta_1 M/B_i + \delta_2 Lev_i + \delta_3 D_i Size_i + \delta_4 D_i M/B_i + \\ & \delta_5 D_i Lev_i) + \varepsilon_i \qquad (5) \end{aligned}$$

Given my estimation strategy, I slightly adjust the above estimation, so as to make the estimates more comparable for firms across different industries, by allowing the empirical estimators of μ_0 and λ_0 to vary across industries:

$$g\text{-}score = \beta_2 = \mu_j + \mu_1 Size_i + \mu_2 M/B_i + \mu_3 Lev_i \qquad (6)$$

$$c\text{-}score = \beta_3 = \lambda_j + \lambda_1 Size_i + \lambda_2 M/B_i + \lambda_3 Lev_i \qquad (7)$$

where *j* indexes the industry. This leads to:

$$X_i = \beta_0 + \beta_1 D_i +$$

$$R_i\,(\mu_j + \mu_1 Size_i + \mu_2 M/B_i + \mu_3 Lev_i) +$$

$$D_i R_i\,(\lambda_j + \lambda_1 Size_i + \lambda_2 M/B_i + \lambda_3 Lev_i) +$$

$$(\delta_0 Size_i + \delta_1 M/B_i + \delta_2 Lev_i + \delta_3\, D_i\, Size_i + \delta_4\, D_i\, M/B_i + \delta_5\, D_i\, Lev_i) + \varepsilon_i \qquad (8)$$

Finally, in line with the above stated building up a commitment to conservatism over several years, I estimated equation (8) using 5-year rolling windows instead of annually.

4.4 Results

4.4.1 Descriptive Statistics

Table 1 reports the descriptive statistics of the sample. The mean (median) value of the accounting conservatism measure is 1.44 (1.45). Note that this measure is held constant for all firm-years at its before adoption value, which partially accounts for the low standard deviation (0.20) relative to the mean. The mean and median values of this measure remain fairly the same when estimated for each firm-year (not reported), which is in line with stickiness of firms' accounting conservatism[52]. The mean value of *Strength* indicates that 77.5% of the observations are from firms incorporated in states that adopt strong state antitakeover laws as defined by Armstrong et al (2012). Note that this is mainly due to firms incorporated in Delaware, which

52 Not surprisingly, the standard deviation nearly doubles.

comprise almost 60% of the overall sample observations. Finally, Table 1 indicates that average annual acquisitions make up 2.26% of the firms' assets.

4.4.2 Event Study

I start my investigation by testing whether stock market participants react differently to the news of the passage of SAL conditional on the *ex-ante* measure of conservatism of affected firms. More specifically, I argue that if equity holders expect conservatism to discipline managers, they should react less negatively to the dissemination of the SAL news for firms that have adopted more conservative reporting prior to these unexpected events. In order to test this, I conduct event studies on the dates of the first news reports relating to the passage of SAL in Delaware. I focus on three dates to mitigate any concerns regarding anticipation. That is, while the news regarding the adoption of the antitakeover laws was made public on January 27th 1988, news regarding the Delaware Bar Association's recommendation was made public on January 5th 1988 and December 22nd of 1987. As it is common in Delaware that recommendations of the Bar Association for a change in corporate law are swiftly enacted by the state legislature Herzel (1988)[53], it is likely that equity holders already price in the effects of antitakeover laws prior to the actual passage date.

Table 2 reports the results of the regressions of the 5-day cumulative abnormal returns (i.e., [-2, 2]) on the *ex-ante* measure of conservatism. For each of the three dates, the regression is run for both the full sample of Delaware

[53] This was also the case with the state antitakeover laws: only 22 days after the Bar Association's recommendation, Delaware's House of Representatives passed the proposal by 40 to 0 (Reuters, 1988).

firms, as well as a sample that is truncated at 5% (based on the CARs) in order to mitigate the effect of outliers. The results strongly indicate that stock markets react more positively to the news of SAL when firms report more conservatively. For instance, on January 5th 1988, the date eliciting the strongest reactions, cumulative abnormal returns were 1.6% higher for every one standard deviation increase in the firms' conservative reporting. Note that the market reaction should be interpreted as the anticipated *net* effects of conservatism on investment efficiency: equity holders should price in the effects of both the potential disciplinary effects of conservatism, as well as any potential dysfunctional effects on efficient risk taking. As such, the market's positive reaction is in line with stock markets attributing a positive role to conservatism in the efficiency of investment decisions. Note however that these results only speak to how equity holders *revalue* conservatism in situations where managers have *increased* discretion, i.e., are shielded from the market of corporate control. One should therefore be cautious in generalizing these finding to normal circumstances.

4.4.3 Main Results

Table 3 presents the results of the tests of Hypothesis 1. Column 1 reports the estimates of Equation 1, which is the regression of acquisition activity of firms between 1980 and 1995 on the indicator for the adoption of state antitakeover legislation (*SAL*) and an interaction term between *SAL* and accounting conservatism measure (*SAL x Conserv*) in the year prior to the adoption of state antitakeover law. The positive and significant coefficient of

SAL indicates that the less conservatively reporting firms[54] increase their acquisition activity after state antitakeover are adopted in their state of incorporation. Note that under the assumption that these laws did not affect firms' investment opportunity set, this finding can be tentatively interpreted to be in line with managerial overinvestment – this issue will be revisited in section 4.4.5 Also note that the increase of 1.04% is almost half of the overall sample average of annual acquisitions, indicating the strong economic significance of state antitakeover laws[55].

Given the aim of this study, the main variable of interest is the interaction of *SAL* and the *ex-ante* conservatism measure *Conserv*. More specifically, whereas the coefficient of *SAL* captures the differential response of two firms that operate in the same state, but where only one of these firms is incorporated in a state that passes an antitakeover law, the coefficient of the interaction term captures the moderating effect of conservatism on the differential response. The negative and significant coefficient indicates that – in line with Hypothesis 1 – the acquisition activity of firms that reported more conservatively prior to the adoption is affected less by the passage of the state antitakeover laws. In fact, a two standard deviation increase (0.40) in the level of conservatism fully offsets the increased acquisition activity due to the adoption as estimated by the coefficient of *SAL* (1.05% versus 1.04%, respectively)[56].

