Journal of Accounting and Economics ▮ (▮▮▮▮) ▮▮▮–▮▮▮

Contents lists available at ScienceDirect

Journal of Accounting and Economics

journal homepage: www.elsevier.com/locate/jae

# Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism ☆

Xiumin Martin [a], Sugata Roychowdhury [b,*]

[a] Washington University, St. Louis, USA
[b] Boston College, USA

ARTICLE INFO

*Article history:*
Received 22 January 2014
Received in revised form
4 September 2014
Accepted 12 September 2014

*JEL classification:*
G1
G2
G21
G30
M40
M41
M44

*Keywords:*
Credit default swaps
Timely loss recognition
Conservatism
Financial market developments
Lender monitoring

ABSTRACT

This paper investigates whether the initiation of trading in credit default swaps (CDSs) on a borrowing firm's outstanding debt is associated with a decline in that firm's reporting conservatism. CDS investments can modify lenders' payoffs on their loan portfolios by providing insurance on negative credit outcomes. The onset of CDS trading reduces lenders' incentives to continuously monitor borrowers and also their demand that borrowers report conservatively. Additionally, borrowers expect CDS-insured lenders to be more intransigent in renegotiations triggered by defaults and covenant violations. Since conservatism can trigger earlier covenant violations, borrowers have heightened incentives to report less conservatively in the post-CDS period. Using a differences-in-differences research design, we observe a decline in borrowing firms' reporting conservatism after CDS trade initiation. This effect is more pronounced when reputation costs lenders face from reducing monitoring are lower, when debt contracts outstanding at the time of CDS trade initiation have more financial covenants, and when lenders who monitor borrowers more regularly in the pre-CDS period enter into CDS contracts to hedge their credit exposures.



☆ This paper has benefited immensely from comments by Guojin Gong. This paper has also benefited from comments by Richard Frankel, Ewa Sletten, Ross L. Watts, Regina Wittenberg-Moerman, Jerry Zimmerman (editor), an anonymous referee, workshop participants at Arizona State University, Columbia University, London Business School, George Mason University, Georgetown University, Georgia State University, New York University, Pennsylvania State University, Washington University at St. Louis, University of Maryland, UCLA, University of Missouri at St. Louis, Dartmouth University, Singapore Management University, University of Texas at Dallas, and participants at the American Accounting Association conference, the Nick Dopuch conference at Washington University, the University of Minnesota Empirical Accounting Research Conference, Harvard University, Yale School of Management Conference on Accounting and the Financial Crisis, and CAPANA Conference. All remaining errors are ours.

* Corresponding author.
*E-mail addresses:* xmartin@wustl.edu (X. Martin), sugata.roychowdhury@bc.edu (S. Roychowdhury).

http://dx.doi.org/10.1016/j.jacceco.2014.09.006


Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

## 1. Introduction

This paper investigates the influence of financial developments in the capital markets on accounting practices. Financial market developments can affect the payoffs and incentives of existing contractual parties to the firm. Since contractual considerations can influence accounting practices (Watts and Zimmerman, 1986), an altered contractual environment can potentially induce changes in these practices. The financial market innovation we focus on is credit default swaps (CDSs), widely used in recent times to manage lenders' credit risk exposures and their regulatory capital. The accounting practice we study is conservatism, in the context of the borrower–lender relationship. Our primary interest is in examining whether the advent of CDS trade initiation on a firm's outstanding debt is associated with a change in that firm's reporting conservatism.

The credit default swap is a contract in which the buyer, generally called the protection buyer, makes a series of payments to the seller, generally called the protection seller. In exchange, the protection buyer receives a payoff from the protection seller if a credit instrument (such as a loan or a bond) goes into default or experiences any other "credit event" specified in the CDS contract (such as restructuring, bankruptcy, or credit-rating downgrade). By acquiring a CDS contract, the protection buyer transfers the credit risk associated with its investment (such as a loan or a bond) to the protection seller, while retaining legal ownership of the investment. The risk-shifting via CDS contracts allows lenders, particularly banks, to better manage their regulatory capital since the risk weight assigned to a loan can be based on the credit rating of the counter-party in the CDS contract rather than the original borrower.[1] As an example, AIG discloses in its Annual Report that $150 billion of its notional CDSs outstanding at the end of 2009 reflected contracts it wrote to provide regulatory capital relief to financial institutions for their corporate loans (Saretto and Tookes, 2013). The overall CDS market has grown tremendously in recent years, with the notional amount increasing from $180 billion in 1998 to $57 trillion at the end of June 2008 (Stulz, 2010).[2]

Investments in CDS contracts by banks can potentially have an influence on the reporting practices of those clients. Upon granting a loan, lenders generally face an asymmetric payoff on their investment: if the borrowing firm remains solvent, lenders receive their principal and earned interest, while bankruptcy entitles them to the orderly liquidation value of the borrower. The literature argues that this asymmetric payoff underlies lenders' demand for conservatism in the financial statements of borrowers (Watts and Zimmerman, 1986; Watts, 2003). Under conservative reporting, which requires stricter verification standards for recognizing good news in earnings relative to bad news, the book value of a firm provides lenders with a lower-bound estimate for the firm's orderly liquidation value. Ensuring that borrowers do not deviate from conservative reporting practices post-loan-initiation arguably requires continuous monitoring by lenders over the life of the loan. Indeed, continuous lender monitoring, in particular by banks, seems to be a salient feature of the traditional lender–borrower relationship (Gorton and Khan, 1993; Roberts and Sufi, 2009; Acharya et al., 2014).

The availability of CDS contracts alters lenders "downside" payoffs and can thus influence the lender–borrower relationship. In the event of borrower insolvency (in practice, any pre-specified credit event in the CDS contract), lenders are now entitled to settlement payouts from CDS sellers. Coverage from a CDS contract thus reduces the asymmetry in the payoffs to lenders' claims, and provides them greater bargaining power upon the occurrence of pre-specified credit events such as defaults and violations (Bolton and Oehmke, 2010). Lenders' less asymmetric claim structure post-CDS, and their higher bargaining power in renegotiations, potentially diminish their reliance on continuous monitoring to protect the value of their claims and relatedly, their demand for conservatism in borrowers' financial reports. Furthermore, lenders' reduced reliance on continuous monitoring is expected to be accompanied by higher intransigence on their part in renegotiations with borrowers who experience credit events (Fink, 2004; Hu and Black, 2006; Bolton and Oehmke, 2010; Stulz, 2010; Subrahmanyam et al., 2014). Since conservative accounting policies are associated with earlier covenant violations (Zhang, 2008; Nikolaev, 2010), borrowers have increased incentives to report less conservatively after CDS trade initiation, because they anticipate tougher renegotiations if they trigger covenant violations. The joint effect of borrowers' incentives to avoid renegotiations and lenders' incentives to avoid monitoring costs can lead to less conservative reporting by borrowers after CDS trade initiation.

A post-CDS reduction in borrowers' reporting conservatism need not necessarily be a foregone conclusion as it can make third parties such as CDS sellers apprehensive about borrowers' credit quality. Note that CDS sellers do not own control rights with respect to the underlying loan and typically eschew any direct contractual involvement with borrowers. Nevertheless, it is possible that lenders maintain their demand for conservatism, to avoid reputation costs (for example, with CDS sellers) arising from negative credit event realizations that are attributed to their reduced monitoring of financial statements. Further, it may be difficult for borrowers to deviate from past conservatism for the sake of maintaining reporting consistency. Other stakeholders to the firm unprotected by CDS investments, such as shareholders and lenders, may step up their monitoring to ensure that borrowers continue to report conservatively. These alternative scenarios illustrate the importance of empirically investigating changes in borrowers' reporting conservatism upon CDS trade initiation.

[1] BASEL II states that guarantees issued by or protection provided by entities with a lower risk weight than the counterparty exposure is assigned the risk weight of the guarantor or protection provider.

[2] The size of the CDS market fell sharply in the second half of 2008 in the wake of the financial crisis, but was still high at $41 trillion at the end of 2008. The Bank for International Settlements (BIS) has statistics on the CDS market since the end of 2004 based on survey data. See http://www.bis.org/statistics/derstats.htm.

Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

We identify a sample of 529 firms who experience CDS trade initiation between 2002 and 2009. The empirical exercise involves the identification of an event, CDS trade initiation, and an examination of whether there is a change in conservatism around this event. To control for selection bias and endogeneity, we follow the Ashcraft and Santos (2009) matched-sample design. In the first stage of this analysis, we predict CDS trade initiation based on credit rating, leverage, profit margin, size, return volatility and market-to-book.[3] In the second stage, we use the CDS-trade-initiation probability-estimates yielded by the first stage model to construct a sample of 525 unique matched firms from the same industry as the treatment firms, but with no CDS trading throughout the sample period. Finally, in a difference-in-difference regression, we compute the change in conservatism for CDS firms from the two years prior to CDS-trade-initiation to the two years after, and compare that change to the corresponding change for matched non-CDS firms.

We define conservatism as the asymmetric requirement of weaker verification standards for recognizing losses than for recognizing gains. We measure conservatism as the greater timeliness of earnings with respect to negative returns relative to positive returns (Basu, 1997). Our results indicate that the initiation of CDS trading is on average associated with a decline in the financial reporting conservatism of underlying borrowers. The specific coefficients imply that asymmetric timeliness of loss recognition declines by around 20% after the onset of CDS trading. The decline seems economically significant and is statistically significant at the 5% level.

In cross-sectional analyses, we test predictions on when a post-CDS decline in conservatism is more likely. We reason that a decline in reporting conservatism should be more prominent when the underlying borrowers are ex ante more informationally opaque and riskier. Among such borrowers, lenders are likely to face higher monitoring costs prior to CDS trade initiation, and lower reputation costs from the potentially adverse consequences of reducing their demand for conservatism after CDS trade initiation (Parlour and Winton, 2013; Ashcraft and Santos, 2009). Further, if lenders reduce their demand for conservatism in financial statements, then borrowers with poorer *ex ante* credit quality are expected to be more responsive to such a reduction. The evidence indicates that the decline in conservatism is concentrated among borrowers that are smaller and carry speculative-grade credit ratings prior to CDS trade initiation.

Our cross-sectional tests also indicate a greater decline in conservatism at the time of CDS trade initiation when borrowers' outstanding private debt at that time includes more financial covenants. Since conservative reporting is hypothesized to accelerate covenant violations (Zhang, 2008), our finding is consistent with borrowers having higher incentives to avoid violations after CDS trade initiation, presumably because they expect greater lender-intransigence in the associated renegotiations.

We reason that the influence of lenders' incentives to avoid monitoring costs in the post-CDS period will manifest more in firms that are subject to continuous lender monitoring in the pre-CDS period. Banks represent the class of lenders most likely to continuously monitor borrowers, via their sophisticated and timelier private insights into various aspects of a firm's managerial practices, including their reporting choices. Banks also provide the data necessary in their Y-9C reports to identify those among them with a greater likelihood of having entered into CDS contracts on their clients. We find that a decline in conservatism after CDS trade initiation is more pronounced among borrowers with loans outstanding from banks that are likely to have hedged their credit exposures via CDS contracts.

Additional tests examine the possibility that the observed association between CDS trade initiation and a decline in conservatism is endogenous, in the sense that lenders are motivated to enter into CDS contracts when they anticipate a decline in reporting conservatism at the borrower. Our empirical analysis indicates that the decline in borrower conservatism we observe is unlikely driven by the change in conservatism anticipated at the time of CDS trade initiation. Indeed this result is confirmed again when we substitute the matched-sample method with a Heckman two-stage procedure to control for selection bias, and when we employ an expanded prediction model for CDS trade initiation that incorporates banks' tendency to offload credit risk (Minton et al., 2009). Our results are not sensitive to alternative identification of matched control firms. Finally, they are also robust to using a conservatism measure that relies on earnings time-series properties (Basu, 1997) and not on equity returns.

The primary contribution of our study is in providing evidence that financial market developments influence financial reporting practices because they alter contractual parties' payoffs and incentives. In particular, CDS trade initiation in a firm's outstanding debt leads to a decline in its reporting conservatism because it alters the firm's relationship with its lenders.

The rest of the paper proceeds as follows. Section 2 discusses related literature and develops testable hypotheses. Section 3 describes sample selection and presents descriptive statistics. Section 4 introduces variable measurement and empirical methods. Section 5 reports empirical findings. We conclude in Section 6.

## 2. Related literature and hypothesis development

### 2.1. The CDS market

Single-name CDSs are typically written on specific securities issued by firms, for example public bonds outstanding. CDS buyers make payments to CDS sellers via insurance premiums expressed as an annualized percentage of the notional value

[3] The first-stage model indicates that CDS trade initiation is more likely for firms with higher credit rating, leverage, profit margin and size, and lower return volatility, consistent with adverse selection concerns in the CDS market creating a bias towards firms with higher credit quality and transparency.

Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

of the transaction. For example, if the CDS spread of the underlying firm is 0.5%, a bank buying $10 million worth of protection from the CDS seller must pay the seller $50,000 annually. The payments continue until either the CDS contract expires or until the occurrence of a pre-specified credit event (e.g., default, bankruptcy, credit-rating downgrade or restructuring). See Appendix A for a sample CDS contract. There are typically two types of CDS sellers: (a) monoline insurers such as AIG and Ambac who primarily operate on the sell-side and (b) financial institutions and hedge funds including J.P. Morgan and Goldman Sachs who serve as market-makers. Both types of CDS sellers hedge their open risk exposures on CDSs across various derivative instruments and across numerous investors in each instrument (Weistroffer, 2009).

CDS contracts provide a convenient channel for hedging to lenders, even though the loans responsible for lenders' credit risk exposures to the underlying borrowers are distinct from the specific securities (i.e., bonds) that CDS contracts are written on.[4] The CDS protection benefits banks with respect to regulatory requirements. BASEL II states that by entering into CDS contracts, a bank can substitute the credit risk of the borrower by the credit risk of the CDS seller in computing risk-weighted assets (BASEL II, page 49, Article 141). CDS purchases can therefore allow for lower commitment of regulatory capital to the loan, which in turn frees funds for alternative productive investments. The CDS insurance also allows originating lenders to maintain lending relationships with their borrowers while reducing the risk profile of their loan portfolios (Venokur et al., 2008; Saretto and Tookes, 2013).

CDSs are derivative instruments, and hence are available for trade not only to lenders seeking insurance on their loan exposures, but also to speculators (Stulz, 2010; Lewis, 2010). Nonetheless, there is evidence that banks are increasingly using the CDS market to hedge the credit exposures they originate through their lending business.[5] In 2006, banks entering into CDS contracts related to their loan portfolios constituted 20% of the market for CDS purchases. By comparison, banks writing CDS contracts on their loan portfolios constituted only 9% of total CDS selling, implying that banks in general maintain net purchase positions, reflecting their incentives to use CDSs for hedging and managing regulatory capital (British Bankers Association (BBA), 2006, also see Appendix B).

### *2.2. CDS contracts, lender monitoring and borrowers' conservatism: primary hypothesis*

CDS contracts allow loan originators to “economize” on their regulatory capital and share their risk exposures with the rest of the economy (Deutsche Bank Research, 2009). The literature has investigated whether lenders pass on the benefits from CDS investments to their clients. Ashcraft and Santos (2009) fail to find evidence that CDS trade initiation is associated negatively with interest spreads that lenders demand from their corporate borrowers. However, Saretto and Tookes (2013) document that S&P 500 firms with CDS contracts trading on their debt are able to maintain higher leverage ratios and longer debt maturities, consistent with such firms benefiting from fewer supply-side frictions in lending.

