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Does institutional shareholder activism stimulate corporate information flow?☆

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ABSTRACT

Activist shareholders have an incentive to communicate and cooperate with other major shareholders. However, the impact of their activity on information flow surrounding targeted firms is largely unknown. We explore this aspect using a prolific proponent: labor unions. Following the mailing of proxies containing union-sponsored shareholder proposals, trading volume increases significantly and at-issue bond yield spreads of targeted firms are lower compared to matched firms. Subsequent difference-in-differences analyses show that stock prices of targeted firms become more informative as a result of activism, affirming the intuition that activism results in a reduction of differential information between outside investors.

1. Introduction

The shareholder proposal mechanism continues to be an important tool used by institutional shareholder activists to make changes to corporations. In each proxy season, hundreds of US listed companies receive, screen, and include these proposals (along with boards' responses) in proxies mailed to their shareholders. In this process sponsors are permitted, and motivated, to communicate and share information with other groups of major shareholders. In this paper we investigate the impact of union shareholder proposals on the information flow and stock price informativeness associated with targeted firms.

While a variety of institutional shareholders engage in activism and share information, activism by labor unions and their affiliated funds provide a particularly interesting setting for this investigation. First, union activism represents a clear case of which there is diversity of information sets to be shared among institutional investors during the activism process. Second, the tests have more power due to the prolific nature of union activism over the past two decades. Since the 1990 s, unions have occupied a prominent space within the spectrum of corporate stakeholders – that of shareholder activists (e.g., Gillan and Starks, 2007). The ultimate success of these efforts necessarily relies on engagement with, and support from, other large shareholders.

Studying 1362 shareholder proposals sponsored by unions and labor-affiliated funds during the 1988 to 2010 proxy seasons, we find that trading volume increases in the period immediately following the proxy mailing date. Bonds that are issued during this time period enjoy relatively lower yield spreads compared to those issued by comparable untargeted firms, supporting the view that communication associated with activism reduces information risk. Further, difference-in-differences (DiD) analyses show that stock prices of targeted firms become more informative relative to a matched set of firms using the information-based trading measure introduced by Llorente et al. (2002) over the one-year period following the activism. These effects are pronounced for targeted firms with high institutional equity ownership, implying that interactions between unions and other institutional shareholders facilitate the flow of unions' firm-specific information to other market participants. Our DiD results also indicate that the more informative prices are not due to a reduction in the layer of information asymmetry that arises from the informational mismatch between managers and outsiders, as documented by Luez and Verrecchia (2000). Collectively, these results suggest that shareholder activism by unions add to the information flow surrounding targeted firms.

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Our work contributes to several strands of the literature. First, we add to the shareholder activism literature. Existing work generally evaluates the benefits of shareholder activism through the announcement returns around the proposal event and its subsequent impact on metrics such as operating performance or corporate governance quality. For instance, Brav et al. (2008) study these dimensions associated with activism by hedge funds, while Prevost et al. (2012) focus on union activism.[1] To our knowledge, our study is the first to focus on the impact of shareholder activism on the information flow between different groups of investors. Second, we contribute to the debate over whether diverse information in financial markets attracts or deters the transmission and/or production of more information. In recent theoretical work, Goldstein and Yang (2015) show that greater diversity of information among different groups of large investors improves stock price informativeness. Goldstein and Yang (2015) consider the case where traders observe other traders' information through trading activity. However, to the extent that non-union institutional investors learn private information from unions during the activism, there should be a reduction in the aspects of information they are uncertain about (e.g. outside the range of their expertise). We contribute to this work by documenting increases in information-based trading as a result of interactions among union shareholders and other institutional shareholders associated with the activism process. Finally, while the majority of existing work (e.g. Hilary, 2006; Bova, 2013) focuses on the negative informational effects related to labor's role as employees (i.e. the effect of union presence on information flow from corporate insiders to outsider investors), we investigate the hitherto unexamined impact unions, in their alternative stakeholder role as shareholders, have on the improved information flow among corporate outsiders.