[54] To facilitate interpretation, the minimum value of conservatism has been subtracted from each observation.

[55] This is even more surprising as state antitakeover laws – if anything – are aimed at preventing acquisitions undesired by target firms' management.

[56] Another way to demonstrate the economic significance of conservatism on acquisition activity is to use the interquartile range: firms at the 3rd quartile of conservative reporting increased their acquisition activity 40% *less* than those at the 1st quartile. This should mitigate

In addition to the mitigating effect of conservatism on inefficient overinvestment, the negative coefficient of the interaction term can also be – tentatively – interpreted as mitigating the incentives of managers to (inefficiently) reduce risk. In particular, Gormley and Matsa (2016) report that after the adoption of state antitakeover laws, managers of affected firms take on inefficiently reduce their firm's riskiness by way of acquisitions (of unrelated businesses and cash rich targets). Note that this interpretation stands in strong contrast to the potential dysfunctional effects of conservatism on efficient risk taking. Caution is however required when interpreting the coefficient of the interaction term with regards to risk taking for two reasons. Firstly, it is possible that even though more conservatively reporting firms on average increase their overall acquisition activity less than less conservatively reporting firms, they may disproportionally increase their risk reducing acquisitions. Secondly, the potential dysfunctional effect of conservatism on efficient risk taking is unobservable (*i.e.*, *forgoing* risky but positive NPV investments), and therefore the test results reported in Table 3 are not a direct test of this hypothesis.

Columns 2a and 2b report the estimates of alternative specification aimed at testing Hypothesis 1 that are more akin to specifications of prior work (*e.g.*, Balakrishnan et al, 2016). More specifically, column 2a reports the results of a specification similar to Equation 1, but one that differs in the fixed-effects that are used (*i.e.*, instead of state-year and industry-fixed effects only year fixed effects are used). The specification reported in column 2b further includes additional control variables (*i.e.*, *Q* and *Cash Flow*). The results

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concerns that the effect of conservatism is driven by outliers that affect the standard deviation of the conservatism measure.

materially remain the same, which holds that the findings are not sensitive to the specification used. Overall, in line with Hypothesis 1, the results of Table 3 suggest that conservatism disciplines managerial behaviour when insulated from the threat of takeover.

4.4.4 Robustness

The results presented in Table 3 may suffer from several issues regarding the composition of the sample and identification. In this section the analyses aimed at addressing these concerns are reported and discussed. The first concern is that given that approximately 60% of the sample firms are incorporated in Delaware, the previously reported findings are mainly driven by this subsample. I address this concern by estimating Equation 1 for the sample that remains after dropping all the observations from firms incorporated in Delaware. As reported in column (1) of Table 4, the results are materially unaffected. A second concern with regard to the sample composition is the inclusion of observations that are too long after the adoption of the state antitakeover laws. To address this issue, I restrict the sample period to only the 11 years surrounding the adoption (*i.e.*, the 5 years before and after). The results of this analysis are reported in column (2) of Table 4, and again remain highly similar to those of the full sample.

I next turn to concerns regarding identification. Firstly, I conduct a test to validate the causal effect of state antitakeover laws by exploiting the variation in the strength of these laws across states. More specifically, if the state antitakeover laws indeed insulate managers from takeover threats and are therefore the reason for increased acquisition activity, we should observe more pronounced effects for firm incorporated in states where the adopted laws were stronger than in states where the impact of the state antitakeover laws was limited. Towards this end, following Armstrong et al (2012), I divide the sample into observations from states with so-called weak and strong antitakeover laws. The results are reported in column (3a) and (3b) of Table 4. In line with the above stated expectations, the results strongly indicate both that state antitakeover laws increase managerial propensity to acquire, as well as that conservatism has a disciplinary effect on the increased managerial discretion in investment decisions. In particular, whereas in the subsample of firms subject to limited changes in the protection against takeover threats no discernable change in acquisition activity occurs (column 3a), a highly significant increase (both statistically and economically) emerges for firms in states where the adoption materially insulated managers (column 3b). Accordingly, the mitigating effect of conservatism is only significant in the subsample of firms with a material change in takeover protection. Note that the insignificant effect of the coefficient on the interaction term in column (3a) is especially informative regarding identification. That is, this regression is tantamount to a placebo test using random non-event years. In particular, a potential concern could be that conservatism always obstructs acquisitions, regardless of the underlying incentive (*i.e.*, potential efficiency). Assuming that takeover threats positively affect acquisition efficiency, the insignificant coefficient of the interaction term dispels this concern.