A traditional lender–borrower relationship is typically characterized by lending institutions, in particular banks, continuously monitoring borrowers after loan initiation (Fama, 1985; James, 1987; Roberts and Sufi, 2009; Acharya et al., 2014). A key question in the literature has been whether the development of the CDS market weakens lenders' incentives to monitor borrowers (Duffee and Zhou, 2001; Morrison, 2005; Ashcraft and Santos, 2009; Marsh, 2009; Stulz, 2010; Parlour and Winton, 2013). A similar issue exists with other credit-risk-transfer mechanisms such as loan sales in the primary and secondary market (Pennacchi, 1988; Gorton and Pennacchi, 1995; Ball et al., 2008; Bushman and Wittenberg-Moerman, 2012). The CDS market, however, differs from the loan sale market in some important respects. In a loan sale, both the risk exposure on the loan and control rights, including the right to monitor and administer the loans, are typically transferred to the loan buyer.[6] Even in cases that loan sales are partial, as when lead arrangers bring in syndicate participants, the latter are in a better position to detect any shirking in monitoring by the originating lender than CDS sellers, who do not have any direct access to the borrower. In contrast, in a CDS contract, the credit risk transfers to the CDS seller, but control rights remain with the original lender. Thus, the agency issues are potentially more severe with CDS contracts (Parlour and Winton, 2013).

CDS availability significantly alters the lender–borrower dynamic. Consider a bank that invests in a CDS contract on an underlying borrower's bonds. Upon the occurrence of a pre-specified credit event, for example a payment default, renegotiations are less crucial for the bank to preserve the value of its claim, because of its existing CDS insurance. This provides the bank incentives to (a) be more inflexible in renegotiations and (b) rely on credit event triggers rather than continuous monitoring to ensure the value of its claim. Hu and Black (2006) and Bolton and Oehmke (2010) refer to this as the phenomenon of the “empty creditor.” Subrahmanyam et al. (2014) document that bankruptcy probability and credit risk assessments by rating agencies both increase with CDS trade inception, a finding that persists after controlling for selection bias therein, as well post-CDS changes in firm fundamentals such as leverage. They attribute the heightened financial risk of borrowers to lenders' intransigence in the post-CDS period.

In the context of the traditional lender–borrower relationship, accounting conservatism is hypothesized to provide an efficient means for debt-holders to monitor their credit risk (Watts and Zimmerman, 1986; Basu, 1997; Watts, 2003; Frankel

[4] For example, JP Morgan Chase reports $48 billion in notional CDS purchases to hedge the credit risk of its loan portfolio in its 2009 Annual Report (Saretto and Tookes, 2013).

[5] According to a survey by the British Bankers Association (2006) half of the protection banks bought in the CDS market in 2005 and 2006 were to cover exposures resulting from their lending activities.

[6] Loan sales without recourse constitute the vast majority of transactions (see Gupta et al., 2008).

Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

and Litov, 2008; Zhang, 2008; Nikolaev, 2010; Ahmed et al., 2002; Gormley et al., 2012; Tan, 2013; Donovan et al., 2014). Conservative financial reporting, by recognizing economic losses in a timely manner and deferring the recognition of economic gains, ensures that borrowers' net asset values are understated. Thus, under conservatism, net asset values provide a lower bound on borrowers' ability to repay their debt obligations (Roychowdhury and Watts, 2007). Further, asymmetrically timely loss recognition under conservative reporting is expected to accelerate debt covenant violations and facilitate timelier transfer of control to debt-holders (Zhang, 2008; Nikolaev, 2010). Borrowers' incentives to report less conservatively after loan origination to avoid covenant violations are typically mitigated by continuous monitoring of their financial statements by lenders.

If lenders shift away from continuous monitoring because their claims are insured via CDSs, borrowers now have incentives to report less conservatively. Firms do not necessarily observe the timing of their lending banks' investments into specific CDS contracts, but they can observe CDS trade initiation on their own outstanding bonds. Further, lenders' weakened incentives to monitor the conservatism in borrowers' financial statements can manifest in several ways, including fewer requests for timely financial statements, fewer clarification requests regarding those statements and less frequent on-site visits to verify reported numbers.[7] The reduced scrutiny from lenders seeking to lower monitoring costs provides borrowers the opportunity to report less conservatively. The expected intransigence of lenders in post-CDS renegotiations makes violations particularly unattractive for borrowers and provides them with incentives to report less conservatively.[8]

Systematic empirical evidence regarding a decline in lender monitoring upon CDS trade initiation is limited. Marsh (2009) documents a less positive stock return reaction to borrowers announcing new loans from banks known to transfer credit risk via collateralized loan obligations (CLOs), consistent with the market anticipating weaker monitoring by such banks.[9] Ashcraft and Santos (2009) find that debt financing costs are higher for risky and informationally opaque firms after the onset of CDS trading, which they interpret as evidence of a reduction in lender monitoring among this subset of firms.

Wang and Xia (2014) document that firms borrowing from banks active in loan securitizations (via CLOs) enjoy looser covenants at loan origination, and appear to take on more risk than those borrowing from non-securitization banks. They conclude that banks exert less monitoring effort after loan securitizations, but do not investigate CDSs. Inconsistent with Wang and Xia (2014), Sustersic (2012) finds that new debt agreements in the post-CDS period are more likely to include financial covenants, with less “slack,” relative to those initiated in the pre-CDS period. Interestingly, Sustersic (2012) finds no evidence of increased covenant violation probability in the post-CDS period; her results raise the possibility that borrowers report less conservatively after CDS trade initiation, and thus avoid an otherwise higher level of covenant violation, given the more numerous and tighter covenants. To our knowledge, the specific issue of whether lenders reduce their scrutiny of their borrowers' financial statements upon acquiring CDS contracts, or whether borrowers exhibit any change in their reporting practices, is unaddressed in the literature. Our paper tests the following null hypothesis:

**H1 (null).** The onset of CDS trading in a firm is not associated with a reduction in the firm's reporting conservatism.

H1 is stated in null form because other economic forces exist that could possibly influence the change in lender monitoring and borrower conservatism upon CDS trade initiation, but the net influence of these forces is often ambiguous. For example, it is possible that the demand for conservatism arising from parties other than banks is also altered with the availability of CDS contracts. Parties such as shareholders (with the board of directors and auditors as their fiduciary agents) and public debt-holders not invested in CDSs possibly increase their monitoring of financial statements and their demand for conservatism in anticipation of reduced lender monitoring. However, it is not clear that such parties can completely substitute for the expert monitoring by lenders, in particular banks. Further, CDS contractual provisions can specify that the bank's claim on the underlying firm is junior to those of other parties, motivating banks to retain their demand for conservatism (Sufi, 2007). In spite of this, such modifications in practice tend to be rare; CDS counterparties typically rely on the International Swaps and Derivatives Association Master Agreement to draw up contracts (see Deutsche Bank Research, 2009, and the discussion of standard CDS contracts in Appendix A). Similarly, on the one hand, it is possible that the responsibility for monitoring of financial statements passes from lenders to CDS sellers. On the other hand, the absence of private contractual agreements between CDS sellers and underlying borrowers limits the ability of the former to monitor borrowers on an ongoing basis after CDS trade initiation. Rather, CDS sellers, the largest of whom are monoline insurers, typically establish diversified portfolios of credit risk in which losses generated by one contract are compensated by premiums earned from other contracts. Finally, CDS sellers can price-protect in anticipation of reduced lender monitoring and can also increase CDS prices for lenders if they observe heightened post-CDS borrower default. Nevertheless, the implications of price-protection by CDS sellers are far from obvious. CDS-sellers typically price-protect only on average

[7] As Arping (2012) argues, managers at borrowing firms can typically detect any weakening of monitoring intensity in general.

[8] Other factors can also motivate managers at borrowing firms to report less conservatively in the absence of lender monitoring. Managerial compensation is often linked to earnings, for example via bonus plans. Conservative reporting, by delaying the recognition of gains relative to losses, introduces a deferred component to managers' compensation (Leone et al., 2006). It is also argued that reporting conservatism restricts managers' ability to operate or invest in projects that are potentially detrimental to the firm's health but generate private benefits for managers (Ball and Shivakumar, 2005). In addition, Roychowdhury (2010) points to the possibility that conservative reporting can weaken managers' incentives to invest in risky projects.

[9] Marsh (2009) does not observe the same evidence with banks investing in CDSs, but cautions that his sample is not appropriate for the test, since it excludes firms actively traded in the CDS market.

Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

across both lenders and speculators *ex ante*, and cannot always attribute *ex post* borrower defaults to reduced lender monitoring. We test this latter possibility in our cross-sectional analysis, discussed next.

### 2.3. CDS contracts and underlying borrowers' conservatism: cross-sectional hypotheses

#### 2.3.1. The role of reputation costs

If lenders reduce their demand for conservatism after acquiring CDSs, they may face reputation costs with CDS sellers and other loan syndicate participants (in cases that the lenders are also lead arrangers) if the loan subsequently performs poorly. This can respectively manifest in higher CDS prices in subsequent transactions with CDS sellers and/or lower willingness of potential syndicate participants to participate in future loans. Thus, lenders' propensity to lower monitoring of financial statements after entering into CDS contracts is expected to be greater when reputation effects are weaker.

Parlour and Winton (2013) argue that reputation effects in the CDS market will be weaker when riskier borrowers are involved. The intuition is as follows: if borrowers already deemed to be riskier were to default or experience any other credit event, it is more difficult for external parties to attribute this negative outcome to a lack of lender monitoring. Consequently, CDS-protected lenders are more likely to reduce monitoring of riskier borrowers. The evidence in Ashcraft and Santos (2009) suggesting a more pronounced decline in lender monitoring among riskier borrowers post-CDS-trade initiation is consistent with Parlour and Winton (2013). In the loan syndication market, Gopalan et al. (2011) find that the reputation loss suffered by lead arrangers in the event of borrower bankruptcies is lower when outstanding loans to the insolvent borrowers already have high yields (consistent with these loans being deemed as high-risk).

Furthermore, since continuous monitoring of riskier and opaque borrowers should be costlier, we expect that lenders protected by CDS investments will find it more efficient to rely on pre-specified credit events to trigger renegotiations or payments by CDS sellers. In turn, riskier borrowers will generally tend to be closer to triggering covenant violations and renegotiations, and hence are expected to be more responsive to any decline in lenders' demand for conservatism. Thus, we expect any decline in conservatism after CDS-trade initiation to be more pronounced for riskier borrowers with more opaque information environments. We identify firms that are smaller in size and have credit ratings below investment grade as riskier firms with lower-quality information environments. Thus, we test the following prediction:

**H2 (alternate).** A decline, if any, in conservatism after the onset of CDS trading is more pronounced for smaller borrowers and borrowers with credit rating below investment grade.

#### 2.3.2. The role of covenants

Zhang (2008) documents that firms who report more conservatively are timelier in violating covenants upon the realization of a negative event, proxied for by a negative price shock. Nikolaev (2010) documents a positive association between conservatism and the presence of financial covenants in public debt contracts, interpreting this as evidence of the complimentary role they play in facilitating timely transfer of control to lenders. In general, the literature suggests that conservatism is more likely to facilitate transfer of control to lenders in the presence of financial covenants. Upon CDS trade initiation, we expect borrowers to have greater incentives to lower conservatism in response to reduced lender scrutiny when their existing debt contracts include more financial covenants. This is particularly true because borrowers rationally expect lenders to be more inflexible in renegotiations triggered by covenant violations when lenders are CDS-protected (Hu and Black, 2006; Bolton and Oehmke, 2010; Subrahmanyam et al.,2014). In formulating our hypothesis, we focus on the number of financial covenants in private debt contracts with banks. Private debt contracts are much more likely to include financial covenants than public debt (Begley and Freedman, 2004; Chava and Roberts, 2008), and banks/financial institutions are also more likely to hedge their underlying exposures via sophisticated derivative instruments such as CDSs (Acharya and Johnson, 2007).

**H3 (alternate).** A decline, if any, in conservatism after the onset of CDS trading is more pronounced for borrowers with a larger number of financial covenants in their existing private debt contracts at the time of CDS trade initiation.

#### 2.3.3. Lender identity and continuous monitoring

CDS trade initiation and its influence on lenders' monitoring of financial statements is relevant only when the lender–borrower relationship is characterized by continuous monitoring in the first place. In other words, if post-loan initiation, lenders rely not on continuous monitoring but on credit events such as defaults or violations to trigger further scrutiny of borrowers even in the pre-CDS period, then CDS investments have little scope of altering lenders' monitoring strategy. We identify borrowers that are subject to regular monitoring by their lenders in the pre-CDS period by focusing on CDS firms that have loans outstanding to banks. Fama (1985) describes banks as providers of "inside debt", in that they have the greatest access to private information and are most likely class of lenders to monitor the firm closely after loan initiation. Consistent with this James (1987) finds a positive equity return upon the announcement of new bank credit agreements. Additionally, consistent with continuous monitoring after loan initiation, Roberts and Sufi (2009) document that a large majority (75%) of private credit agreements are renegotiated, with only a small minority (18%) of these renegotiations motivated by defaults and/or covenant violations. Data available for banks also enables us to identify whether lending banks exhibit a greater likelihood of having invested in CDS contracts on underlying borrowers upon CDS availability. This

Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

identification provides the additional advantage of isolating instances when CDS investments are made by banks hedging underlying loans, rather than by traders assuming speculative positions.

In summary, we test the following hypothesis:

**H4 (alternate).** A decline, if any, in conservatism after the onset of CDS trading is more pronounced among borrowers whose banks exhibit a greater likelihood of having hedged their respective loan exposures via CDS contracts upon CDS availability.

## 3. Sample selection

### *3.1. Firms with traded CDS contracts*

CDS contracts are traded in the over-the-counter (OTC) market, almost entirely populated by institutional investors. Unlike an organized exchange such as the NYSE, the information on CDS trading must be gathered from market participants on the basis of their voluntary participation in periodic surveys. We collect information on CDS contracts from Datastream. Datastream covers approximately 13,000 single-name CDS contracts for firms domiciled in 70 countries. Among U.S. firms, there are 8,041 single-name CDS contracts with either senior debt (93%) or subordinated debt (7%) as the underlying securities.[10] Datastream collects CDS data from two main sources: CMA Datavision CDS series and Thomson Reuters CDS series. We only focus on the CMA CDS series because Mayordomo et al. (2014) find that CMA database quotes lead the price discovery process relative to quotes provided by other databases including Markit, GFI, Reuters EOD and JP Morgan. CMA in turn collects data directly from the trading desks of buy-side CDS market participants. Note that the CMA series are no longer offered through Datastream after the 3rd quarter of 2010. This change does not affect our empirical analysis, as our sample of CDS trade initiations ends in 2009.

We identify 1,193 U.S. firms that have single-name CDS contracts traded between January 2002 and December 2009. The CDS sample ends in 2009 to facilitate computation of asymmetric timeliness of loss recognition after CDS trade initiation for all firms in the sample. For each of these firms, we identify the first fiscal year that the firm trades at least one US-dollar-denominated CDS contract. We merge these 1,193 firms with Compustat and CRSP to collect financial variables used in the subsequent empirical analyses.[11] After deleting financial firms and requiring all firms to have at least one observation during both pre- and post-CDS trade-initiation periods, we are left with 529 unique U.S. non-financial firms with required financial variables.