2. Literature review and hypotheses development

As with other individual and institutional shareholders, unions engage in the shareholder activism process via shareholder proposals. Shareholder proposals are submitted under Securities Exchange Commission (SEC) Rule 14a-8, which allows shareholders who meet a minimal ownership threshold to place a proposal alongside management proposals on the proxy. Unions have long played a leading role in the shareholder activism landscape. As noted by Prevost et al. (2012), unions have a lengthy track record of prolific activism stemming to the early 1990 s, making them highly experienced, and visible, institutional shareholder proponents. For example, according to Copland and O' Keefe (2013) institutional investors affiliated with organized labor sponsored about one-third of all shareholder proposals submitted during 2006–2013.

There is an ongoing debate as to whether union shareholder activism is motivated by wealth-maximizing objectives that align with the interests of other institutional investors, or instead is a mechanism to promote the welfare of the unions' member constituents. For example, Copland and O' Keefe (2013) contend that "labor-affiliated pension funds have tended to focus their shareholder-proposal activism on companies and sectors that seem to have little to do with share value but may be related to labor-organizing efforts or other labor disputes with company management, or otherwise a political agenda." Support for this view is provided by Agrawal (2012), who finds empirical evidence of AFL-CIO union funds pursuing objectives consistent with worker interests. Similarly, Del Guercio and Woidtke (2014) find that directors who comply with union proposals experience a net loss in external board seats, and interpret this as evidence that the external directorships market views union proposals as self-serving. On the contrary, Cunat et al. (2012) show that board decisions to implement corporate governance-related proposals add the most value when sponsored by union and public pension funds. Ertimur et al. (2011) document that unions are not more likely to target highly unionized companies, or firms involved in disputes with labor, in comparison to non-union activists. Consistent with Schwab and Thomas' (1998, p. 1023) view that "other shareholders are generally able to distinguish, on a case-by-case basis, which hat the union shareholder is wearing", Ertimur et al. (2010) show that voting recommendations by proxy advisor firms (i.e., Institutional Shareholder Services) are less likely and shareholder voting support is significantly lower when union activists represent the interests of both shareholders and workers.[2]

Despite the mixed empirical findings regarding the underlying motivation for union activism, we contend that union activists generally have a motivation to cooperate with other investors in order to attain their activism goals (Schwab and Thomas, 1998). In seeking this support, the sharing of different pieces of information among different groups of institutional investors stimulates corporate information flow. We investigate if union activism aimed at corporate governance issues increases information-based trading by improving the availability and intensity of information flow between significant shareholders. As discussed by Schwab and Thomas (1998), unions' role as dual stakeholders affords them access to information in some companies that other shareholders may not have due to their regular involvement with companies, their analysis of industry wide information, and the input of specialist advisors. To the extent that unions are well connected, highly experienced, and visible players in the shareholder activism arena and to the extent that they are motivated to cooperate with other institutional investors to achieve their activism objectives, the activism process serves as a mechanism for facilitating the flow of additional information available to union proponents directed towards other significant shareholders. Accordingly, we examine if union activism is associated with trading activities and information based trading:

H1: Firms targeted by unions on corporate governance issues are associated with improved information flow and stock price informativeness

We expect that the effect of union activism on the information environment to be conditional on the level of institutional equity ownership at targeted firms. As noted by Dennis and Weston (2001) and Chemmanur et al. (2013), institutional investors have an economic advantage in the precision and cost of collecting information. Indeed, prior work suggests that institutional shareholders possess an informational advantage over retail investors (e.g., Szewczyk et al., 1992; Alangar eta l., 1999; Bartov et al., 2000). However, some institutions may be better placed than others in their access to different pieces of information, resulting in an information dissemination role within institutional investor communication networks. Anecdotal evidence suggests that unions work with other institutional investors to achieve common objectives. For example, Laroux (2012) points out that while unions do not typically hold large proportions of equity in U.S. corporations, they exert a disproportionate amount of influence due to

[1] Gillan and Starks (2007) note that several studies show activism results in short-term positive abnormal returns but the impact on longer term changes in shareholder wealth, operating performance, and corporate governance quality is less clear.