An identifying assumption is that state antitakeover laws unexpectedly increased managerial investment discretion, allowing me to use the pre-adoption levels of conservatism as an exogenous source of variation. Prior studies that investigate direct effects of state antitakeover laws often consider the possibility of reverse causality as a threat to the exogeneity condition, arguing the possibility that the passage of the laws may have been the result of lobbying efforts by firms that may have benefited from their adoption (Bertrand and Mullainathan, 2003; Armstrong et al, 2012). As these firms would anticipate the adoption of these laws, they could take actions that affect the outcome of interest. While several studies investigate this issue and dismiss the lobbying argument (*e.g.*, Romano, 1987; Bertrand and Mullainathan, 2003; Armstrong et al, 2012), I address this concern following the methodology advanced by Bertrand and Mullainathan (2003) and used by several others towards this end (e.g., Armstrong et al, 2012; Valta, 2012). More specifically, I replace the *SAL* (and the interaction counterparts) with four indicator variables: *SALm1* is an event time indicator variable that equals one if the firm is incorporated in a state one year before the adoption of state antitakeover laws and zero otherwise, *SALt0* is an event time indicator variable that equals one if the firm is incorporated in a state during the year of adoption and zero otherwise, *SALp1* is an event time indicator variable that equals one if the firm is incorporated in a state one year after the adoption of state antitakeover laws and zero otherwise, and *SAL2plus* is an event time indicator variable that equals one if the firm is incorporated in a state two or more years after the adoption of state antitakeover laws and zero otherwise. A positive and significant coefficient on *SALm1* would be an indication of causality running in the opposite direction, as it would entail that increased acquisition activity precedes the adoption. As reported in column (4) of Table 4, the estimated coefficients of *SALm1* and *SALt0* are nevertheless insignificantly different from zero. Moreover, the coefficients of the first three event time variables gradually increase in both magnitude and significance, in line with the previous studies' findings that the antitakeover laws were not anticipated and suggest that the adoption preceded the increase in acquisition activity. Furthermore, consistent with the tests of the strength of antitakeover laws (column 3a), the insignificant coefficient on both the *SALm1* and its associated interaction with *Conserv* can be interpreted as a placebo test: in years without the frictions imposed by state antitakeover laws, conservatism does not impede acquisition activity. Note that this argument needs not to hold *after* the adoption: as targeted firms can require higher premiums to accept a bid when insulated by antitakeover laws (Comment and Schwert, 1995), a significantly negative coefficient on the interaction term can no longer be interpreted unambiguously.

Another concern is that measures of conservatism merely pick up the effect of other governance mechanisms such as the quality of the firm's management team. For instance, Balakrishnan et al (2016) argue the possibility that better and more able managers are more likely to report more conservatively as they better understand the benefits of conservative reporting. As it is also probable that more skilled managers make better investment decisions, the previously reported mitigating effect of conservatism on acquisition activity could merely reflect the superior managerial ability. In line with this argument, García Lara et al (2009) show that firms with strong governance report more conservatively. In order to mitigate this concern, I follow Balakrishnan et al (2016) and re-run the previous test while directly controlling for managerial quality. More specifically, using two measures for managerial quality – *i.e.*, the *MA-Score* by Demerjian et al (2012) and industry-adjusted operating performance (Balakrishnan et al, 2016), both measured in the year prior to the adoption – I split my sample in firms with high (above median) and low (below median) managerial quality, and re-run the regression based on Equation 1. As reported in Table 5, for both measures of managerial quality firms with low managerial quality show a much stronger increase in post-adoption acquisitions activity (i.e., the coefficient on *SAL* in columns 1a vs 1b and 2a vs 2b). Note that the coefficient of *SAL* in column 1b even becomes significantly negative, in line with the conjecture that good managers are less willing to acquire targets that can require high premiums in the post-adoption period. More importantly, while the absolute magnitude and *t*-statistics of the coefficient on *SAL x Conserv* are larger for the low managerial quality subsamples (column 1a and 2a), the mitigating effect of conservatism is still negative for the high managerial quality subsamples (1b and 2b) and significant when managerial quality is measure by industry-adjusted operating performance (and barely insignificant when managerial quality is measure by the *MA-Score*). I further use an alternative specification (unreported) in which the managerial quality proxies are included as interaction terms. The estimated coefficients of *SAL x Conserv* remain negative and significant, revealing that the mitigating effect of conservatism on acquisition activity is distinct from managerial quality.

A limitation of the specification used in the previous analyses is that firms from states where *no* antitakeover laws were adopted cannot be used as a control group, as the staggered implementation of antitakeover laws renders it impossible to delineate the pre and post period. In order to circumvent this, I estimate Equation 1 separately for each year during which antitakeover laws were adopted as a final robustness test. More specifically, given that the pre and post period is now imposed by the year of adoption, I can compare the change in acquisition behaviour of firms from each cohort of state antitakeover laws with all the firms from states where such laws were not passed. The results are reported in Table 6, and the estimated coefficients are in line with the previous findings. More specifically, all the estimated coefficients have the same sign as the results in Table 2, and are statistically different from zero for all but two year (*i.e.*, 1988 and 1989), providing further evidence of the robustness of the main findings.

4.4.5 Consequences

The main focus of this study is to ascertain the effect of conservatism on the *efficiency* of investment decisions. While the results of the previous analyses strongly suggest that firms that report less conservatively exhibit an increase in acquisition activity after the adoption of state antitakeover laws, it does not necessarily follow that the increased acquisition activity is inefficient. Put differently, the change in the *level* of acquisition activity, does not speak unambiguously to its consequences for the acquirers' performance and risk profile. For instance, it is possible that state antitakeover laws allow managers to pursue more risky, yet (in the long run) more profitable acquisitions, as these laws reduce the threat of forced turnover due to the absence of short term gains[57]. Especially given the proposed potential dysfunctional effects of conservatism on efficient risk taking (Roychowdhury, 2010; Kravet, 2014), it would be premature to label an increase in acquisition activity as inefficient.

[57] Theoretical arguments regarding the effect of takeover protection on risk taking and innovation are mixed. On the one hand, it is argued that when takeover threat is high, managers may be less inclined to tie their human capital to the firm which would otherwise foster the creation of innovative products (Shleifer and Summers, 1988). On the other hand, it is argued that takeover threats mitigate moral hazard issues which if left unchecked would lead to less innovation (Jensen and Ruback, 1983). The empirical evidence is equally mixed (see Chemmanur and Tian (2018) and Atanassov (2013) for empirical evidence in line with the respective theoretical arguments). Lel and Miller (2015) provide evidence that antitakeover laws increase the propensity to replace poorly performing CEOs.

This section therefore discusses the results of the tests aimed at uncovering the effect of conservatism on changes in firms' performance (Hypothesis 2) and riskiness (Hypothesis 3) after the adoption of state antitakeover laws.