### *3.2. Matched control firms*

The initiation of CDS contracts balances credit risk preferences between the protection seller and the protection buyer. In particular, firms' credit risk and growth opportunities potentially influence the demand and supply of CDS contracts (Ashcraft and Santos, 2009). To address this sample selection issue, we follow Ashcraft and Santos (2009) and implement the matched-sample design developed in the literature. Specifically, we augment the model in Ashcraft and Santos (2009) and estimate the following logistic model to predict the initiation of CDS trading (firm subscripts are suppressed for brevity):

$$\begin{aligned}\text{Prob}(\text{CDS}_t=1)=\Phi(&\beta_0+\beta_1\text{INVESTMENT GRADE}_{t-1}+\beta_2\text{RATING}_{t-1}+\beta_3\text{LEV}_{t-1}+\beta_4\text{PROFIT MARGIN}_{t-1}\\&+\beta_5\text{SIZE}_{t-1}+\beta_6\text{RETURN VOLATILITY}_{t-1}+\beta_7\text{MB}_{t-1})+\varepsilon_t\end{aligned} \quad (1)$$

where CDS is an indicator variable equal to one for firms with CDSs traded between 2002 and 2009, and zero otherwise. We include INVESTMENT GRADE, RATING, LEV, and PROFIT MARGIN to account for firms' credit risk. INVESTMENT GRADE is an indicator variable equal to one if a firm has an S&P credit rating above BB+, and zero otherwise. RATING is an indicator variable equal to one if a firm has an S&P credit rating, and zero otherwise.[12]

LEV is book leverage, equal to a firm's total debt (short-term debt plus long-term debt) scaled by total assets. PROFIT MARGIN is net income scaled by sales. We also include firm size (SIZE), return volatility (RETURN VOLATILITY), and market-to-book ratio (MB) to consider the effect of overall information environment and growth opportunities on the demand and supply of CDS contracts. SIZE is the natural logarithm of market value of equity. RETURN VOLATILITY is the standard

[10] Single-name CDS contract is one where there is just one reference entity. The reference entity can be any borrower, but is most often one of a few hundred widely traded companies (corporate or financials) or a handful of governments (sovereigns). The CDS contract that we are interested in is the single-name one where the reference entity is a corporation. In addition to the single-name CDSs, there are basket default swaps (BDSs), index CDSs, and funded CDSs (also called a credit-linked notes) etc.

[11] The 1,193 firms include multiple subsidiaries for the same parent holding firms. For such firms, we collect financial variables for the parent holding firms only when merging with the Compustat and CRSP databases.

[12] We have tested robustness to using an ordinal variable capturing the credit rating of the firm, in lieu of the indicator variables INVESTMENT GRADE and RATING. All our subsequent results are robust to this alternative specification of the first stage model.

**Table 1**
Logistic regression results on probability of CDS trade initiation.

| Dependent variable = Prob(CDS = 1) | | | |
|---|---|---|---|
| Variable | Coeff. Est. | | *p*-Value |
| Intercept | −6.485 | | < 0.001 |
| INVESTMENT GRADE | 0.691 | | < 0.001 |
| RATING | 1.356 | | < 0.003 |
| LEV | 1.476 | | < 0.001 |
| PROFIT MARGIN | 0.106 | | < 0.001 |
| SIZE | 0.439 | | < 0.001 |
| RETURN VOLATILITY | −2.201 | | < 0.001 |
| MB | 0.023 | | 0.331 |
| Pseudo $R^2$ | | 0.46 | |
| Model significance | 1,940.55 | | < 0.001 |
| Likelihood ratio | 21,145.97 | | < 0.001 |
| Percent concordant | | 91.50% | |
| Percent discordant | | 8.11% | |
| Number of firm-years | | 138735 | |

This table reports coefficient estimates from estimating a logistic model to predict the onset of credit default swap (CDS) trading. The dependent variable, CDS, is equal to 1 if a CDS is traded on a firm, and 0 otherwise. Independent variables include INVEST GRADE, an indicator variable equal to 1 if a firm has a S&P credit rating above BB+, and 0 otherwise; RATING, an indicator variable equal to 1 if a firm has a S&P credit rating, and 0 otherwise; SIZE, natural logarithm of market value; MB, the ratio of market value of equity to book value of equity; LEV, leverage equal to total debt scaled by total assets; PROFIT MARGIN is net income scaled by sales; RETURN VOLATILITY is standard deviation of monthly stock return within a fiscal year. The sample period spans 1997–2009, containing firms without CDS traded and firms with CDS traded during this period. For firms with CDS traded, only firm-years prior to the onset of CDS trading are included in the sample. Robust standard errors are estimated and are clustered at the firm level.

deviation of monthly stock return within a fiscal year, and MB is the ratio of market value of equity to book value of equity. We use all Compustat firms with available information during the period 1997–2009.[13]

Table 1 reports regression results of estimating Eq. (1). As shown, the model specified in Eq. (1) predicts the onset of CDS trading well, as evidenced by good model fit, high proportion of concordant pairs (91.5%) and low proportion of discordant pair (8.1%). The results indicate that firms with higher credit rating, leverage, profit margin and market cap, along with lower stock return volatility are more likely to have CDS trade initiation during the sample period. These findings are generally in line with an adverse selection explanation: given banks (potential protection buyers) possess superior private information about the debt instruments that they originated, the protection seller is more likely to offer CDS contracts for firms with less credit risk (firms with higher credit rating and higher profit margin) and a more transparent information environment (larger firms). The positive relation between leverage and the likelihood of CDS trade initiation suggests greater market demand of credit risk protection (via CDS contracts) for high-leverage firms.

Next, we utilize a matching procedure to construct a control sample of non-CDS firms (i.e., firms with no CDSs trade during the sample period). Specifically, based on the estimation results of Eq. (1), we obtain the estimated likelihood of CDS trade initiation for all Compustat firms. For each CDS firm (i.e., firms with a CDS trade during the sample period), we identify three non-CDS firms within the CDS firm's two-digit SIC industry that have the closest estimated likelihood to the CDS firm. The comparison of estimated likelihoods is made in the fiscal year prior to the year of CDS-trade-initiation.[14] We allow the same non-CDS firm to be matched to multiple CDS firms to minimize the distance in their estimated probability of CDS trade initiation.[15] However, once assigned as a match, a control (i.e., non-CDS) firm enters the sample only once every year, even if it serves as a match for more than one treatment (i.e., CDS) firm; thus, every control firm-year observation is unique. The matched-sample design generates 525 unique non-CDS-firm matches for the 529 CDS firms.

### 3.3. Descriptive statistics

Table 2 Panel A presents the sample distribution based on the CDS-trade-initiation year for the CDS sample and the matched non-CDS sample. The year 2004 witnessed the largest number of firms with CDS trade initiation (297 firms, or 56.1% of the CDS sample). The number of CDS trade initiations quickly declined afterwards for two reasons. First, since we select only the first traded CDS contract for each underlying firm, by construction we will observe a decline in the number of

[13] Our results are robust to setting the initial year of prediction to 2002, that is, the year of the first CDS trade initiation in our sample (and using corresponding explanatory variables from 2001) and therefore estimating the model parameters over 2001–2009.

[14] Our process of identifying more than one matching non-CDS firm for every CDS firm is similar to that in Lee (1997) and Chen and Martin (2011).

[15] We limit the distance in their estimated probability of CDS trade initiation to within 20% points. As a result, some CDS firms may have fewer than three matching non-CDS firms.

Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

**Table 2**
Sample distribution.

**Panel A: Sample distribution by CDS onset year for both CDS and non-CDS firms**

| | CDS | | Non-CDS | |
|---|---|---|---|---|
| Year | *N* | % | *N* | % |
| 2002 | 1 | 0.19 | 2 | 0.38 |
| 2003 | 128 | 24.2 | 237 | 45.14 |
| 2004 | 297 | 56.14 | 184 | 35.05 |
| 2005 | 41 | 7.75 | 32 | 6.1 |
| 2006 | 13 | 2.46 | 17 | 3.24 |
| 2007 | 41 | 7.75 | 43 | 8.19 |
| 2008 | 3 | 0.57 | 5 | 0.95 |
| 2009 | 5 | 0.95 | 5 | 0.95 |
| Total | 529 | 100 | 525 | 100 |

**Panel B: Sample distribution by industry**

| | CDS | | Non-CDS | |
|---|---|---|---|---|
| Industry (based on 1-digit SIC) | *N* | % | *N* | % |
| Mining, mineral and construction | 57 | 10.78 | 40 | 7.62 |
| Food, apparel, petroleum refining, and paper and printing | 122 | 23.06 | 113 | 21.52 |
| Rubber, stone, computer, transportation equipment | 125 | 23.63 | 130 | 24.76 |
| Railroad transportation and electric and gas | 107 | 20.23 | 121 | 23.05 |
| Retail and wholesale | 55 | 10.40 | 51 | 9.72 |
| Business service | 47 | 8.88 | 55 | 10.47 |
| Public service | 13 | 2.46 | 12 | 2.28 |
| Government service | 3 | 0.57 | 3 | 0.58 |
| Total | 529 | 100 | 525 | 100 |

This table reports sample distribution by the CDS onset year in Panel A and by industry in Panel B, for both CDS firms and their matched firms (Non-CDS). For the match firms, the CDS onset year is assumed from their matched CDS firms.

**Table 3**
Summary statistics.

**Panel A: Pre-CDS trading period**

| | CDS | | Non-CDS | | Mean diff |
|---|---|---|---|---|---|
| Variable | Mean | Median | Mean | Median | |
| EPS | 0.033 | 0.055 | 0.023 | 0.050 | 0.010* |
| *R* | 0.178 | 0.153 | 0.186 | 0.153 | −0.008 |
| *D* | 0.273 | 0.000 | 0.293 | 0.000 | −0.021 |
| SIZE | 8.849 | 8.770 | 8.123 | 7.969 | 0.726*** |
| MB | 1.142 | 0.884 | 1.240 | 0.920 | −0.099** |
| LEV | 0.297 | 0.268 | 0.291 | 0.270 | 0.005 |

**Panel B: Post-CDS trading period**

| | CDS | | Non-CDS | | Mean diff |
|---|---|---|---|---|---|
| Variable | Mean | Median | Mean | Median | |
| EPS | −0.007 | 0.042 | −0.010 | 0.038 | 0.003 |
| *R* | 0.093 | 0.057 | 0.094 | 0.059 | −0.001 |
| *D* | 0.429 | 0.000 | 0.432 | 0.000 | −0.003 |
| SIZE | 8.424 | 8.307 | 7.627 | 7.480 | 0.798*** |
| MB | 1.112 | 0.786 | 1.333 | 0.817 | −0.221*** |
| LEV | 0.328 | 0.309 | 0.325 | 0.308 | 0.003 |

This table reports sample mean and median for main variables in the empirical analysis for both CDS firms and their matching firms (non-CDS) for both pre-CDS onset period and post-CDS onset period. The pre-CDS onset period covers two years prior to the onset of CDS and the post-CDS onset period covers two years after the onset of CDS. For non-CDS firms, the onset year is assumed from their matching firms. The sample period spans 2000–2011. EPS is net income scaled by prior year market value of equity; *R* is 12 month compounded returns starting 9 months before the fiscal year end. *D* is an indicator variable coded 1 if *R* is less than 0, and 0 otherwise. SIZE, natural logarithm of market value; MB, the ratio of market value of equity to book value of equity; LEV, leverage equal to total debt scaled by total assets.

**Table 4**
Pearson and Spearman correlations between selected variables.

| | Variables | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|---|
| (1) | EPS | 1.00 | **0.21** | **−0.19** | **0.32** | **0.13** | **−0.22** | 0.01 | **0.12** |
| (2) | *R* | **0.09** | 1.00 | **−0.73** | **0.08** | **0.08** | −0.03 | −0.01 | **0.11** |
| (3) | *D* | **−0.10** | **−0.83** | 1.00 | **−0.13** | **−0.07** | **0.06** | 0.00 | **−0.15** |
| (4) | SIZE | **0.16** | **0.09** | **−0.13** | 1.00 | **0.28** | **−0.33** | **0.23** | **0.13** |
| (5) | MB | 0.01 | **0.14** | **−0.15** | **0.38** | 1.00 | **−0.23** | **−0.06** | −0.01 |
| (6) | LEV | **−0.07** | −0.04 | **0.05** | **−0.31** | **−0.43** | 1.00 | 0.02 | **−0.08** |
| (7) | CDS | **0.04** | 0.00 | 0.00 | **0.24** | −0.04 | 0.03 | 1.00 | **−0.05** |
| (8) | POST | **0.16** | **0.13** | **−0.15** | **0.14** | **0.08** | **−0.09** | **−0.05** | 1.00 |

This table reports Pearson (above diagonal) and Spearman (below diagonal) corrections among variables used in the empirical analysis. The sample period spans 2000–2011. EPS is net income scaled by prior year market value of equity; *R* is 12 month compounded returns starting 9 months before the fiscal year end. *D* is an indicator variable coded 1 if *R* is less than 0, and 0 otherwise. SIZE is natural logarithm of market value of a firm; MB is market value of equity to book value of equity of a firm; LEV is leverage equal to total debt scaled by total assets; CDS is an indicator variable equal to 1 if a firm has a CDS traded over the sample period. POST is an indicator variable equal to 1 if a year falls in the two years after the onset of CDS trading, and 0 if a year falls in the two years before the onset of CDS trading for CDS firms. The match firms take on the same value of POST as the matched CDS firms in the pre- and post-CDS-trade-initiation year, respectively. Bold figures indicate significant level less than 1%.

CDS trade initiation over time. Second, the drastic decline may also foreshadow the looming financial crisis—by the end of 2008, CDS trades were initiated on only 3 new firms (0.6% of the CDS sample). By construction, we observe similar distribution for non-CDS firms. Table 2 Panel B reports the sample distribution by industry. As shown, CDS firms are primarily concentrated in the rubber, stone, computer, and transportation equipment industries (23.6% of the CDS sample). In addition, 23.1% of all CDS firms belong to the food, apparel, petroleum refining, and paper and printing industries, while 20.2% belong to the railroad transportation and electric and gas industries.

Table 3 reports descriptive statistics of variables used in subsequent analyses across the CDS sample and the matched non-CDS sample. The two samples exhibit characteristics that are generally similar in economic magnitudes. However, there are a few statistically significant differences between the two samples. During the two-year period prior to CDS trading, CDS firms have slightly better earnings performance. In addition, CDS firms are significantly larger in market capitalization and exhibit lower growth potential (lower MB).[16] The differences in firm size and growth between CDS firms and non-CDS firms are similar during the two-year period *after* the onset of CDS trading, although earnings performance becomes statistically similar between these two groups of firms. The similar magnitudes of differences between firm characteristics across CDS and non-CDS firms in Panels A and B suggest that these characteristics are unlikely to be driving the increasing difference in conservatism between CDS and non-CDS firms as one moves from Panel A to Panel B.

We report Pearson and Spearman correlations among variables used in our empirical tests in Table 4. As shown in Column (7), the correlations between CDS (an indicator variable equal to one for CDS firms, and zero for non-CDS firms) and earnings performance, firm size, growth, and leverage confirm the univariate patterns observed in Table 3.