[2] Another strand of the union literature (e.g. Kleiner and Bouillon, 1988; Hilary, 2006; Chen et al., 2011; Chung et al., 2012; Bova, 2013) considers whether managers strategically withhold financial information to improve their bargaining position with strong unions, resulting in greater information asymmetry between the firm and market participants. In contrast, our study examines if union shareholders play a role in reducing asymmetry between market participants.

their proximity to other institutional investors that do.[3] Laroux (2012) notes that unions often align themselves with large public pension funds in order to achieve their objectives.[4] Therefore, to the extent that there are scale economies in acquiring and using information for institutional owners, and since unions have an incentive to communicate with other shareholders in order to increase the likelihood of a successful activism outcome (Schwab and Thomas, 1998), the proportion of institutional ownership should play a significant role in the ability of union activism to affect information flow and stock price informativeness:

H2: Increases in information flow and information based trading associated with union shareholder activism is directly related to the level of institutional equity ownership.

In addition, we investigate the pricing implications of information flow associated with activism. A long line of theoretical and empirical work shows that information risk is priced into financial assets. Bond yield spreads provide advantages over stock returns in that spreads are deterministic and have clearly defined components related to default, liquidity, and information. Prior research demonstrates that the quality of the information held by outsiders is reflected by yield spreads. For example, Livingston and Zhou (2010) show that bonds with differing Moody and S&P ratings are associated with higher yields, indicating that impaired information related to disagreement about the risks of the firms' underlying assets leads to a higher cost of debt capital. Therefore, to the extent that targeting acts as an impetus for improved information flow among corporate outsiders, we expect that the yield spreads of targeted firms should be systematically lower after controlling for other determinants of yield spreads. Further, following Hypothesis *H2*, the effect should be increasing in the level of equity held by institutions. These conjectures collectively lead to Hypothesis *H3*:

H3: Information flow associated with union shareholder activism is priced into bonds issued by targeted issuers, and the pricing effect is stronger when institutional equity ownership levels are higher.

3. Data and sample description

We focus on proposals that address corporate governance issues: these proposals are most likely to be associated with significant information effects due to the likelihood of being of interest to other institutional investors. Our shareholder proposal data is obtained from two sources. For the period 1988–2002, a comprehensive list of shareholder proposals sponsored by labor unions and other proponents that went to a vote is obtained from the Investor Responsibility Research Center (IRRC). The 1988–2002 segment of the sample is comprised of 467 labor union-sponsored proposal events for which mailing dates are available in SEC Edgar. For the 2003–2010 period, we obtain labor-sponsored shareholder proposals from GMI Ratings' Shareholder Proposal Database. The 2003–2010 portion of the dataset is based on 895 proposals. Together, the primary sample comprises 1362 labor-sponsored proposals that are spread out over 1086 proxies. For each proposal event, we collect the proxy mailing date from the proxy's cover letter to shareholders.

Table 1
Labor-sponsored governance-related shareholder proposals.

Panel A1: CII members	**No. proposals**	**Proportion of sample**
United Brotherhood of Carpenters	292	0.21
American Federation of State, County and Municipal Employees	120	0.09
AFL-CIO	107	0.08
International Brotherhood of Electrical Workers	106	0.08
Teamsters	103	0.08
Sheet Metal Workers International Association	71	0.05
Communication Workers of America	43	0.03
Laborers International	41	0.03
Service Employees International Union	38	0.03
Central Laborers	35	0.03
Massachusetts Laborers	33	0.02
International Union of Operating Engineers	23	0.02
Trowel Trades	21	0.02
Plumbers and Pipefitters	13	0.01
United Food and Commercial Workers International Union	12	0.01
United Auto Workers	2	0.00
Panel A2: all other proponents	302	0.22
Panel B: proposal content		
Board structure and composition	543	0.40
Executive compensation	97	0.07
Voting	533	0.39
Antitakeover	34	0.03
Other	155	0.11
Panel C: voting outcomes		
Votes-for percentage		0.37
Proposals achieving majority voting support	329	0.24
No. proposal observations	1362	

Table 1 provides numbers of proponents, issues addressed by proposals, and voting outcomes for the primary sample of 1362 union-sponsored proposals covering the sample period 1989–2010.