Table 7 presents the estimation results of Equation 1, where the dependent variable is either a measure of firm performance aimed at capturing overall investment efficiency, or a measure of firm riskiness. More specifically, for the former I use industry-adjusted return on assets (measured both before and after depreciation deduction) and yearly stock returns, whereas the latter is measure as the yearly standard deviation of the stock returns. The results strongly suggest that conservatism has a disciplinary effect on managerial decisions. That is, the coefficient of *SAL* is significantly negative in all the regressions with performance measures as the dependent variable. More importantly with respect to Hypothesis 2, the coefficient of *SAL x Conserv* is significantly positive for the regressions of one of the profitability measures and the stock returns. Most salient are the results in column (4) of Table 7: while less conservatively reporting firms reduce their riskiness in line with managers' abuse of investment discretion to reduce their firm's risk (Gormley and Matsa, 2016), more conservatively reporting firms effectively take on *more* risk, a finding that is very hard to reconcile with the proposed potential dysfunctional effect of conservatism on efficient risk taking. Thus, overall the results presented in Table 7 indicate that conservatism prevents managers from pursuing their inefficient desire to reduce their firm's risk, validating the more positive stock market reactions for conservatively reporting firms upon the news of state antitakeover laws adoptions.

4.5 Conclusion

Extant empirical accounting research mostly focusses on the role of accounting in the supply of information for valuation and monitoring purposes, but (implicitly) regards the outcome of the underlying economic activity pursued by firms as independent of the accounting method used. More recent work endogenizes the role of accounting by recognizing that the quality of information provision by firms can improve investment efficiency by mitigating both underinvestment, through reduction of information asymmetry between firms and external supplier of capital, and overinvestment, by facilitating contracting and monitoring (Biddle and Hilary, 2006; Biddle et al, 2009).

This study extends this line of research by focusing on the role of one specific principle in accounting, *i.e.*, conservatism, on investment efficiency. The reason for the focus on conservatism is twofold. Firstly, despite the central role of conservatism in accounting, there are both contrasting theoretical predictions and mixed empirical evidence regarding the effect of conservatism on investment efficiency. Secondly, the aforementioned literature ascribes the beneficial role of higher accounting quality mostly to its role in reducing information asymmetry between parties internal and external to the firm. The effect of conservatism on investment efficiency however stems from its asymmetric reporting of good and bad outcomes coupled with managerial incentives and factors that affect managerial investment discretion (*e.g.*, covenants). In other words, while information asymmetry gives rise to the need for conservatism, the mechanism through which conservatism affects investment efficiency is *distinct* from its effect on information asymmetry.

Exploiting the staggered and unanticipated passage of state antitakeover laws in order to circumvent endogeneity concerns, I find evidence strongly in line with a disciplinary effect of conservatism on managerial investment discretion. More specifically, I find that investors react less negatively to an increase in managerial discretion for firms that report more conservatively. Using a difference-in-difference setup, I find that firms that report more conservatively do not increase their acquisition investments, while those reporting less conservatively do. Furthermore, while both the operating profitability, stock performance and riskiness of less conservatively reporting firms decline after increases in managerial discretion, more conservatively reporting firms' performance is unaffected. Overall, the evidence of this chapter suggests that accounting conservatism mitigates inefficient investment that can be attributed to increased managerial discretion.

Table 1: Summary Statistics.

This table reports the summary statistics for the main sample, which refers to all firm-years between 1980 till 1995 for which the conservatism measure could be calculated. *Conservatism* is based on the *c-score* measure of Khan and Watts (2009) and is calculated as discussed in section 3.2. Note that it is estimated in the last fiscal year before the adoption of state antitakeover laws by the state of incorporation, and is subsequently held constant for all firm-years. *Strength* refers to whether the change in the *SAL* (state antitakeover law) was high, as defined by Armstrong et al. (2012). *Acquisitions* is the ratio of acquisitions to total assets. *Cash Holdings* is the ratio of cash and short-term investments to total assets. *Cash Flow* is the ratio of operating income before depreciation to lagged total assets. *Leverage* is the ratio of the sum of debt in current liability and long-term debt to the market value of equity. *Profitability1* is the industry-year (at 2 digit SIC level) adjusted ratio of operating income after depreciation to lagged total assets, and *Profitability2* is constructed similarly but based on operating income before depreciation. *Q* is the market value of assets to book value of assets following Kaplan and Zingales (1997). *Return Volatility* is the yearly standard deviation of the stock returns for each fiscal year. *Stock Returns* are the yearly stock returns.

	Mean	*St.Dev*	*Q1*	*Median*	*Q3*	*N*
Conservatism	1.444	0.20	1.367	1.454	1.523	24967
Strength	0.775	0.42	1.000	1.000	1.000	24967
Acquisitions	2.263	7.90	0.000	0.000	0.278	24967
Cash Holdings	0.114	0.14	0.019	0.057	0.155	24964
Cash Flow	0.153	0.13	0.090	0.147	0.214	24967
Leverage	0.570	0.83	0.078	0.288	0.723	24967
Profitability1	0.018	0.12	-0.035	0.013	0.072	24967
Profitability2	0.021	0.13	-0.038	0.014	0.078	24967
Q	1.488	0.92	0.980	1.203	1.657	24307
Return Volatility	0.029	0.02	0.018	0.025	0.036	20006
Stock Returns	0.162	0.45	-0.124	0.093	0.361	22158

Table 2: Event study.

This table reports the estimates from regressions of 5-day cumulative abnormal returns (*i.e.*, [-2; 2]) of three events related to the dissemination of news regarding the adoption of state antitakeover laws in Delaware on *Conservatism* (calculated as discussed in section 3.2).The samples consist either of all firms incorporated in Delaware at the time of the news (indicated by *Full*) or 90% of those firms as a results of a symmetric 5% truncation based on the 5-day cumulative abnormal returns (indicated by *Truncated*). *t*-statistics are presented in parentheses and standard errors are clustered at the state of incorporation level. ***, **, and * denote significance at the 1%, 5%, and 10% level for two-tailed tests, respectively.