## 4. Empirical methodology

### 4.1. *Measurement of accounting conservatism*

The aspect of conservatism we are interested in is the asymmetric timeliness of earnings in recognizing losses versus gains. We measure conservatism following Basu (1997), that is, the greater timeliness of earnings with respect to negative returns relative to positive returns. In particular, we estimate the following pooled cross-sectional model with standard errors clustered at the firm level (firm subscripts are suppressed for brevity):

$$EPS_t = \beta_0 + \beta_1 D_t + \beta_2 R_t + \beta_3 D_t \times R_t + \varepsilon_t \qquad (2)$$

where EPS is net income for fiscal year t scaled by year-beginning market value of equity; *R* is the 12-month compound stock returns ending three months after the end of fiscal year t; *D* is a indicator variable equal to one if *R* is negative, and zero otherwise.

In the above model, the sensitivity of earnings to economic gains is captured by $\beta_2$ and the sensitivity of earnings to economic losses is captured by $(\beta_2+\beta_3)$. If verification standards imposed for recognizing losses are lower than those imposed for recognizing gains, earnings will recognize economic losses in a timelier manner than economic gains. Hence, the association between earnings and stock returns should be incrementally higher when stock returns are negative, i.e., $\beta_3 > 0$ (Basu, 1997). We thus use $\beta_3$ to measure the extent to which earnings are reported conservatively. In the robustness tests, we also use Basu (1997)'s earnings time-series model to measure asymmetric timeliness of loss recognition.

[16] The probability-score-match controls only for the extent to which a combination of these firm characteristics (SIZE, MB, LEV) contribute to the probability of CDS trade initiation, rather than the individual characteristics.

Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

### 4.2. Research design

We use the difference-in-difference method in all empirical tests. Specifically, to examine the influence of CDS trade initiation on accounting conservatism, we expand Basu's (1997) baseline model as specified in Eq. (2) by including two indicator variables: the first identifies whether a firm has at least one CDS traded over the sample period and the second captures whether a firm-year observation falls in the two-year period after CDS trade initiation. We estimate the following model using ordinary least square regression:

$$\begin{aligned}\mathrm{EPS_t} = {} & \beta_0+\beta_1 R_t+\beta_2 D_t+\beta_3 D_t \times R_t+\beta_4\mathrm{CDS}+\beta_5\mathrm{CDS}\times R_t \\ & +\beta_6\mathrm{CDS}\times D_t+\beta_7\mathrm{CDS}\times D_t\times R_t+\beta_8\mathrm{POST}+\beta_9\mathrm{POST}\times R_t \\ & +\beta_{10}\mathrm{POST}\times D_t+\beta_{11}\mathrm{POST}\times D_t\times R_t+\beta_{12}\mathrm{CDS}\times\mathrm{POST} \\ & +\beta_{13}\mathrm{CDS}\times\mathrm{POST}\times R_t+\beta_{14}\mathrm{CDS}\times\mathrm{POST}\times D_t \\ & +\boldsymbol{\beta}_{15}\mathbf{CDS}\times\mathbf{POST}\times\mathbf{D_t}\times\mathbf{R_t}+\sum_{j=1}^{N}\lambda_j\mathrm{ADDITIONAL\ CONTROLS_j} \\ & +\sum_{i=1}^{K}\gamma_i\mathrm{INDUSTRY_i}+\sum_{m=1}^{N}\delta_m\mathrm{YEAR}_m+\varepsilon_t \end{aligned} \tag{3}$$

where CDS is an indicator variable equal to one for firms with a CDS traded during the sample period, and zero for matched control firms. POST is an indicator variable equal to one (zero) if an observation falls in the two-year period following (preceding) CDS trade initiation for both the CDS firm and its matched control firms. Since sample CDS trade initiations span 2002 to 2009, the examination of change in conservatism from two years before to two years after CDS trade initiation implies that the overall period of our analysis extends from 2000 to 2011. Industry and year fixed effects are included. Standard errors are clustered at the firm level to account for serial correlation within a firm (Petersen, 2009).[17] All the other variables are as defined in Eq. (2). To ensure that the results are not entirely driven by matched control firms, we also estimate Eq. (3) for CDS firms only (thereby dropping all terms relating to the indicator variable CDS).

Prior studies suggest that firm size, market-to-book ratio, and leverage may affect accounting conservatism (e.g., Basu, 1997; LaFond and Roychowdhury, 2008; LaFond and Watts, 2008). Therefore, we include these firm characteristics and their interactions with the three terms in Basu's (1997) model in Eq. (3) as additional control variables.

Our primary interest is the effect of CDS trade initiation on asymmetric timeliness of loss recognition (H1). Hence, we test whether the coefficient $\beta_{15}$, which captures the change in accounting conservatism of CDS firms relative to their matched firms, is significantly different from zero.

## 5. Empirical results

### 5.1. Primary tests

Table 5 reports regression results on the change in asymmetric loss recognition timeliness around the onset of CDS trading. The first two columns of Table 5 summarize results of estimating Eq. (3). As shown, the coefficient on $D \times R$ is significantly positive (coefficient = 1.216, $p$-value $< 0.001$), suggesting that non-CDS firms are more timely in recognizing economic losses than economic gains in the two-year period prior to CDS trading. The coefficient on CDS $\times D \times R$, which captures the difference in conservatism between CDS and non-CDS firms prior to the initiation of CDS trading, is positive and statistically significant (coefficient = 0.231, $p$-value = 0.015). Hence, prior to CDS trading, CDS firms exhibit higher levels of accounting conservatism than their matched non-CDS firms.

Comparing the pre-trading period with the post-trading period, non-CDS firms appear to have no change in the timeliness of recognizing economic losses in the two-year period after CDS trading, as evidenced by the statistically insignificant coefficient on POST $\times D \times R$ (coefficient = −0.012, $p$-value = 0.884). Importantly, we find a significantly negative coefficient on CDS $\times$ POST $\times D \times R$ (coefficient = −0.295, $p$-value = 0.036), suggesting that relative to matched control firms, CDS firms reduce asymmetric timeliness in loss recognition after the onset of CDS trading. The combined coefficient on CDS $\times$ POST $\times R$ and CDS $\times$ POST $\times D \times R$ (−0.254) is significantly negative, indicating that CDS firms also experience a significant decline in overall (and not just asymmetric) timeliness of loss recognition after CDS trade initiation compared to their match firms. These findings reject the null and support the notion that CDS firms experience a decline in accounting conservatism around the initial years of CDS trading. Economically, the incremental decline in asymmetric timeliness for CDS firms relative to non-CDS firms is about 20% of the accounting conservatism level of the CDS firms before the onset of CDS trading (= 0.295/(1.216 + 0.231)).[18]

Next, we estimate Eq. (3) for CDS firms only to ensure that our findings are not driven by the change in accounting conservatism for matched control firms. We therefore exclude all terms related with the indicator variable CDS and control

[17] Results are quantitatively similar if we cluster the standard errors at the year and two-digit SIC industry levels.

[18] This decline of 20.3% represents a net amount, inclusive of a 21.2% decline in the conservatism of CDS firms around CDS trade initiation. The matched non-CDS firms exhibit a (statistically insignificant) decline of 0.09% in their conservatism around the pseudo-CDS-trade-initiation dates assigned to them. For a comparison of magnitudes, Ettredge et al. (2012) findings imply a 46% increase in asymmetric timeliness of earnings following earnings restatements.

Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

**Table 5**
OLS regression results on the relation between asymmetric loss recognition timeliness and the onset of CDS trading.

| | Dependent variable = $EPS_t$ | | | |
|---|---|---|---|---|
| | CDS firms and matched firms | | CDS firms only | |
| Variable | Coeff. Est. | *p*-Value | Coeff. Est. | *p*-Value |
| $R_t$ ($\beta_1$) | −0.076 | 0.295 | −0.037 | 0.847 |
| $D_t$ ($\beta_2$) | 0.148 | 0.014 | 0.327 | 0.002 |
| $D_t \times R_t$ ($\beta_3$) | 1.216 | 0.000 | 1.589 | 0.000 |
| CDS ($\beta_4$) | −0.017 | 0.193 | | |
| CDS × $R_t$ ($\beta_5$) | −0.033 | 0.308 | | |
| CDS × $D_t$ ($\beta_6$) | 0.048 | 0.024 | | |
| CDS × $D_t \times R_t$ ($\beta_7$) | 0.231 | 0.015 | | |
| POST ($\beta_8$) | −0.023 | 0.023 | 0.009 | 0.572 |
| POST × $R_t$ ($\beta_9$) | 0.024 | 0.260 | 0.045 | 0.052 |
| POST × $D_t$ ($\beta_{10}$) | 0.038 | 0.045 | −0.029 | 0.144 |
| **POST × $D_t \times R_t$ ($\beta_{11}$)** | −0.012 | 0.884 | **−0.293** | **0.012** |
| CDS × POST ($\beta_{12}$) | 0.005 | 0.708 | | |
| CDS × POST × $R_t$ ($\beta_{13}$) | 0.041 | 0.208 | | |
| CDS × POST × $D_t$ ($\beta_{14}$) | −0.073 | 0.010 | | |
| CDS × POST × $D_t \times R_t$ ($\beta_{15}$) | **−0.295** | **0.036** | | |
| Additional controls | Included | | Included | |
| Intercept ($\beta_0$) | 0.087 | 0.000 | −0.047 | 0.625 |
| Year and industry fixed effects | Included | | Included | |
| *F*-test: ($\beta_{13}+\beta_{15}$) | −0.254 | 0.059 | | |
| *F*-test: ($\beta_9+\beta_{11}$) | | | −0.248 | 0.001 |
| Number of firm-years | 4,428 | | 1,996 | |
| Adjusted $R^2$ (%) | 34.51 | | 17.17 | |

This table reports multivariate regression results on the relation between Basu's (1997) measure of asymmetric loss recognition timeliness and the onset of CDS trading. The sample period spans 2000–2011. Firms in financial industries are excluded. The dependent variable is EPS, defined as net income scaled by prior year's market value of equity. *R* is twelve-month buy-and-hold returns starting nine months before the fiscal year end. *D* is an indicator variable equal to one if *R* is negative, and zero otherwise. CDS is an indicator variable equal to one if a firm has a CDS contract traded over the sample period, and zero for matched control firms. The control sample is chosen based on the matched-sample design, using the estimated probability of CDS trade initiation as described in Table 1. POST is an indicator variable equal to one if a year falls in the two-year period after the CDS-trade-initiation year, and zero if a year falls in the two-year period prior to the CDS-trade-initiation year for CDS firms. The match control firms take on the same value of POST as the matched CDS firms in the pre- and post-CDS-trade-initiation year, respectively. Additional controls include firm size, market-to-book ratio, book leverage, and their corresponding interaction terms with *R*, *D*, and *D* × *R*. Year and industry fixed effects are included. *P*-values are derived based on robust standard errors clustered at the firm level.

for several firm attributes that may affect accounting conservatism. The last two columns of Panel A report the results. We find that the coefficient on POST × *D* × *R* is significantly negative, supporting the notion that CDS firms experience a decline in accounting conservatism after the onset of CDS trading.

Taken together, results presented in Table 5 suggest that regardless of whether CDS firms are benchmarked with matched control firms or are used as their own controls, they become less asymmetric timely in reporting economic losses after the onset of CDS trading. Thus, CDS trade initiation has a net negative effect on accounting conservatism of borrowing firms.

## 5.2. *Cross-sectional tests*

### 5.2.1. *The role of reputation costs*

To test H2, we examine whether the change in accounting conservatism around CDS trading initiation varies with firm size, and credit rating. We partition our sample of treatment and control firm-years into two groups based on size and credit rating and estimate Eq. (3) within each subsample.

Table 6, Panel A presents results for two sub-samples partitioned based on size – specifically, firms below median market value of equity and those above. As shown, the coefficient on CDS × POST × *D* × *R* is significantly negative for firms with below median market value (coefficient = −0.518, *p*-value = 0.005), but insignificant for firms with above median market value (coefficient = 0.132, *p*-value = 0.544). An *F*-test of the statistical difference in this coefficient estimate across these two subsamples yields a *p*-value of 0.037.

Panel B presents results for two sub-samples partitioned based on S&P long-term credit rating – specifically, firms below investment-grade rating and those above. We find that the coefficient on CDS × POST × *D* × *R* is significantly negative for firms with below investment grade ratings (coefficient = −0.377, *p*-value = 0.023). We observe a smaller and insignificant

**Table 6**
Cross-sectional analysis of the relation between asymmetric loss recognition timeliness and the onset of CDS trading conditional on firm size and credit rating.

**Panel A: Conditional on firm size**

| | Dependent variable = $EPS_t$ | | | |
|---|---|---|---|---|
| | Below median MVE | | Above median MVE | |
| Variable | Coeff. Est. | *p*-Value | Coeff. Est. | *p*-Value |
| $R_t$ ($\beta_1$) | −0.011 | 0.941 | −0.050 | 0.329 |
| $D_t$ ($\beta_2$) | 0.121 | 0.456 | 0.068 | 0.235 |
| $D_t \times R_t$ ($\beta_3$) | 1.005 | 0.002 | 0.780 | 0.041 |
| CDS ($\beta_4$) | −0.051 | 0.039 | 0.006 | 0.407 |
| CDS × $R_t$ ($\beta_5$) | −0.027 | 0.606 | −0.036 | 0.037 |
| CDS × $D_t$ ($\beta_6$) | 0.085 | 0.018 | 0.007 | 0.710 |
| CDS × $D_t \times R_t$ ($\beta 7$) | 0.289 | 0.026 | 0.098 | 0.351 |
| POST ($\beta_8$) | −0.038 | 0.018 | −0.001 | 0.865 |
| POST × $R_t$ ($\beta_9$) | 0.052 | 0.106 | −0.005 | 0.783 |
| POST × $D_t$ ($\beta_{10}$) | 0.058 | 0.032 | −0.005 | 0.801 |
| POST × $D_t \times R_t$ ($\beta_{11}$) | −0.025 | 0.813 | −0.117 | 0.271 |
| CDS × POST ($\beta_{12}$) | 0.022 | 0.391 | −0.010 | 0.301 |
| CDS × POST × $R_t$ ($\beta_{13}$) | 0.051 | 0.352 | 0.026 | 0.232 |
| CDS × POST × $D_t$ ($\beta_{14}$) | −0.147 | 0.002 | 0.010 | 0.725 |
| **CDS × POST × $D \times R_t$ ($\beta_{15}$)** | **−0.518** | **0.005** | **0.132** | **0.544** |
| Additional controls | Included | | Included | |
| Intercept ($\beta_0$) | −0.279 | 0.005 | 0.048 | 0.079 |
| Year and industry fixed effects | Included | | Included | |
| *F*-test: ($\beta_{13}+\beta_{15}$) | −0.467 | 0.0593 | 0.158 | 0.462 |
| *F*-test: $\beta_{15}$ across subsamples (*p*-value) | | −0.65 (0.037) | | |
| Number of firm-years | 2,214 | | 2,214 | |
| Adjusted $R^2$ (%) | 34.49 | | 15.04 | |