Table 1 provides descriptive statistics for the primary sample of proposals. Unions that are members of the Council of Institutional Investors (CII), a non-profit organization that advocates corporate governance and shareholder rights, are heavily represented in our dataset: Panel A shows that unions that are 2013 members of the CII are responsible for 78% of proposals submitted during the sample period. The most frequent union proponent is the United Brotherhood of Carpenters (UBC) with a total of 292 proposals over the sample period, while the least frequent is the United Auto Workers with 2 proposals.[5] The remaining unions and labor-affiliated funds (e.g., Longview Fund) make up the remaining 22% of the sample. In Panel B, we follow the approach taken by Prevost et al. (2012) by categorizing the issues addressed by proposals into four broad categories: Board structure and composition, executive compensation, voting-related, and antitakeover. Union proposals that focus on board- and voting-related issues account for nearly 80% of all proposals with approximately 40% in each category, with executive compensation and antitakeover proposals accounting for about 10%. Proposals that cannot be placed in these four categories make up the remaining 10%.[6] Finally, Panel C

[3] Recent work documents that proposals sponsored by institutional investors targeting corporate governance issues are more likely to be supported by other shareholders (Ferri and Sandino, 2009), and that the consequences of activism are associated with the identity of the sponsor (Ertimur et al., 2010) and voting support by shareholders (Cotter and Thomas, 2007).

[4] Laroux (2012) further states "[of] the 124 members of the highly influential Council of Institutional Investors, half are public pension funds. The remainder comprises corporate pension funds (32), union pension funds (22) and special-purpose funds (8) such as Ceres. When public pension funds and unions align on matters of Council policy, they hold sway."

[5] The AFSCME is the second most prolific fund in our sample, with a total of 120 proposals that went to a vote over the sample period. In contrast to the other labor unions in our sample, the AFSCME represents workers in the public sector. However, the AFSCME is owned by the AFL-CIO (third most prolific fund), suggesting that its objectives and access to information at target companies is similar to private sector unions.

[6] As an example of a proposal outside these four categories, the Central Laborers Pension, Welfare & Annuity Funds submitted the following proposal in 2006:

provides a summary of voting outcomes achieved by proposals. On average, labor-sponsored proposals achieve a votes-for percent of 37% when the proposal is taken to a vote at the annual meeting, with 329 proposals in our sample (24%) achieving majority support (>50%) from shareholders.

4. Short-term effect of union activism on abnormal trading volume

4.1. Methodology

We begin our empirical analysis by analyzing abnormal changes in trading activity during the period immediately following the mailing date of proxy statements that contain shareholder proposals submitted by labor unions. We estimate abnormal changes in trading volume associated with union activism with log-transformed relative volume. As described by Campbell and Wasley (1996), market model abnormal trading volume (v_{it}) is obtained as (actual – predicted) volume, i.e. $v_{it} = V_{it} - (\alpha_i + \beta_i V_{mt})$: V_{it} is the log-transformed percentage of shares traded on day t, i.e. $V_{it} = \frac{n_{it} \times 100}{S_{it}}$ where n_{it} is the number of shares traded for firm i on day t and S_{it} is firm i's outstanding shares; α_i and β_i are based on least-squares estimation over a 100-day (−131, −31) estimation period relative to each proxy mailing date; and V_{mt} is the market volume measure on day t and is measured as $V_{mt} = \frac{1}{N}\sum_{i=1}^{N} V_{it}$, where N is the number of securities in the CRSP value-weighted market index.