	22-Dec-87		*05-Jan-88*		*27-Jan-88*	
	Full	*Truncated*	*Full*	*Truncated*	*Full*	*Truncated*
Constant	-0.114**	-0.04	-0.446***	-0.285***	-0.145***	-0.061**
	(-2.21)	(-1.17)	(-7.64)	(-7.03)	(-3.43)	(-2.36)
Conservatism	0.093***	0.039	0.341***	0.225***	0.097***	0.036**
	(2.64	(1.64)	(8.51)	(8.06)	(-3.33)	(2.01)
Adj. R^2	0.00	0.00	0.05	0.05	0.01	0.00
N	1,495	1,352	1,495	1,349	1,495	1,352

Table 3: Acquisition activity and accounting conservatism before and after adoption of SAL.

This table reports the estimates from firm-level panel regressions of acquisition activity (defined as the ratio of acquisitions to total assets) of firms between 1980 and 1995 on the indicator for the adoption of state antitakeover legislation (*SAL*) and an interaction term between *SAL* and accounting conservatism measure (*SAL x Conserv*) in the year prior to the adoption of *SAL*. *SAL* takes on the value of 1 in the years of and after the adoption of state antitakeover laws, and zero otherwise. For the construction of accounting conservatism, see caption of Table 1. Model (1) follows the methodology of Gormley and Matsa (2016), which includes firm, state-year and industry-year fixed effects. Model (2) includes firm and (fiscal) year fixed effects. *Q* and *Cash Flow* are defined as in Table 1. *t*-statistics are presented in parentheses and standard errors are clustered at the state of incorporation level. ***, **, and * denote significance at the 1%, 5%, and 10% level for two-tailed tests, respectively.

	(1)	*(2)*	
		(a)	*(b)*
SAL	1.040***	0.690**	0.806**
	(2.85)	(2.56)	(2.43)
SAL x Conserv	-2.627***	-1.539***	-0.906*
	(-5.61)	(-4.50)	(-1.97)
Q			0.369***
			(8.25)
Cash Flow			3.299***
			(4.79)
Firm FE	Yes	Yes	Yes
Year FE	No	Yes	Yes
State-year FE	Yes	No	No
Industry-year FE	Yes	No	No
Adj. R^2	0.11	0.10	0.10
N	24,967	24,967	20,719
N Clusters	30	30	30

Table 4: Robustness.

This table reports the estimates from firm-level panel regressions of acquisition activity (defined as the ratio of acquisitions to total assets) of firms between 1980 and 1995. In columns (1), (2), (3a) and (3b), acquisitions activity is regressed on the indicator for the adoption of state antitakeover legislation (*SAL*) and an interaction term between *SAL* and accounting conservatism measure (*SAL x Conserv*) in the year prior to the adoption of *SAL* (both defined as in table 3). The observations in these regressions however vary; (1) excludes observations of firms incorporated in Delaware, (2) only includes observations that are no more than 5 years away from the adoption of state antitakeover laws, (3a) and (3b) only include observations of firms from states that passed either weak or strong antitakeover laws respectively. Column (4) includes all firm-year observation, but acquisition activity is regressed on indicator variables that replace *SAL* (and the interaction counterparts) with four indicator variables: *SALm1* is an event time indicator variable that equals one if the firm is incorporated in a state one year before the adoption of state antitakeover laws and zero otherwise, *SALt0* is an event time indicator variable that equals one if the firm is incorporated in a state during the year of adoption and zero otherwise, *SALp1* and *SAL2plus* are an event time indicator variables that equals one if the firm is incorporated in a state either one, or two or more years after the adoption of state antitakeover laws respectively, and zero otherwise. All models include firm, state-year and industry-year fixed effects and standard errors are clustered at the state of incorporation level. *t*-statistics are presented in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% level for two-tailed tests, respectively.

	w/o DE	*(-5/5+)*	*SAL*		*Dynamic*
			Weak	Strong	
	(1)	*(2)*	*(3a)*	*(3b)*	*(4)*
SAL	1.183*	0.961***	-0.021	1.781***	
	(1.85)	(2.92)	(-0.02)	(4.29)	
SAL × Conserv	-2.473***	-2.802***	-1.87	-3.646***	
	(-3.77	(-4.38)	(-1.15)	(-4.85)	
SALm1					-0.416
					(-0.52)
SALt0					1.057
					(1.66)
SALp1					1.397*
					(1.78)
SAL2plus					0.724
					(1.01)
SALm1 × Conserv					0.300
					(0.28)
SALt0 × Conserv					-3.219***
					(-4.70)
SALp1 × Conserv					-2.974***
					(-2.91)
SAL2plus × Conserv					-2.418***
					(-3.79)
Adj. R^2	0.07	0.11	0.08	0.12	0.11
N	10,798	18,618	5,358	19,309	24,967
N Clusters	29	30	16	14	30

Table 5: Managerial quality.

This table reports the estimates from firm-level panel regressions of acquisition activity (defined as the ratio of acquisitions to total assets) of firms between 1980 and 1995 on the indicator for the adoption of state antitakeover legislation (*SAL*) and an interaction term between *SAL* and accounting conservatism measure (*SAL x Conserv*) in the year prior to the adoption of *SAL*. *SAL* takes on the value of 1 in the years of and after the adoption of state antitakeover laws, and zero otherwise. For the construction of accounting conservatism, see caption of Table 1. The observations in these regressions however vary; column (1a) and (1b) include observations that had respectively below or above median values of the *MA-Score* (see Demerjian et al (2012) for construction) in the year before the adoption of antitakeover laws in their state of incorporation. Column (2a) and (2b) include observations that had respectively below or above median values of 2-digit SIC industry-adjusted profitability (ratio of operating income after depreciation to lagged total assets) in the year before the adoption of antitakeover laws in their state of incorporation. All models include firm, state-year and industry-year fixed effects and standard errors are clustered at the state of incorporation level. *t*-statistics are presented in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% level for two-tailed tests, respectively.