**Panel B: Conditional on credit rating**

| | Dependent variable = $EPS_t$ | | | |
|---|---|---|---|---|
| | Below investment grade | | Above investment grade | |
| Variable | Coeff. Est. | *p*-Value | Coeff. Est. | *p*-Value |
| $R_t$ ($\beta_1$) | −0.025 | 0.783 | −0.007 | 0.863 |
| $D_t$ ($\beta_2$) | 0.167 | 0.078 | 0.005 | 0.878 |
| $D_t \times R_t$ ($\beta_3$) | 1.131 | 0.000 | 0.064 | 0.750 |
| CDS ($\beta_4$) | −0.030 | 0.133 | −0.004 | 0.537 |
| CDS × $R_t$ ($\beta_5$) | −0.030 | 0.457 | −0.014 | 0.474 |
| CDS × $D_t$ ($\beta_6$) | 0.072 | 0.025 | −0.010 | 0.511 |
| CDS × $D_t \times R_t$ ($\beta_7$) | 0.276 | 0.016 | 0.026 | 0.765 |
| POST ($\beta_8$) | −0.034 | 0.009 | 0.004 | 0.664 |
| POST × $R_t$ ($\beta_9$) | 0.033 | 0.194 | 0.003 | 0.852 |
| POST × $D_t$ ($\beta_{10}$) | 0.058 | 0.019 | −0.006 | 0.743 |
| POST × $D_t \times R_t$ ($\beta_{11}$) | 0.029 | 0.754 | −0.094 | 0.532 |
| CDS × POST ($\beta_{12}$) | 0.008 | 0.682 | 0.004 | 0.542 |
| CDS × POST × $R_t$ ($\beta_{13}$) | 0.047 | 0.252 | 0.004 | 0.838 |
| CDS × POST × $D_t$ ($\beta_{14}$) | −0.108 | 0.006 | −0.004 | 0.862 |
| **CDS × POST × $D_t \times R_t$ ($\beta_{15}$)** | **−0.377** | **0.023** | **−0.089** | **0.603** |
| Additional controls | Included | | Included | |
| Intercept ($\beta_0$) | −0.109 | 0.050 | 0.057 | 0.027 |
| Year and industry fixed effects | Included | | Included | |
| *F*-test: ($\beta_{13}+\beta_{15}$) | −0.330 | 0.071 | −0.085 | 0.614 |
| *F*-test: $\beta_{15}$ across subsamples (*p*-value) | | −0.288 (0.075) | | |
| Number of firm-years | 3,000 | | 1,428 | |
| Adjusted $R^2$ (%) | 30.55 | | 11.93 | |

**Table 6** (*continued*)

**Panel C: Conditional on lender reputation factor**

| | Dependent variable = $EPS_t$ | | | |
|---|---|---|---|---|
| | Below median reputation factor | | Above median reputation factor | |
| Variable | Coeff. Est. | *p*-Value | Coeff. Est. | *p*-Value |
| $R_t(\beta 1)$ | −0.028 | 0.776 | 0.046 | 0.552 |
| $D_t(\beta 2)$ | 0.161 | 0.062 | 0.089 | 0.032 |
| $D_t \times R_t(\beta 3)$ | 1.241 | 0.000 | 0.553 | 0.000 |
| $CDS(\beta 4)$ | −0.036 | 0.104 | 0.002 | 0.856 |
| $CDS \times R_t(\beta 5)$ | −0.042 | 0.312 | −0.001 | 0.973 |
| $CDS \times D_t(\beta 6)$ | 0.071 | 0.035 | −0.002 | 0.868 |
| $CDS \times D_t \times R_t(\beta 7)$ | 0.342 | 0.000 | 0.030 | 0.484 |
| $POST(\beta 8)$ | −0.034 | 0.126 | 0.020 | 0.064 |
| $POST \times R_t(\beta 9)$ | 0.067 | 0.053 | 0.000 | 0.993 |
| $POST \times D_t(\beta 10)$ | 0.069 | 0.022 | −0.014 | 0.371 |
| $POST \times D_t \times R_t(\beta 11)$ | 0.024 | 0.768 | −0.094 | 0.209 |
| $CDS \times POST(\beta 12)$ | 0.015 | 0.626 | −0.012 | 0.271 |
| $CDS \times POST \times R_t(\beta 13)$ | 0.038 | 0.512 | 0.010 | 0.729 |
| $CDS \times POST \times D_t(\beta 14)$ | −0.146 | 0.003 | 0.034 | 0.080 |
| **$CDS \times POST \times D_t \times R_t(\beta 15)$** | **−0.585** | **0.005** | **0.177** | **0.478** |
| Additional controls | Included | | Included | |
| Intercept ($\beta_0$) | −0.276 | 0.001 | −0.068 | 0.107 |
| Year and industry fixed effects | Included | | Included | |
| *F*-test: $(\beta_{13}+\beta_{15})$ | −0.547 | 0.005 | 0.187 | 0.463 |
| *F*-test: $\beta_{15}$ across subsamples (*p*-value) | | −0.762 (0.014) | | |
| Number of firm-years | 2214 | | 2214 | |
| Adjusted $R^2$ (%) | 33.2 | | 19.9 | |

This table reports cross-sectional analysis of firm size, credit rating, and the common factor constructed based on the former two variables on the relation between asymmetric timely loss recognition and the onset of CDS trading. The sample period spans 2000–2011. Firms in financial industries are excluded. The dependent variable is EPS, defined as net income scaled by prior year's market value of equity. In Panel A, the sample is partitioned into large firms and small firms based on the full sample median market value of equity prior to the year of CDS onset. In Panel B, the sample is partitioned into firms with investment-grade credit rating and firms without investment-grade credit rating prior to the year of CDS onset. In Panel C, the sample is partitioned into high lender reputation and low lender reputation based on the full sample median lender reputation factor. Lender reputation factor is derived from the principal component analysis based on the two variables: natural logarithm of firm market value of equity and long-term S&P credit rating. Credit rating is defined by an ordinal variable ranging between 1 (AAA) and 19 (CCC-) for firms with S&P long term debt rating; we assign a value of 20 for firms in default stage, and 21 for firms with no debt rating. *R* is twelve-month buy-and-hold returns starting nine months before the fiscal year end. *D* is an indicator variable equal to one if *R* is negative, and zero otherwise. CDS is an indicator variable equal to one if a firm has a CDS contract traded over the sample period, and zero for matched control firms. The control sample is chosen based on the matched-sample design, using the estimated probability of CDS trade initiation as described in Table 1. POST is an indicator variable equal to one if a year falls in the two-year period after the CDS-trade-initiation year, and zero if a year falls in the two-year period prior to the CDS-trade-initiation year for CDS firms. The match control firms take on the same value of POST as the matched CDS firms in the pre- and post-CDS-trade-initiation year, respectively. Additional controls include firm size, market-to-book ratio, book leverage, and their corresponding interaction terms with *R*, *D*, and $D \times R$. Year and industry fixed effects are included. *P*-values are derived based on robust standard errors clustered at the firm level.

decline in conservatism for firms with credit rating above investment grade (coefficient = −0.089, *p*-value = 0.603). The difference in the decline in conservatism between the two groups is significant at the 10% level.

In Panel C we construct a common factor, Reputation Factor, based on both firm size and credit rating using principle-components. Size is defined as natural logarithm of market value of equity. Credit rating is defined by an ordinal variable ranging between 1 (AAA) and 19 (CCC-) for firms with S&P long term debt rating; we assign a value of 20 for firms in default stage, and 21 for firms with no debt rating. The common factor thus captures the contribution of both size and credit rating to ex ante risk of borrower default, which is expected to be associated negatively with lenders' incentives to maintain post-CDS monitoring out of concern for reputation costs. As constructed, Reputation Factor varies positively with size and negatively with credit rating. We partition the sample based on whether the value of Reputation Factor for a particular firm-year is above or below the median value for that year. The results show that the coefficient on $CDS \times POST \times D \times R$ is significantly negative for firms with low Reputation Factor (coefficient = −0.585, *p*-value < 0.005), but insignificant for firms with high Reputation Factor (coefficient = 0.177, *p*-value = 0.478). An *F*-test of the statistical difference in this coefficient estimate across these two subsamples yields a *p*-value of 0.014. The decline in conservatism for CDS firms when their lenders are likely to bear lower reputation costs is about 35% of their conservatism level in the pre-CDS period $(=(-0.585+0.024)/(0.342+1.241))$.

The results in Panels A, B and C collectively indicate that the post-CDS decline in borrower conservatism is more pronounced when banks entering into CDS contracts face lower reputation costs (as in borrowers with smaller size and

**Table 7**
Cross-sectional analysis of the relation between asymmetric loss recognition timeliness and the onset of CDS trading conditional on loan contracts with financial covenants.

| | Dependent variable = $EPS_t$ | | | |
|---|---|---|---|---|
| | Number of financial covenants > 3 | | Number of financial covenants < 2 | |
| Variable | Coeff. Est. | *p*-Value | Coeff. Est. | *p*-Value |
| $R_t(\beta 1)$ | −0.043 | 0.727 | −0.173 | 0.415 |
| $D_t(\beta 2)$ | 0.117 | 0.277 | 0.080 | 0.597 |
| $D_t \times R_t(\beta 3)$ | 1.002 | 0.001 | 0.838 | 0.027 |
| CDS($\beta 4$) | −0.010 | 0.670 | −0.013 | 0.349 |
| CDS × $R_t(\beta 5)$ | −0.081 | 0.034 | 0.066 | 0.301 |
| CDS × $D_t(\beta 6)$ | 0.024 | 0.605 | −0.004 | 0.913 |
| CDS × $D_t \times R_t(\beta 7)$ | 0.231 | 0.243 | −0.191 | 0.449 |
| POST($\beta 8$) | −0.008 | 0.678 | −0.030 | 0.228 |
| POST × $R_t(\beta 9)$ | 0.029 | 0.472 | 0.086 | 0.182 |
| POST × $D_t(\beta 10)$ | 0.028 | 0.387 | −0.038 | 0.434 |
| POST × $D_t \times R_t(\beta 11)$ | −0.053 | 0.738 | −0.496 | 0.005 |
| CDS × POST($\beta 12$) | −0.024 | 0.103 | 0.033 | 0.164 |
| CDS × POST × $R_t(\beta 13)$ | 0.123 | 0.031 | −0.064 | 0.402 |
| CDS × POST × $D_t(\beta 14)$ | −0.029 | 0.534 | −0.038 | 0.468 |
| **CDS × POST × $D_t \times R_t(\beta 15)$** | **−0.497** | **0.006** | **0.246** | **0.397** |
| Additional Controls | Included | | Included | |
| Intercept ($\beta_0$) | −0.037 | 0.580 | −0.065 | 0.251 |
| Year and industry fixed effects | Included | | Included | |
| *F*-test: ($\beta_{13}+\beta_{15}$) | −0.374 | 0.024 | 0.182 | 0.274 |
| *F*-test: $\beta_{15}$ across subsamples (*p*-value) | | −0.556 (0.067) | | |
| Number of firm-years | 1153 | | 711 | |
| Adjusted $R^2$ (%) | 37.01 | | 33.04 | |

This table reports cross-sectional analysis based on financial covenants in loan contracts to the firms in our sample. The sample period spans 2000 to 2011. Banks lending to CDS and non-CDS firms in the sample are identified using data obtained from the LPC (Loan Pricing Corporation)'s Dealscan database. Loan contracts that are outstanding at least two years prior to the CDS trade initiation date but mature at least two years after that date are identified from the LPC database. Among all loans outstanding, the number of financial covenants is measured for the loan with the maximum number of financial covenants in the year prior to the onset of CDS trading. The sample is partitioned into more (few) covenants groups in which firms have loans contracts outstanding with the number of financial covenants exceeding 3 (below 2). The dependent variable is EPS, defined as net income scaled by prior year's market value of equity. *R* is twelve-month buy-and-hold returns starting nine months before the fiscal year end. *D* is an indicator variable equal to one if *R* is negative, and zero otherwise. CDS is an indicator variable equal to one if a firm has a CDS contract traded over the sample period, and zero for matched control firms. POST is an indicator variable equal to one if a year falls in the two-year period after the CDS-trade-initiation year, and zero if a year falls in the two-year period prior to the CDS-trade-initiation year for CDS firms. The match control firms take on the same value of POST as the matched CDS firms in the pre- and post-CDS-trade-initiation year, respectively. Additional controls include firm size, market-to-book ratio, book leverage, and their corresponding interaction terms with *R*, *D*, and $D \times R$. Year and industry fixed effects are included. *p*-Values are derived based on robust standard errors clustered at the firm level.

poorer credit ratings) from reducing monitoring. We do not observe any evidence of a decline in conservatism among larger firms with higher credit ratings, suggesting either lenders do not reduce their demand for conservatism among such firms or borrowers maintain their consistency with respect to conservative accounting practices (presumably to satisfy the demands of other stakeholders).

### *5.2.2. The role of covenants*

In examining the role of covenants (H3), we first identify the number of financial covenants in firms' private debt contracts from the LPC (Loan Pricing Corporation)'s Dealscan database. Due to data availability requirements, we are left with a smaller sample size of 3,074 firm-years containing 417 unique CDS firms and 317 unique matched non-CDS firms. We next identify all loan contracts outstanding in the fiscal year *prior* to CDS trade initiation and maturing *after* CDS trade initiation. Among these loan contracts, we consider the one with the maximum number of financial covenants because this number likely represents the binding covenant intensity. Subsequently we partition the sample based on whether the number of loan covenants exceeds the sample top quartile (i.e., number of covenants > 3) or whether that number is below the bottom quartile (i.e., number of covenants < 2). Table 7 reports results for this partition. Firms exhibit a much more prominent decline in conservatism when the number of financial covenants in existing loan contracts at the time of the CDS trade initiation is above the 75th percentile. From an economic perspective, the accounting conservatism of CDS firms declines about 45% relative to their pre-CDS trade initiation period. When the number of financial covenants in existing loan contracts at the time of the CDS is below the 25th percentile, firms do not exhibit a significant decline in conservatism.

Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

### 5.3. The issue of lender identity

We first identify banks with outstanding loans to the CDS firms and the matched control firms in our sample. The partial effect for a bank of investing in a CDS contract on a borrower would be to lower the risk weight assigned to the loan on that borrower and to increase the bank's CDS holdings.[19] However, it is difficult to observe the effect of a CDS contract on a single borrower on the bank's risk-weighted assets or on its CDS portfolio. Therefore, we employ a reverse approach. For each bank, we can observe whether there is a change in the proportion of total assets bearing a risk weight lower than 100% in a given year, and also whether their overall CDS holdings increases/decreases in any given year. We reason that banks for whom the proportion of assets weighted at lower than 100% rises, or banks that exhibit an increase in overall CDS holdings, *in the same year as* CDS trade initiation on an underlying borrower are more likely to have hedged their exposure to the specific borrower via the newly available CDS contracts.

Table 8 presents results of testing H4. In Panel A we present results for two sub-samples partitioned based on whether there was an increase in the proportion of banks' assets risk-weighted at lower than 100% in the same year as CDS trade initiation. We identify banks lending to CDS and non-CDS firms in our sample using data obtained from the LPC (Loan Pricing Corporation) Dealscan database, and the risk weights on banks' assets from Federal Reserve's Y-9C reports. We find that the coefficient on CDS × POST × $D$ × $R$ is significantly negative for firms whose banks exhibit an increase in the proportion of assets that bear risk weights lower than 100% (coefficient = −0.452, $p$-value = 0.003), but is actually positive and statistically significant for firms whose banks do not exhibit an increase in the proportion of assets that bear risk weights lower than 100% (coefficient = 0.286, $p$-value = 0.028). An $F$-test of the statistical difference in this coefficient estimate across these two subsamples yields a $p$-value of 0.003.