	Low MA-Score	*High MA-Score*	*Low IA. Profit*	*High IA. Profit*
	(1a)	*(1b)*	*(2a)*	*(2b)*
SAL	3.044***	-0.894*	1.549***	0.541
	(5.97)	(-1.80)	(2.83)	(1.47)
SAL x Conserv	-4.326***	-1.154	-3.093***	-2.038**
	(-8.47)	(-1.70)	(-4.38)	(-2.63)
Adj. R^2	0.12	0.10	0.11	0.12
N	11,465	11,433	11,173	13,450
N Clusters	27	29	27	28

Table 6: Year of adoption.

This table reports the estimates from firm-level panel regressions of acquisition activity (defined as the ratio of acquisitions to total assets) of firms between 1980 and 1995. Note that – different that the other tables – the sample not only includes firms from states that adopted antitakeover laws, but also from states that never did. Each column contains observations from states that adopted state antitakeover laws in the year as indicated by the column title, and firms from states that never adopted antitakeover laws. Acquisition activity is regressed on the indicator for the adoption of state antitakeover legislation (*SAL*) and an interaction term between *SAL* and accounting conservatism measure (*SAL x Conserv*) in the year prior to the adoption of *SAL*. Note that *SAL* takes on the value of 1 in the years of and after the year indicated in the column title, and zero otherwise. For the construction of accounting conservatism, see caption of Table 1. All models only include firm fixed effects and standard errors are clustered at the state of incorporation level. *t*-statistics are presented in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% level for two-tailed tests, respectively.

	1985	*1986*	*1987*	*1988*	*1989*	*1990*	*1991*
SAL	2.220***	2.919***	1.662*	0.582	0.01	1.479**	0.64
	(4.80)	(6.64)	(1.87)	(0.98)	(0.02)	(2.54)	(1.44)
SAL x Conserv	-5.134***	7.494***	4.162**	-2.627	-0.438	-9.256***	-5.849***
	(-3.51)	(-4.65)	(-2.04)	(-1.48)	(-0.15)	(-3.49)	(-4.09)
Adj. R^2	0.12	0.14	0.15	0.12	0.11	0.11	0.11
N	16,894	16,398	15,836	22,457	7,951	4,897	3,873
N Clusters	47	46	45	39	31	24	21

Table 7: Consequences.

This table reports the estimates from firm-level panel regressions of operating profitability, stock performance or riskiness measures of firms between 1980 and 1995 on the indicator for the adoption of state antitakeover legislation (*SAL*) and an interaction term between *SAL* and accounting conservatism measure (*SAL x Conserv*) in the year prior to the adoption of *SAL*. *SAL* takes on the value of 1 in the years of and after the adoption of state antitakeover laws, and zero otherwise. For the construction of accounting conservatism, see caption of Table 1. The specific dependent variables are indicated as column titles and are constructed as described in the caption of Table 1. All models include firm, state-year and industry-year fixed effects and standard errors are clustered at the state of incorporation level. *t*-statistics are presented in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% level for two-tailed tests, respectively.

	Profitability	*Profitability2*	*Stock Returns*	*Return Volatility*
	(1)	*(2)*	*(3)*	*(4)*
SAL	-0.012**	-0.016***	-0.076***	-0.003**
	(-2.43)	(-3.42)	(-2.77)	(-2.63)
SAL x Conserv	0.008	0.016*	0.152**	0.006***
	(0.98)	(1.81)	(2.70)	(3.61)
Adj. R^2	0.45	0.47	0.17	0.66
N	24,678	24,673	21,816	19,966
N Clusters	30	30	30	30

English summary

A primary function of capital markets is to efficiently allocate capital, which entails the flow of capital to investments with the highest returns commensurate with their risk (Tobin, 1984). The market of corporate assets fulfills a similar function, where ideally productive assets are transferred to those most equipped to manage them (Manne, 1965). A large body of scientific inquiry identifies informational frictions as impediments to the efficient functioning of these markets. To mitigate these adverse effects, firms provide financial information to parties external to the firm. For the most part, this information is generated by the firms' accounting function and its presentation and disclosure are guided by accounting standards. The studies comprising this dissertation empirically investigate several aspects of the role financial accounting information plays in asset and capital markets.

Chapter 2 investigates the voluntary provision of information in the market of corporate assets sales and finds evidence which suggests that these disclosures are used to signal the turnaround of poor prior performance and financial distress. Chapter 3 investigates the effect of changes in accounting standards (i.e., IFRS) on the asymmetric distribution of information amongst capital market participants. The results indicate that IFRS adoption in cross-listed firm's domestic market improves the liquidity of ADRs, in line with the reduction of the information disadvantage of investors trading on U.S.

exchanges. Finally, chapter 4 finds that a commitment to providing conservative accounting information disciplines the allocation of capital by managers when state legislation increases managerial discretion.

Nederlandse samenvatting

(Summary in Dutch)

Een primaire functie van kapitaalmarkten is het efficiënt alloceren van kapitaal, hetgeen inhoudt dat kapitaal naar investeringen stroomt met het hoogste rendement gegeven de bijbehorende risico (Tobin, 1984). De markt voor activa vervult een vergelijkbare functie, waarbij idealiter productieve activa worden overgedragen aan diegenen die het best zijn toegerust om ze te beheren (Manne, 1965). Echter, zoals in andere aspecten van het leven, beletten fricties het optimaal functioneren van markten, waardoor de voordelen voor de samenleving niet altijd ten volle benut worden. Een groot aantal wetenschappelijke studies heeft zich gewijd aan het identificeren en analyseren van deze fricties, en stelt vast dat de asymmetrische verdeling van informatie tussen verschillende partijen, gekoppeld aan hun uiteenlopende belangen, leidt tot verhoogde financieringskosten (Jensen en Meckling, 1976), kredietbeperking (Stiglitz and Weiss, 1981), of zelfs een complete ineenstorting van markten (Akerlof, 1970). Om deze nadelige effecten te reduceren, verstrekken bedrijven financiële informatie aan externe partijen. Deze informatie wordt grotendeels gegenereerd door het accountingsysteem van de bedrijven en de rapportering ervan wordt geleid door verslaggevingsstandaarden. De studies die dit proefschrift omvat, onderzoeken empirisch verschillende aspecten van de rol die financiële informatie speelt in markten voor activa en kapitaal.

Hoofdstuk 2 onderzoekt de informatievoorziening op de markt voor activa. Een groot deel van de overdracht van activa tussen bedrijven bestaat uit de verkoop van bedrijfsonderdelen, waarbij bedrijven een deel van hun activiteiten afstoten en andere behouden. Uit eerdere onderzoek blijkt dat het afstoten van bedrijfsonderdelen wordt gebruikt om het domein van de activiteiten van de onderneming te wijzigen (Maksimovic en Phillips, 2001), waarbij activa worden verkocht aan bedrijven die ze efficiënter kunnen inzetten (Hite et al, 1987). Bovendien fungeert de verkoop van bedrijfsonderdelen als een primaire financieringsbron (Lang et al., 1995; Bates, 2005, Arnold et al, 2017), wat de verkopende bedrijven in staat stelt hun aandacht te richten op de activiteiten waar ze de meeste waarde kunnen toevoegen. Hoewel uit de literatuur blijkt dat het afstoten van bedrijfsonderdelen voor zowel verkoper als koper waarde toevoegt (Eckbo en Thorburn, 2013), is er nog onvoldoende kennis van het proces van verkoop van bedrijfsmiddelen (Borisova et al, 2013). Wat bekend is, is dat er een gebrek is aan voldoende openbare informatie over de kwaliteit van bedrijfsonderdelen, aangezien de rapportageverplichtingen met betrekking tot specifieke delen van bedrijven minder streng zijn dan die voor de onderneming als geheel, en bedrijven niet geneigd zijn om dergelijk gedetailleerde informatie vrijwillig te verstrekken uit zorg voor hun concurrentiepositie (Botosan en Stanford, 2005). Dit betekent dat potentiële kopers mogelijk hoge zoekkosten moeten maken, wat op zijn beurt leidt tot een minder efficiënte toewijzing van bedrijfsmiddelen. Een openbare aankondiging dat bepaalde bedrijfsonderdelen beschikbaar zijn voor verkoop kan worden ingezet om zoekkosten te verminderen en het aantal potentiële kopers te vergroten, waardoor de kans op een efficiëntere toewijzing van bedrijfsmiddelen toeneemt.

Het onderzoek naar de informatievoorziening door bedrijven tijdens het proces van verkoop van bedrijfsonderdelen leidt tot de volgende bevindingen. We laten zien dat in 42% van de voltooide transacties het verkopende bedrijf haar voornemen om te verkopen van tevoren aankondigt, en stellen vast dat deze vooraankondigingen leiden tot economisch en statistisch significante positieve marktreacties. Onze analyses geven verder aan dat deze vooraankondigingen worden gebruikt om de ommekeer van slechte prestaties en financiële problemen te signaleren. Bovendien duiden onze resultaten er enigszins op dat niet vooraf aangekondigde transacties de verkoop van zeer gewilde bedrijfsonderdelen betreffen, mogelijk geïnitieerd door buitenlandse bieders. Deze schijnbaar verschillende aanleidingen voor de twee soorten transacties komen ook overeen met onze belangrijkste bevinding dat markten positiever reageren op vooraf aangekondigde transacties dan op niet vooraf aangekondigde transacties. Meer specifiek, vooraf aangekondigde transacties impliceren dat de toekomst van de resterende activiteiten, hetgeen de meerderheid van de activiteiten van de verkopende onderneming vormt, zal verbeteren. Het rendement op de niet vooraf aangekondigde transacties lijkt echter een premie te weerspiegelen voor het afgestoten bedrijfsonderdeel, dat over het algemeen een minderheidsdeel van de verkopende onderneming vormt. Een belangrijke implicatie van onze resultaten is dat investeerders vooraankondigingen geloofwaardig achten en dat het merendeel van de waarderingseffecten van de vooraf aangekondigde afstotingen in de aandelenkoers van de verkopende onderneming worden opgenomen voorafgaand aan de aankondiging van de transactie. Dit, in combinatie met onze resultaten die suggereren dat de beslissing om de intentie om een onderdeel af te stoten van tevoren aan te kondigen gerelateerd is aan de motivatie voor de verkoop, heeft implicaties voor empirische analyses van de verkoop van bedrijfsonderdelen.

Hoofdstuk 3 onderzoekt het effect van wijzigingen in verslaggevingsstandaarden op de asymmetrische verdeling van informatie tussen beleggers, zoals kan worden afgeleid uit veranderingen in de transactiekosten. Geconfronteerd met het risico te handelen met partijen die beter geïnformeerd zijn, beschermen beleggers zich door een hogere (lagere) prijs te vragen (bieden) of zich geheel te onthouden van handel, hetgeen een optimale toewijzing van risico en kapitaal belet. De bezorgdheid om minder informatie te hebben is met name van belang wanneer de handel in effecten van een bedrijf op meer dan één beurs plaatsvindt, waarbij één beurs dichter bij het hoofdkantoor van de onderneming ligt. Dit komt doordat de meeste waarde relevante informatie in de nabijheid van die beurs wordt gegenereerd, vaak wordt gecommuniceerd in de taal van het land van herkomst van het bedrijf, en is opgesteld in overeenstemming met de verslaggevingsstandaarden van het land van herkomst van het bedrijf (Halling et al., 2007), hetgeen leidt tot een informatieachterstand van de beleggers op buitenlandse beurzen. De invoering van International Financial Reporting Standards (IFRS) in meer dan 120 landen, wat een van de grootste wijzigingen in verslaggevingsstandaarden tot op heden vormt, stelt ons in staat om het effect van verslaggevingsstandaarden op internationale kapitaalstromen te onderzoeken.