Table 8, Panel B presents results for two sub-samples partitioned on whether banks exhibit an increase in CDS portfolio holdings in the same year as CDS trade initiation on underlying borrowers. CDS portfolio holdings of banks are obtained from Federal Reserve's Y-9C reports. As shown, the coefficient on CDS × POST × $D$ × $R$ is significantly negative for firms whose banks exhibit an increase in CDS holdings in the year of CDS trade initiation (coefficient = −1.020, $p$-value = 0.000), but insignificant for firms whose banks do not exhibit an increase in CDS holdings (coefficient = −0.113, $p$-value = 0.310). An $F$-test of the statistical difference in this coefficient estimate across these two subsamples yields a $p$-value of 0.100.

Finally, Panel C presents results based on identifying firms whose banks exhibit either an increase in the proportion of assets bearing risk weights below 100% or an increase in CDS holdings. We find that the coefficient on CDS × POST × $D$ × $R$ is significantly negative when either condition is satisfied (coefficient = −0.522, $p$-value = 0.000). Economically, the decline in conservatism for CDS firms when their lenders are likely to enter into CDS contracts for hedging is about 27% of their conservatism level in the pre-CDS period (=(-0.522 + 0.022)/(0.245 + 1.584)). In contrast, the conservatism level for CDS firms relative to their counterparts when lenders unlikely enter into CDS contracts actually increases significantly in the post-CDS period relative to matched control firms (coefficient = 0.437, $p$-value = 0.003). This may be evidence of a selection bias: banks that do not hedge their exposures to underlying borrowers even when CDS contracts are available are likely to be the ones that intend to increase their monitoring of borrowers' financial statements. An $F$-test of the statistical difference in this coefficient estimate across the two subsamples yields a $p$-value below 0.000 (rounded).[20]

We conduct a variety of tests to check the robustness of our results. First, we modify the first-stage prediction model following Minton et al. (2009), to incorporate bank propensity to invest in derivative securities, including interest rate, foreign exchange, equity and commodity derivatives. The findings we obtain in Tables 5–7 with respect to both our primary and cross-sectional hypotheses are robust to using this first stage model. Second, we identify as matched control firms the *two* rather than the three nearest neighbors (with replacement) within 20% of the CDS-trade-initiation probability score for each treatment firm. We also modify the thresholds to include control firms within 5% rather than 20% of the treatment firm's CDS-trade-initiation probability score. Finally, we use a *one year* window rather than a two-year window for the pre- and post-CDS-trade-initiation analysis. The results (untabulated) are robust to all these alternative procedures.

### 5.4. Additional analysis

#### 5.4.1. Endogeneity between expected change in conservatism and CDS trade initiation

In our final analysis, we examine the possibility that CDS trade initiation is more likely when lenders anticipate a decline in borrower conservatism. While existing literature does not raise this possibility, it has a testable empirical prediction: a negative association between *expected* change in conservatism and CDS trade initiation. Note that in our primary tests, we

[19] Loans to corporate entities are assigned a risk weight between 20% and 150% under the standardized approach to credit risk. The risk weight declines if the bank is hedged on its exposure to a specific borrower via CDS contracts if the credit rating of the CDS sellers is higher than that of the borrower. According to Basel II, a risk weight of 150% is assigned to loans rated below BB-, 100% to loans rated above BB- but below AA-, and 20% for loans with CDS protection where CDS sellers are rating above AA-.

[20] We repeat all our cross-sectional analyses in Section 5.2 within the sample of firms that are likely to have hedged their credit exposures via CDSs (that is, firms whose banks exhibit either an increase in the proportion of assets bearing risk weights below 100% or an increase in CDS holdings). The results on cross-sectional variation obtained in Tables 6 and 7 are robust to using this sub-sample; that is, we find a more pronounced decline in conservatism (a) when borrowers are risky and informationally opaque ex ante, and (b) when borrowers' debt contracts outstanding at the time of CDS trade initiation include a larger number of financial covenants.

**Table 8**
Cross-sectional analysis of the relation between asymmetric loss recognition timeliness and the onset of CDS trading conditional on bank characteristics.

**Panel A: The proportion of banks' assets bearing lower risk weights**

| | Dependent variable = $EPS_t$ | | | |
|---|---|---|---|---|
| | Increase in proportion of assets with lower risk-weight | | No increase in proportion of assets with lower risk-weight | |
| Variable | Coef. | $P>t$ | Coef. | $P>t$ |
| $R_t(\beta 1)$ | −0.190 | 0.121 | 0.015 | 0.880 |
| $D_t(\beta 2)$ | 0.279 | 0.004 | 0.167 | 0.018 |
| $D_t \times R_t(\beta 3)$ | 1.734 | 0.000 | 1.010 | 0.000 |
| CDS($\beta$4) | 0.004 | 0.855 | −0.047 | 0.010 |
| CDS × $R_t(\beta 5)$ | −0.091 | 0.054 | 0.086 | 0.035 |
| CDS × $D_t(\beta 6)$ | −0.003 | 0.928 | 0.042 | 0.138 |
| CDS × $D_t \times R_t(\beta 7)$ | 0.173 | 0.121 | −0.147 | 0.069 |
| POST($\beta$8) | 0.003 | 0.914 | −0.005 | 0.802 |
| POST × $R_t(\beta 9)$ | 0.011 | 0.804 | 0.045 | 0.254 |
| POST × $D_t(\beta 10)$ | 0.001 | 0.980 | −0.021 | 0.514 |
| POST × $D_t \times R_t(\beta 11)$ | −0.101 | 0.378 | −0.345 | 0.000 |
| CDS × POST($\beta$12) | −0.027 | 0.368 | 0.032 | 0.179 |
| CDS × POST × $R_t(\beta 13)$ | 0.099 | 0.129 | −0.041 | 0.453 |
| CDS × POST × $D_t(\beta 14)$ | −0.042 | 0.422 | 0.000 | 0.998 |
| **CDS × POST × $D_t \times R_t(\beta 15)$** | **−0.452** | **0.003** | **0.286** | **0.028** |
| Additional controls | Included | | Included | |
| Intercept ($\beta_0$) | 0.016 | 0.825 | −0.069 | 0.192 |
| Year and industry fixed effects | Included | | Included | |
| *F*-test: ($\beta_{13}+\beta_{15}$) | −0.353 | 0.010 | 0.245 | 0.038 |
| *F*-test: $\beta_{15}$ across subsamples (*p*-value) | | −0.598 (0.003) | | |
| Number of firm-years | 1,135 | | 1,651 | |
| Adjusted $R^2$ (%) | 33.13 | 35.01 | | |

**Panel B: Bank CDS holding change**

| | Dependent variable = $EPS_t$ | | | |
|---|---|---|---|---|
| | Increase in CDS holdings | | No increase in CDS holdings | |
| Variable | Coef. | $P>t$ | Coef. | $P>t$ |
| $R_t(\beta 1)$ | −0.361 | 0.014 | 0.001 | 0.991 |
| $D_t(\beta 2)$ | 0.115 | 0.357 | 0.226 | 0.000 |
| $D_t \times R_t(\beta 3)$ | 1.633 | 0.000 | 1.151 | 0.000 |
| CDS($\beta$4) | 0.023 | 0.461 | −0.025 | 0.131 |
| CDS × $R_t(\beta 5)$ | −0.248 | 0.001 | 0.036 | 0.312 |
| CDS × $D_t(\beta 6)$ | 0.074 | 0.148 | 0.017 | 0.524 |
| CDS × $D_t \times R_t(\beta 7)$ | 0.742 | 0.000 | −0.067 | 0.350 |
| POST($\beta$8) | 0.003 | 0.903 | −0.002 | 0.931 |
| POST × $R_t(\beta 9)$ | −0.046 | 0.370 | 0.064 | 0.075 |
| POST × $D_t(\beta 10)$ | 0.027 | 0.522 | −0.002 | 0.946 |
| POST × $D_t \times R_t(\beta 11)$ | 0.076 | 0.554 | −0.173 | 0.045 |
| CDS × POST($\beta$12) | −0.099 | 0.052 | 0.013 | 0.541 |
| CDS × POST × $R_t(\beta 13)$ | 0.354 | 0.001 | −0.019 | 0.685 |
| CDS × POST × $D_t(\beta 14)$ | −0.118 | 0.158 | −0.028 | 0.453 |
| **CDS × POST × $D_t \times R_t(\beta 15)$** | **−1.020** | **0.000** | **−0.113** | **0.310** |
| Additional controls | Included | | Included | |
| Intercept ($\beta_0$) | 0.142 | 0.160 | −0.189 | 0.002 |
| Year and Industry fixed effects | Included | | Included | |
| *F*-test: ($\beta_{13}+\beta_{15}$) | −0.665 | 0.002 | −0.133 | 0.188 |
| *F*-test: $\beta_{15}$ across subsamples (*p*-value) | | −0.906 (0.101) | | |
| Number of firm-years | 435 | | 2,351 | |
| Adjusted $R^2$ (%) | 37.02 | | 33.01 | |

**Panel C: Banks with an increase in either the proportion of assets bearing lower risk weights or in CDS holdings**

| | Dependent variable = $EPS_t$ | | | |
|---|---|---|---|---|
| | Increase in either the proportion of assets bearing lower risk weights or in CDS holdings | | No increase in either | |
| Variable | Coef. | $P>t$ | Coef. | $P>t$ |
| $R_t(\beta 1)$ | 0.112 | −0.393 | 0.026 | 0.803 |
| $D_t(\beta 2)$ | 0.252 | 0.004 | 0.194 | 0.010 |

**Table 8** (*continued* )

| | | | | |
|---|---|---|---|---|
| $D_t \times R_t(\beta 3)$ | 1.584 | 0.000 | 1.104 | 0.000 |
| CDS($\beta$4) | 0.002 | 0.899 | −0.047 | 0.021 |
| CDS × $R_t(\beta 5)$ | −0.086 | 0.046 | 0.095 | 0.036 |
| CDS × $D_t(\beta 6)$ | 0.009 | 0.793 | 0.037 | 0.241 |
| CDS × $D_t \times R_t(\beta 7)$ | 0.245 | 0.014 | −0.240 | 0.007 |
| POST($\beta$8) | 0.004 | 0.856 | 0.000 | 1.000 |
| POST × $R_t(\beta 9)$ | −0.005 | 0.895 | 0.073 | 0.102 |
| POST × $D_t(\beta 10)$ | 0.005 | 0.880 | −0.028 | 0.461 |
| POST × $D_t \times R_t(\beta 11)$ | 0.022 | 0.816 | −0.482 | 0.000 |
| CDS × POST($\beta$12) | −0.030 | 0.265 | 0.039 | 0.138 |
| CDS × POST × $R(\beta 13)$ | 0.118 | 0.048 | −0.070 | 0.242 |
| CDS × POST × $D_t(\beta 14)$ | −0.047 | 0.320 | 0.012 | 0.795 |
| **CDS × POST × $D_t \times R_t(\beta 15)$** | **−0.522** | **0.000** | **0.437** | **0.003** |
| Additional controls | Included | | Included | |
| Intercept ($\beta_0$) | −0.127 | 0.086 | −0.082 | 0.315 |
| Year and industry fixed effects | Included | | Included | |
| *F*-test: ($\beta_{13}+\beta_{15}$) | −0.404 | 0.001 | 0.367 | 0.005 |
| *F*-test: $\beta_{15}$ across subsamples (*p*-value) | | −0.960 (0.000) | | |
| Number of firm-years | 1,296 | | 1,490 | |
| Adjusted $R^2$ (%) | 0.311 | | 0.369 | |

This table reports cross-sectional analysis of characteristics of banks lending to the firms in our sample. The sample period spans 2000–2011. Firms in financial industries are excluded. The dependent variable is EPS, defined as net income scaled by prior year's market value of equity. Panel A presents results for two sub-samples partitioned on whether there was an increase in the proportion of banks' assets risk-weighted at lower than 100% relative to total bank assets in the same year as CDS trade initiation. Banks lending to CDS and non-CDS firms in the sample are identified using data obtained from the LPC (Loan Pricing Corporation)'s Dealscan database, and the risk weights on banks' assets are from Federal Reserve's Y-9C reports. Panel B presents results for two sub-samples partitioned based on whether banks exhibit an increase in CDS portfolio holdings in the same year as CDS trade initiation on underlying borrowers. CDS portfolio holdings of banks are obtained from Federal Reserve's Y-9C reports. Panel C presents for situations in which either the conditions in Panel A or Panel B hold. In other words, banks exhibit an increase in either lower-risk-weighted assets or an increase in CDS portfolio holdings. *R* is twelve-month buy-and-hold returns starting nine months before the fiscal year end. *D* is an indicator variable equal to one if *R* is negative, and zero otherwise. CDS is an indicator variable equal to one if a firm has a CDS contract traded over the sample period, and zero for matched control firms. The control sample is chosen based on the matched-sample design, using the estimated probability of CDS trade initiation as described in Table 1. POST is an indicator variable equal to one if a year falls in the two-year period after the CDS-trade-initiation year, and zero if a year falls in the two-year period prior to the CDS-trade-initiation year for CDS firms. The matched control firms take on the same value of POST as the CDS firms in the pre- and post-CDS-trade-initiation year, respectively. Additional controls include firm size, market-to-book ratio, book leverage, and their corresponding interaction terms with *R*, *D*, and $D \times R$. Year and industry fixed effects are included. *p*-Values are derived based on robust standard errors clustered at the firm level.

are interested in whether there is an actual *ex post* decline in conservatism after CDS trade initiation as lenders lower monitoring of financial statements. Key to distinguishing between the two possibilities is the measurement of, and the imposition of a control for, expected change in conservatism. Accordingly, we modify our empirical research design to match control firms to treatment firms based on expected change in conservatism.

To accommodate the cross-sectional nature of our conservatism measure, we develop a novel approach for measuring expected change in conservatism. Khan and Watts (2009) demonstrate that the asymmetric timeliness of earnings (i.e., the Basu measure) varies monotonically across deciles of CSCORE, a firm-specific measure of conservatism. We partition firms in the Compustat universe into quintiles of CSCORE in the year prior to CDS trade initiation, that is, year $t-1$ (where year $t$ is the year of CDS trade initiation). We then estimate the cross-sectional Basu measure within each of these quintiles and assign the corresponding asymmetric timeliness coefficient to all firms in that quintile. Holding quintile membership constant, we measure the cross-sectional Basu measure for year $t+1$, that is, the year after CDS trade initiation. The actual change in the asymmetric timeliness coefficient for the CSCORE quintile from year $t-1$ to $t+1$ serves as a proxy for the expected change in conservatism for every firm within that quintile. Thereafter, we augment our first stage model with expected change in conservatism.

The association between CDS trade initiation and expected change in conservatism is significantly negative. Note that our measure of expected conservatism suffers from hindsight bias, since it relies on actual change in conservatism measured *ex post*. Therefore, caution is warranted in interpreting the negative association observed in the first stage as a causal relation between expected conservatism change and CDS trade initiation. This potentially reduces the power of our second-stage regression to detect an actual change in conservatism for CDS firms. Upon matching on the CDS trading probability score from the modified first stage model, we detect no difference in *expected* change in conservatism between matched and control firms. In the second stage, we still observe an actual decline in conservatism among CDS firms upon CDS trade initiation, with this decline being much more pronounced relative to the matched non-CDS firms identified earlier, as shown in Table 9.[21]

[21] The sample size reduces slightly due to the enhanced data requirements for estimating our first stage model, which adversely influences the extent to which we can find matches for CDS firms.

**Table 9**
Robustness analysis – change in conservatism included in the first stage model.

| Variable | Dependent variable = $EPS_t$ | |
|---|---|---|
| | Coeff. Est. | *p*-Value |
| $R_t$ ($\beta_1$) | −0.028 | 0.767 |
| $D_t$ ($\beta_2$) | 0.182 | 0.007 |
| $D_t \times R_t$ ($\beta_3$) | 1.105 | 0.000 |
| CDS ($\beta_4$) | −0.035 | 0.044 |
| CDS × $R_t$ ($\beta_5$) | −0.017 | 0.666 |
| CDS × $D_t$ ($\beta_6$) | 0.036 | 0.157 |
| CDS × $D_t \times R_t$ ($\beta_7$) | 0.154 | 0.185 |
| POST ($\beta_8$) | −0.006 | 0.534 |
| POST × $R_t$ ($\beta_9$) | 0.011 | 0.626 |
| POST × $D_t$ ($\beta_{10}$) | −0.007 | 0.761 |
| POST × $D_t \times R_t$ ($\beta_{11}$) | −0.110 | 0.302 |
| CDS × POST ($\beta_{12}$) | 0.005 | 0.744 |
| CDS × POST × $R_t$ ($\beta_{13}$) | 0.048 | 0.177 |
| CDS × POST × $D_t$ ($\beta_{14}$) | −0.022 | 0.487 |
| **CDS × POST × $D_t \times R_t$ ($\beta_{15}$)** | **−0.256** | **0.085** |
| Additional controls | Included | |
| Intercept ($\beta_0$) | −0.088 | 0.143 |
| Year and industry fixed effects | Included | |
| *F*-test: ($\beta_{13}+\beta_{15}$) | −0.208 | 0.097 |
| Number of firm-years | 4,002 | |
| Adjusted $R^2$ (%) | 29.05 | |

This table reports regression results on the relation between Basu's (1997) measure of asymmetric loss recognition timeliness and the onset of CDS trading based on a first stage model that includes expected change in asymmetric loss recognition timeliness from one year before to one year after the onset of CDS trading. For computing expected change in asymmetric timeliness, firms in Compustat are sorted into quintiles each fiscal year based on CSCORE, which in turn is computed following Khan and Watts (2009). For every firm-year, the change in the Basu coefficient from year −1 to year +1 for the CSCORE quintile the firm belongs to serves as that firm's expected change in conservatism. The sample period spans 2000–2011. The dependent variable is EPS, defined as net income scaled by prior year's market value of equity. *R* is twelve-month buy-and-hold returns starting nine months before the fiscal year end. *D* is an indicator variable equal to one if *R* is negative, and zero otherwise. CDS is an indicator variable equal to one if a firm has a CDS contract traded over the sample period, and zero for matched control firms. POST is an indicator variable equal to one if a year falls in the two-year period after the CDS-trade-initiation year, and zero if a year falls in the two-year period prior to the CDS-trade-initiation year for CDS firms. The matched control firms take on the same value of POST as the CDS firms in the pre- and post-CDS-trade-initiation year, respectively. Additional controls include firm size, market-to-book ratio, book leverage, and their corresponding interaction terms with *R*, *D*, and *D* × *R*. Year and industry fixed effects are included. *p*-Values are derived based on robust standard errors clustered at the firm level.

### 5.4.2. *Alternative control for selection bias in CDS trade initiation*

The difference-in-difference research design used in our main analysis should mitigate the concern that our results are driven by firm characteristics that conceivably also determine CDS trade initiation. To further test the robustness of our analysis to an alternative specification, we use the Heckman two-stage procedure where the first stage models the probability of firms experiencing CDS trade initiations, and the second stage model includes the inverse mills ratio derived from the first stage to control for the selection bias. We estimate model (1) in the first stage; the second stage results are reported in column (1) Table 10. The coefficient on the inverse-mills ratio is positive, but statistically insignificant at conventional levels. The coefficient on POST × $D_t \times R_t$ continues to be negative and statistically significant, with a *p*-value of 0.014. In addition, we expand the first stage model by including the expected change in accounting conservatism as discussed in the previous sub-section; the corresponding second stage results are presented in column (2). The coefficient on the inverse-mills ratio in this specification is significantly positive, indicating the presence of selection bias. Importantly, our results on a significant decline in borrower conservatism upon CDS trade initiation are robust to this alternative procedure of controlling for selection bias.

### 5.4.3. *Robustness to using non-returns based measure of conservatism*

CDS trade initiation potentially influences equity price changes (see for example Boehmer et al., forthcoming), and in turn can conceivably influence the returns-based Basu measure of conservatism. Note that the magnitude and even direction of this influence is not obvious, and it is unlikely that any influence of CDS trade initiation on stock prices can generate the collective evidence we report. Nevertheless, to check robustness to a non-returns-based measure of conservatism, we utilize an alternative measure of asymmetric loss recognition based on an earnings time-series model (Basu, 1997; Ball and Shivakumar, 2005). Specifically, we estimate the following equation using ordinary least square regression with standard errors clustered at the firm level based on the sample consisting of both CDS firms and matched

**Table 10**
Robustness analysis – Heckman two-stage tests.

| | Dependent variable = $EPS_t$ | | | |
|---|---|---|---|---|
| | (1) | | (2) | |
| Variable | Coeff. Est. | *p*-Value | Coeff. Est. | *p*-Value |
| $R_t$ ($\beta_1$) | −0.109 | 0.497 | −0.112 | 0.481 |
| $D_t$ ($\beta_2$) | 0.271 | 0.001 | 0.270 | 0.001 |
| $D_t \times R_t$ ($\beta_3$) | 1.518 | 0.000 | 1.526 | 0.000 |
| POST ($\beta_8$) | −0.005 | 0.676 | −0.003 | 0.797 |
| POST × $R_t$ ($\beta_9$) | 0.050 | 0.018 | 0.048 | 0.025 |
| POST × $D_t$ ($\beta_{10}$) | −0.022 | 0.259 | −0.022 | 0.251 |
| POST × $D_t \times R_t$ ($\beta_{11}$) | **−0.270** | **0.014** | **−0.269** | **0.014** |
| Inverse mills ratio | 0.046 | 0.105 | 0.054 | 0.046 |
| Additional controls | Included | | Included | |
| Intercept ($\beta_0$) | −0.200 | 0.097 | −0.221 | 0.059 |
| Year and industry fixed effects | Included | | Included | |
| *F*-test: ($\beta_{13}+\beta_{15}$) | | | | |
| *F*-test: ($\beta_9+\beta_{11}$) | −0.220 | 0.036 | −0.221 | 0.034 |
| Number of firm-years | 1,996 | | 1,996 | |
| Adjusted $R^2$ (%) | 17.56 | | 18.12 | |

This table reports the second-stage regression results on the relation between Basu's (1997) measure of asymmetric loss recognition timeliness and the onset of CDS trading using Heckman two-stage procedure with CDS firms only. In the first stage, the sample consists of 138,735 firm-year observations, the same sample in Table 1. In column (1), the first stage selection model includes all the explanatory variables used in the first stage model as shown in Table 1. In column (2), the first stage selection model includes all the explanatory variables used in the first stage model as shown in Table 1, as well as the expected change in asymmetric loss recognition timeliness between one year before and one year after CDS trade initiation. For computing expected change in asymmetric timeliness, firms in Compustat are sorted into quintiles each fiscal year based on CSCORE, which in turn is computed following Khan and Watts (2009). For every firm-year, the change in the Basu coefficient from year −1 to year +1 for the CSCORE quintile the firm belongs to serves as that firm's expected change in conservatism. The sample period spans 2000–2011. The dependent variable is EPS, defined as net income scaled by prior year's market value of equity. *R* is twelve-month buy-and-hold returns starting nine months before the fiscal year end. *D* is an indicator variable equal to one if *R* is negative, and zero otherwise. POST is an indicator variable equal to one if a year falls in the two-year period after the CDS-trade-initiation year, and zero if a year falls in the two-year period prior to the CDS-trade-initiation year. Additional controls include firm size, market-to-book ratio, book leverage, and their corresponding interaction terms with *R*, *D*, and $D \times R$. Year and industry fixed effects are included. *p*-Values are derived based on robust standard errors clustered at the firm level.

control firms.

$$\begin{aligned}\Delta E_t = {} & \gamma_0+\gamma_1\Delta E_{t-1}+\gamma_2 D_{t-1}+\gamma_3 D_{t-1}\times\Delta E_{t-1}\\ & +\gamma_4\text{CDS}+\gamma_5\text{CDS}\times\Delta E_{t-1}+\gamma_6\text{CDS}\times D_{t-1}+\gamma_7\text{CDS}\times D_{t-1}\times\Delta E_{t-1}\\ & +\gamma_8\text{POST}+\gamma_9\text{POST}\times\Delta E_{t-1}+\gamma_{10}\text{POST}\times D_{t-1}+\gamma_{11}\text{POST}\times D_{t-1}\times\Delta E_{t-1}\\ & +\gamma_{12}\text{CDS}\times\text{POST}+\gamma_{13}\mathit{POST}\times\text{CDS}\times\Delta E_{t-1}+\gamma_{14}\text{CDS}\times\text{POST}\times D_{t-1}\\ & +\boldsymbol{\gamma_{15}}\textbf{CDS}\times\textbf{POST}\times\mathbf{D}_{t-1}\times\boldsymbol{\Delta}\mathbf{E}_{t-1}+\sum_{j=1}^{N}\lambda_j\text{ADDITIONAL CONTROLS}_j\\ & +\sum_{i=1}^{K}\gamma_i\text{INDUSTRY}_i+\sum_{m=1}^{N}\delta_m\text{YEAR}_m+\varepsilon_t \end{aligned} \quad (4)$$

where $\Delta E_t$ is current year's earnings change, $\Delta E_{t-1}$ is previous year's earnings change, *D* is an indicator variable equal to one for previous earnings decline (i.e., $\Delta E_{t-1}<0$) and zero otherwise, and the other variables are as defined in Eq. (3). In the above equation, we allow earnings persistence to differ between earnings increase and earnings declines. More timely recognition of losses than gains implies that earnings increases are more persistent than earnings declines. Hence, a reduction of accounting conservatism around the CDS-trade-initiation year would require that the coefficient on CDS × POST × $D_{t-1}\times\Delta E_{t-1}$, $\gamma_{15}$, be significantly positive.

Results presented in Table 11 corroborate those presented in Table 5.[22] The coefficient on CDS × POST × $D_{t-1}\times\Delta E_{t-1}$ is significantly positive at the 5% level (coefficient = 1.308, *p*-value = 0.017), implying that reversals in earnings declines are less in the years following CDS trade initiation. Our finding reinforces the conclusion that CDS firms experience a decline in the asymmetric timeliness of loss recognition after the onset of CDS trading.

[22] There is a reduction in sample size, because we impose the requirement that for every year, enough data be available to compute earnings changes for the following year. Since CDS trade initiations span 2002 to 2009, the sample period for this test extends from 1999 to 2011, given that we require lagged data to estimate changes in the earnings time-series measure of conservatism.

**Table 11**
Non-returns based measure of conservatism.

| **Earnings time-series measure** | Dependent variable = $\Delta EPS_t$ | |
|---|---|---|
| | CDS firms and matched firms | |
| Variable | Coeff. Est. | *p*-Value |
| $\Delta E_{t-1}$ ($\gamma_1$) | 0.163 | 0.662 |
| $D_{t-1}$ ($\gamma_2$) | −0.069 | 0.149 |
| $D_{t-1} \times \Delta E_{t-1}$ ($\gamma_3$) | −0.184 | 0.822 |
| CDS ($\gamma_4$) | −0.022 | 0.010 |
| CDS × $\Delta E_{t-1}$ ($\gamma_5$) | 0.307 | 0.044 |
| CDS × $D_{t-1}$ ($\gamma_6$) | 0.000 | 0.982 |
| CDS × $D_{t-1} \times \Delta E_{t-1}$ ($\gamma_7$) | −0.520 | 0.139 |
| POST ($\gamma_8$) | −0.019 | 0.036 |
| POST × $\Delta E_{t-1}$ ($\gamma_9$) | 0.426 | 0.014 |
| POST × $D_{t-1}$ ($\gamma_{10}$) | 0.001 | 0.960 |
| POST × $D_{t-1} \times \Delta E_{t-1}$ ($\gamma_{11}$) | −0.910 | 0.010 |
| CDS × POST ($\gamma_{12}$) | 0.006 | 0.640 |
| CDS × POST × $\Delta E_{t-1}$ ($\gamma_{13}$) | −0.499 | 0.043 |
| CDS × POST × $D_{t-1}$ ($\gamma_{14}$) | 0.012 | 0.611 |
| CDS × POST × $D_{t-1} \times \Delta E_{t-1}$ ($\gamma_{15}$) | **1.308** | **0.017** |
| Additional controls | Included | Included |
| Intercept ($\gamma_0$) | −0.119 | 0.000 |
| Year and industry fixed effects | Included | |
| Number of firm-years | 4,209 | |
| Adjusted $R^2$ (%) | 28.86 | |

This table reports multivariate regression results on the relation between an earnings time-series measure of asymmetric loss recognition timeliness and the onset of CDS trading. The sample period for this test spans 1999–2011. The period includes one extra year (1999) relative to Tables 5–10, due to the requirement of lagged data for estimation of the earnings-time-series-based measure of conservatism. Firms in financial industries are excluded. The dependent variable is $\Delta E_t$, the change in annual earnings before extraordinary item, scaled by lagged total assets. *D* is an indicator variable equal to one if $\Delta E_{t-1}$ is negative, and zero otherwise. CDS is an indicator variable equal to one if a firm has a CDS contract traded over the sample period, and zero for matched control firms. The control sample is chosen based on the matched-sample design, using the estimated probability of CDS trade initiation as described in Table 1. POST is an indicator variable equal to one if a year falls in the three-year period after the CDS-trade-initiation year, and zero if a year falls in the three-year period prior to the CDS-trade-initiation year for CDS firms. The matched control firms take on the same value of POST as the CDS firms in the pre- and post-CDS-trade-initiation year, respectively. Additional controls include firm size, market-to-book ratio, book leverage, and their corresponding interaction terms with $\Delta E_{t-1}$, $D_{t-1}$, and $D_{t-1} \times \Delta E_{t-1}$. Year and industry fixed effects are included. *p*-Values are based on robust standard errors clustered at the firm level.

## 6. Conclusion

Our paper provides evidence that the availability of CDS trades to a firm's lenders leads to a decline in that firm's reporting conservatism. Our findings are consistent with borrowers' incentives to report less conservatively after CDS trade initiation to avoid triggering violations, because they expect CDS-protected lenders to be less accommodating in renegotiations. Even in the presence of heightened incentives, borrowers would be constrained from reporting less conservatively if lenders were to maintain their monitoring. An actual decline in conservatism is thus also indicative of a slackening in lenders' post-CDS scrutiny of financial statements. The results thus imply that either (a) other stakeholders to the firm cannot replicate the monitoring by banks necessary to maintain borrowers' reporting conservatism and/or (b) that banks demand a level of conservatism higher than the equilibrium demand from other stakeholders.

It is possible that the lower reporting conservatism after CDS trade initiation encourages borrowers to also take actions that transfer wealth from debt-holders to shareholders via asset substitution, underinvestment, dividend overpayment, etc. In that sense, the structure of the CDS market during the sample period examined in this study may be off-equilibrium. On the other hand, the threat of more unfavorable renegotiations with lenders upon the occurrence of payment defaults, covenant violations and other credit events possibly disciplines borrowers *ex ante* from expanding attempts to transfer wealth from debt-holders. Thus, CDS availability potentially alters the lender–borrower relation to one that de-emphasizes continuous monitoring and financial reporting conservatism, and relies instead on the discipline imposed by the threat of lender inflexibility in renegotiations triggered by credit events. A thorough investigation into these possible scenarios following the onset of CDS trading is beyond the scope of this study, but can serve as a fertile area for future research.