In een steekproef van 239 bedrijven met niveau 2 of 3 ADRs uit 31 landen, waarvan 27 tussen 2003 en 2012 IFRS hebben ingevoerd, observeren we dat IFRS-implementatie de liquiditeit van ADR's verbetert, hetgeen duidt op een vermindering van de informatieachterstand van beleggers die handelen op Amerikaanse beurzen. Onze resultaten geven verder aan dat de verbetering van de liquiditeit van ADRs afhangt van de kwaliteit van de nationale juridische en regulerende instellingen. Analyses die gericht zijn op het identificeren van de bron van de verbeteringen tonen niet aan dat de superieure kwaliteit van IFRS ten opzichte van de reeds bestaande binnenlandse GAAP's de liquiditeitsverbeteringen verklaren, maar wijzen in de richting van de schaalvoordelen die voortvloeien uit het verminderen van het aantal verslaggevingsstandaarden volgens welke cross-listed bedrijven rapporteren. Gezamenlijk duiden onze resultaten erop dat de invoering van IFRS op de binnenlandse markt van een aan een Amerikaanse beurs genoteerde onderneming de toegang verbetert tot buitenlandse markten die het mandaat niet hebben ingevoerd en dat de kapitaalallocatie van Amerikaanse beleggers die beperkt zijn tot het beleggen op beurzen in de VS mogelijk verbetert. Onze bevindingen bieden verder nieuwe inzichten in de rol van verslaggevingsstandaarden bij concurrentie tussen beurzen.

Ten slotte onderzoek ik in hoofdstuk 4 of de toewijding om conservatief te rapporteren de allocatie van kapitaal door managers disciplineert. Bestaand empirisch onderzoek richt zich voornamelijk op de rol van accounting in de informatievoorziening voor waarderings- en controledoeleinden, maar beschouwt veelal de uitkomst van de onderliggende economische activiteiten van bedrijven als onafhankelijk van de gebruikte verslaggevingsmethode. Meer recent werk endogeniseert de rol van de verslaggeving door te erkennen dat de kwaliteit van informatieverstrekking door bedrijven de efficiëntie van investeringen kan verbeteren, zowel onderinvestering te verminderen als gevolg van een reductie in de informatie-asymmetrie tussen bedrijven en externe leveranciers van kapitaal, als door overinvestering te verminderen door het vergemakkelijken van contracteren en monitoren (Biddle en Hilary, 2006; Biddle et al, 2009). Dit hoofdstuk breidt deze onderzoekslijn uit door zich te richten op het effect van conservatieve verslaggeving op de efficiëntie van bedrijfsinvesteringen, aangezien er, ondanks de centrale rol van conservatisme in verslaggeving, zowel tegenstrijdige theoretische voorspellingen als gemengd empirisch bewijs bestaat met betrekking tot deze relatie.

Terwijl ik gebruik maak van de gespreide en onverwachte implementatie van overnamebeschermingswetten in verschillende staten van de V.S. om endogeniteitsproblemen te verminderen, vind ik bewijs dat duidt op een disciplinerend effect van conservatisme op de discretionaire bevoegdheid van managers. Meer specifiek stel ik vast dat beleggers minder negatief reageren op een toename van de discretionaire bevoegdheid van managers voor bedrijven die conservatiever rapporteren. Met behulp van een "difference-in-difference" analyse concludeer ik dat bedrijven die conservatiever rapporteren hun investeringen in overnames niet verhogen, terwijl bedrijven die minder conservatief rapporteren dat wel doen. Bovendien, terwijl zowel de operationele winstgevendheid, als het rendement en risico op aandelen van minder conservatieve rapporterende bedrijven afnemen na een toename van de discretionaire bevoegdheid van managers, blijven deze voor conservatiever rapporterende bedrijven onaangetast. Over het geheel genomen signaleren de resultaten van dit onderzoek dat een toewijding om conservatief te rapporteren inefficiënte investeringen die voort kunnen vloeien uit toegenomen bestuurlijke discretie belet.

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- Corporate Finance Theory
- Corporate Governance
- Empirical Corporate Finance
- English
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- Microeconomics
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- Publishing Strategy
- Statistical Methods

Teaching activity:

- Financial Accounting Research (Accounting – master core course)
- Research Skills Accounting and Control (Accounting – master core course)
- Taxation (Accounting – master elective)
- Research Skills (Accounting – master)
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- Corporate Finance (Finance – bachelor)
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A primary function of capital markets is to efficiently allocate capital, which entails the flow of capital to investments with the highest returns commensurate with their risk (Tobin, 1984). The market of corporate assets fulfills a similar function, where ideally productive assets are transferred to those most equipped to manage them (Manne, 1965). A large body of scientific inquiry identifies informational frictions as impediments to the efficient functioning of these markets. To mitigate these adverse effects, firms provide financial information to parties external to the firm. For the most part, this information is generated by the firms' accounting function and its presentation and disclosure are guided by accounting standards. The studies comprising this dissertation empirically investigate several aspects of the role financial accounting information plays in asset and capital markets.

Chapter 2 investigates the voluntary provision of information in the market of corporate assets sales and finds evidence which suggests that these disclosures are used to signal the turnaround of poor prior performance and financial distress. Chapter 3 investigates the effect of changes in accounting standards (i.e., IFRS) on the asymmetric distribution of information amongst capital market participants. The results indicate that IFRS adoption in cross-listed firm's domestic market improves the liquidity of ADRs, in line with the reduction of the information disadvantage of investors trading on U.S. exchanges. Finally, chapter 4 finds that a commitment to providing conservative accounting information disciplines the allocation of capital by managers when state legislation increases managerial discretion.

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