## Appendix A. The credit default swap contract

Credit default swaps are generally documented using industry-standard derivative master agreements and standard CDS terms. Unlike equity shares or bonds, which are traded primarily on regulated exchanges, CDS are traded mainly over-the-counter (OTC). In principle, therefore, the contracting parties can agree upon the terms and conditions of the CDS individually – such as definitions of the credit events or settlement procedures. In practice, to facilitate documentation, avoid disputes regarding the occurrence of credit events and settle contracts, CDS contracting parties generally refer to the International Swaps and Derivatives Association (ISDA) Master Agreement. These general terms and conditions – established by ISDA, the central industry body – were introduced in 1999 and have been continuously developed since then. A revised version of the agreement was released in 2003, while the latest amendments were made in 2009 (see Deutsche Bank Research, 2009).

### A.1. *Sample term sheet for a credit default swap (traded by XYZ Bank PLC)*

**Draft terms – credit default swap**

**1. General terms**

| | |
|---|---|
| Trade Date | Aug 5, 2003 |
| Effective Date | Aug 6, 2003 |
| Scheduled Termination Date | Jul 30, 2005 |
| Floating Rate Payer ('Seller') | XYZ Bank plc, London branch |
| Fixed Rate Payer ('Buyer') | ABC Investment Bank plc |
| Calculation Agent | Seller |
| Calculation Agent City | New York |
| Business Day | New York |
| Business Day Convention | Following |
| Reference Entity | Jackfruit Records Corporation |
| Reference Obligation | Primary Obligor: Jackfruit Records |
| Maturity | Jun 30, 2020 |
| Coupon | 0% |
| CUSIP/ISIN | xxxxx |
| Original Issue Amount | USD 100,000,000 |
| Reference Price | 100% |
| All Guarantees | Not Applicable |

**2. Fixed payments**

| | |
|---|---|
| Fixed Rate Payer Calculation Amount | USD 7,000,000 |
| Fixed Rate | 0.3% per annum |
| Fixed Rate Payer Payment Date(s) | Oct 30, Jan 30, Apr 30, Jul 30, starting Oct 30, 2003 |
| Fixed Rate Day Count Fraction | Actual/360 |

**3. Floating payments**

| | |
|---|---|
| Floating Rate Payer Calculation Amount | USD 7,000,000 |
| Conditions to Payment (Or Seller) | Credit Event Notice (Notifying Parties: Buyer<br><br>Notice of Publicly Available Information: Applicable (Public Source: Standard Public Sources. Specified Number: Two) |
| Credit Events | Bankruptcy<br>Failure to Pay (Grace Period Extension: Not Applicable. Payment Requirement: $1,000,000) |
| Obligation(s) | Borrowed money |

**4. Settlement terms**

| | |
|---|---|
| Settlement Method | Physical Settlement |
| Settlement Currency | The currency in which the Floating Rate Payer Calculation Amount is denominated |
| Terms Relating to Physical Settlement | |
| Physical Settlement Period | The longest of the number of business days for settlement in accordance with the then – current market practice of any Deliverable Obligation being Delivered in the Portfolio, as determined by the Calculation Agent, after |

Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

| | |
|---|---|
| | consultation with the parties, but in no event shall be more than 30 days |
| Portfolio | Exclude Accrued Interest |
| Deliverable Obligations | Bond or Loan |
| Deliverable Obligation Characteristics | Not Subordinated<br>Specified Currency – Standard Specified Currencies<br>Maximum Maturity: 30 years<br>Not Contingent<br>Not Bearer<br>Transferable<br>Assignable Loan<br>Consent Required Loan |
| Restructuring Maturity Limitation | Not Applicable |
| Partial Cash Settlement of Loans | Not Applicable |
| Partial Cash Settlement of Assignable Loans | Not Applicable |
| Escrow | Applicable |

**5. Documentation**

Confirmation to be prepared by the Seller and agreed to by the Buyer. The definitions and provisions contained in the 2003 ISDA Credit Derivatives Definitions, as published by the International Swaps and Derivatives Association, Inc., as supplemented by the May 2003

Supplement, to the 2003 ISDA Credit Derivatives Definitions (together, the 'Credit Derivatives Definitions'), are incorporated into the Confirmation

**6. Notice and account details**

| | |
|---|---|
| Telephone, Telex and/or: Facsimile Numbers and Contact Details for Notices | Buyer<br>Phone:<br>Fax:<br>Seller: A.N. Other<br>Phone: +1 212-xxx-xxxx<br>Fax: +1 212-xxx-xxxx |
| Account Details of Seller | 84-7512562-85 |

### *A.2. Risks and characteristics*

*Credit risk*. An investor's ability to collect any premium will depend on the ability of XYZ Bank plc to pay.

*Non-marketability*. Swaps are not registered instruments and they do not trade on any exchange. It may be impossible for the transactor in a swap to transfer the obligations under the swap to another holder. Swaps are customized instruments and there is no central source to obtain prices from other dealers.

## Appendix B

Panel A: Buyers of protection by institution type

| Type of institution | 2000 | 2002 | 2004 | 2006 |
|---|---|---|---|---|
| Banks (including securities firms) | 81 | 73 | 67 | 59 |
| Banks – trading activities | – | – | – | 39 |
| Banks – loan por tfolio | – | – | – | 20 |
| Insurers | 7 | 6 | 7 | 6 |
| Monoline insurers | – | 3* | 2 | 2 |
| Reinsurers | – | | 3 | 2 |
| Other insurance companies | – | 3 | 2 | 2 |
| Hedge funds | 3 | 12 | 16 | 28 |
| Pension funds | 1 | 1 | 3 | 2 |
| Mutual funds | 1 | 2 | 3 | 2 |
| Corporates | 6 | 4 | 3 | 2 |
| Other | 1 | 2 | 1 | 1 |

Panel B: Sellers of protection by institution type

| Type of institution | 2000 | 2002 | 2004 | 2006 |
|---|---|---|---|---|

| | | | | |
|---|---|---|---|---|
| Banks (including securities firms) | 63 | 55 | 54 | 44 |
| Banks—trading activities | – | – | – | 35 |
| Banks—loan portfolio | – | – | – | 9 |
| Insurers | 23 | 33 | 20 | 17 |
| Monoline insurers | – | 21* | 10 | 8 |
| Reinsurers | – | | 7 | 4 |
| Other insurance companies | – | 12 | 3 | 5 |
| Hedge funds | 5 | 5 | 15 | 32 |
| Pension funds | 3 | 2 | 4 | 4 |
| Mutual funds | 2 | 3 | 4 | 3 |
| Corporates | 3 | 2 | 2 | 1 |
| Other | 1 | 0 | 1 | 1 |

*Monoline insurers and reinsurers combined.

This appendix shows the breakdown of market share in percentage for CDS market participants by type of institutions. The source is British Bankers Association (BBA) Credit Derivatives Report (2006).

## References

Acharya, V., Almeida, H., Ippolito, F., Perez, A., 2014. Credit lines as monitored liquidity insurance: Theory and evidence. Journal of Financial Economics 112, 287–319.
Acharya, V., Johnson, T., 2007. Insider trading in credit derivatives. Journal of Financial Economics 84, 110–141.
Ahmed, A., Billings, B., Morton, R., Stanford, M., 2002. The role of accounting conservatism in mitigating bondholder-shareholder conflicts over dividend policy and in reducing debt costs. The Accounting Review 77, 867–890.
Arping, S., 2012. Credit protection and lending relationships. Journal of Financial Stability 10, 7–19.
Ashcraft, A., Santos, J., 2009. Has the CDS lowered the cost of corporate debt? Journal of Monetary Economics 56 (4), 514–523.
Ball, R., Shivakumar, L., 2005. Earnings quality in UK private firms: comparative loss recognition timeliness. Journal of Accounting and Economics 39, 83–128.
Ball, R., Bushman, R., Vasvari, F., 2008. The debt-contracting value of accounting information and loan syndicate structure. Journal of Accounting Research 46, 247–287.
Basu, S., 1997. The conservatism principle and the asymmetric timeliness of earnings. Journal of Accounting and Economics 24, 3–37.
Begley, J., Freedman, R., 2004. The changing role of accounting numbers in public lending agreements. Accounting Horizons 18, 81–96.
Boehmer, E., Chava, S., Tookes, H., 2014. Related securities and equity market quality: the case of CDS. Journal of Financial and Quantitative Analysis. (forthcoming).
Bolton, P., Oehmke, M., 2010. Credit default swaps and the empty creditor problem. Review of Financial Studies 24, 2617–2655.
British Bankers Association (BBA), 2006. BBA Credit Derivatives Report. BBA Enterprises Ltd., London2006.
Bushman, R.M., Wittenberg-Moerman, R., 2012. The role of bank reputation in "certifying" future performance implications of borrowers' accounting numbers. Journal of Accounting Research 50, 883–930.
Chava, S., Roberts, M., 2008. How does financing impact investment? The role of debt covenants. Journal of Finance 63, 2085–2121.
Chen, T., Martin, X., 2011. Do bank affiliated analysts benefit from information sharing? Journal of Accounting Research 49, 633–675.
Deutsche Bank Research, 2009. Credit default swaps heading towards a more stable system.
Donovan, J., Frankel, R., Martin, X., 2014. Accounting conservatism and creditor recovery rate, Working Paper at Washington University in St. Louis.
Duffee, D., Zhou, C., 2001. Credit derivatives in banking: useful tools for managing risk? Journal of Monetary Economics 48, 25–54.
Ettredge, M., Huang, Y., Zhang, W., 2012. Earnings restatements and differential timeliness of accounting conservatism. Journal of Accounting and Economics 53, 489–503.
Fama, E., 1985. What's different about banks? Journal of Monetary Economics 15 (29), 39.
Fink, R., 2004. Credit Default Swaps: Who's Doing What? CFO.com, October 7, 2004.
Frankel, R., Litov, L., 2008. Financial accounting characteristics and debt covenants, Working Paper, Washington University in St. Louis.
Gopalan, R., Nanda, V., Yerramilli, V., 2011. Does poor performance damage the reputation of financial intermediaries? Evidence from the loan syndication market. Journal of Finance 66, 2083–2120.
Gormley, T., Kim, B., Martin, X., 2012. Do firms adjust their timely loss recognition in response to changes in the banking industry. Journal of Accounting Research 50, 159–196.
Gorton, G., Khan, J., 1993. The design of bank loan contracts. Review of Financial Studies 13, 331–364.
Gorton, G., Pennacchi, G., 1995. Banking and loan sales: marketing nonmarketable assets. Journal of Monetary Economics 35, 389–411.
Gupta, A., Singh, A., Zebedee, A., 2008. Liquidity in the pricing of syndicated loans. Journal of Finance 11, 339–376.
Hu, H., Black, B., 2006. The new vote buying: empty voting and hidden (morphable) ownership. Southern California Law Review 79, 811–908.
James, C., 1987. Some evidence on the uniqueness of bank loans. Journal of Financial Economics 19, 217–235.
Khan, M., Watts, R., 2009. Estimation and empirical properties of a firm-year measure of accounting conservatism. Journal of Accounting and Economics 48, 132–150.
LaFond, R., Roychowdhury, S., 2008. Managerial ownership and accounting conservatism. Journal of Accounting Research 46, 101–135.
LaFond, R., Watts, R., 2008. The information role of conservative financial statements. The Accounting Review 83, 447–478.
Lee, I., Do firms knowingly sell overvalued equity?, The Journal of Finance 52, 1997, 1439-1466.
Leone, A.J., Wu, J.S., Zimmerman, J.L., 2006. Asymmetric sensitivity of CEO cash compensation to stock returns. Journal of Accounting and Economics 42, 167–192.
Lewis, M., 2010. The Big Short. W.W. Norton and Company, New York, NY.
Marsh, I., 2009. The effect of lenders' credit risk transfer activities on borrowing firms' equity returns, Working Paper. City University, London.
Mayordomo, S., Pena, J., Schwartz, E., 2014. Are all credit default swap databases equal? European Financial Management 20, 677–713.
Minton, B., Stulz, R., Williamson, R., 2009. How much do banks use credit derivatives to hedge loans? Journal of Financial Service Research 35, 1–31.
Morrison, A., 2005. Credit derivatives, disintermediation, and investment decisions. Journal of Business 78, 621–647.
Nikolaev, V., 2010. Debt covenants and accounting conservatism. Journal of Accounting Research 48, 137–175.
Parlour, C., Winton, A., 2013. Laying off credit risk: loan sales versus credit default swaps. Journal of Financial Economics 107, 25–45.
Pennacchi, G., 1988. Loan sales and the cost of bank capital. Journal of Finance 43, 375–396.
Petersen, M., 2009. Estimating standard errors in finance panel data sets: comparing approaches. Review of Financial Studies 22, 435–480.
Roberts, M., Sufi, A., 2009. Renegotiation of financial contracts: evidence from private credit agreements. Journal of Financial Economics 93, 159–184.
Roychowdhury, S., 2010. Discussion of acquisition profitability and timely loss recognition. Journal of Accounting and Economics 49, 179–183.

Please cite this article as: Martin, X., Roychowdhury, S., Do financial market developments influence accounting practices? Credit default swaps and borrowers' reporting conservatism. Journal of Accounting and Economics (2014), http://dx.doi.org/10.1016/j.jacceco.2014.09.006

Roychowdhury, S., Watts, R.L., 2007. Asymmetric timeliness of earnings, market-to-book and conservatism in financial reporting. Journal of Accounting and Economics 44, 2–31.
Saretto, A., Tookes, H., 2013. Corporate leverage, debt maturity and credit default swaps: the role of credit supply. Review of Financial Studies 26, 1190–1247.
Stulz, R., 2010. Credit default swaps and the credit crisis. Journal of Economic Perspectives 24, 73–92.
Subrahmanyam, M., Tang, D., Wang, S., 2014. Does the tail wag the dog? The effect of credit default swaps on credit risk. Review of Financial Studies 27, 2927–2960.
Sufi, A., 2007. Information asymmetry and financing arrangements: evidence from syndicated loans. Journal of Finance 62, 629–668.
Sustersic, J., 2012. Do traded credit default swaps impact lenders' monitoring activities? Evidence from private loan agreements (Unpublished dissertation). Ohio State University.
Tan, L., 2013. Creditor control rights, state of nature verification, and financial reporting conservatism. Journal of Accounting and Economics 55, 1–22.
Venokur, S., Magidson, M., Singe, A., 2008. Comparing credit default swaps to insurance contracts: did the New York State Insurance Department get it right? Journal of the Law of Investment and Risk Management Products 28, 1–11.
Wang, Y., Xia, H., 2014. Do lenders still monitor when they can securitize loans? Review of Financial Studies 27, 2354–2391.
Watts, R., Zimmerman, J., 1986. Positive Accounting Theory. Prentice-Hall, Englewood Cliffs, NJ.
Watts, R., 2003. Conservatism in accounting. Part I: explanations and implications. Accounting Horizons 17, 207–221.
Weistroffer, C., 2009. Credit default swaps: heading towards a more stable system. Deutsche Bank Research: Current Issues.
Zhang, J., 2008. Efficiency gains from accounting conservatism: benefits to lenders and borrowers. Journal of Accounting & Economics 45, 27–54.