Because it is not clear when behind-the-scenes dialogue between union proponents and other institutions takes place, we estimate cumulated abnormal trading volume over a period of three months (60 trading days), beginning on the proxy mailing date.[7] Mean and median cumulative average abnormal relative trading volume (CAARV) are reported for individual one month (20 trading day) intervals within the overall 60-day time period. We employ the standardized cross-sectional z-statistic discussed by Boehmer et al. (1991) to assess the statistical significance of each window. We also report whether the proportion of positive abnormal trade volume in the event period is statistically different than in the estimation period using the non-parametric generalized sign test (Cowan, 1992) as well as the nonparametric generalized rank statistic proposed by Kolari and Pynnonen (2011) which may be more appropriate if there are non-normalities in the distributions of abnormal volume.

An increase in abnormal volume may indicate information flowing from insiders to outsiders or information flowing between outsiders. For example, Nofsinger and Sias (1999) provide evidence of the latter by showing institutional investors herd toward (away from) undervalued (overvalued) stocks and argue that institutions trade based on value-relevant information about the firm. As our overall abnormal volume results may be evidence of either type information flow (insider to outsider or outsider to outsider), we generate additional results based on subsets sorted on institutional ownership percentage. If heightened communication between institutions associated with union activism is indeed what occurs, then increases in trading volume should be much larger in high institutional ownership stocks.

4.2. Empirical results

Table 2 provides the univariate abnormal relative volume results. Hypothesis *H1* regarding improved information flow is preliminarily tested in Panel A. In general, the results support our hypothesis that labor union activism is associated with greater information flow (via higher trading activities): Cumulated average abnormal volume during the three-month period following the proxy mailing date is 84.51% with the largest increases over the two months immediately following the mailing date. Mean CAARV over the (0, 20) and (21, 40) windows are 53.35 and 51.48%, respectively, and are both statistically significant using the z- and generalized rank test statistics.

We examine if the level of institutional share ownership plays a role in the increased trading activity. Panels B-C bifurcate the sample into above- and below-median institutional ownership subsets by sorting the sample according to institutional ownership. This provides a preliminary test of Hypothesis *H2*, which posits that the information effects of proposals should be stronger for targeted firms with greater levels of institutional ownership. Panel B demonstrates that cumulated (0, 60) abnormal trading volume of firms in the above-median institutional ownership subset is approximately 167%, with the largest increase in the month immediately following the mailing date. In contrast, Panel C illustrates that firms in the below-median subset of institutional ownership are associated with insignificant changes in trading volume over this time period, providing evidence that institutional ownership is a primary driver in the results of Table 2 Panel A.[8] This also supports our contention that our results are caused by a reduction in information asymmetry between different groups of institutional investors rather than between corporate insiders and outside investors.[9]

5. The effect of union activism on at-issue bond yield spreads

5.1. Pricing effects using propensity score matching

The previous analysis suggests that there are informational effects following the submission of union-sponsored proposals, especially among highly institutionally owned firms, that are reflected by abnormal trading volume. In this section, we investigate the pricing impact of such improved information flow by examining the effects of targeting on at-issue corporate bond yield spreads, which are based on information, default, and liquidity components. As noted by Bessembinder and Maxwell (2008) among others, the corporate bond market is dominated by institutions making the at-issue market an appropriate venue for investigating the effects of information flow through institutional channels. A variety of papers (e.g. Butler, 2008; Zhou, 2010; Mansi et al., 2011) show that information is efficiently priced by bond market participants.

"Resolved that the shareholders of Chubb Corporation ("Company") hereby request that our company provide a report, updated semi-annually, disclosing our Company's policies and procedures for political contributions, both direct and indirect, made with corporate funds."

[7] While this interval approximately covers the time period between proxy mailing- and annual meeting dates for most firms, the flow of information is unobservable. The difference-in-differences analyses below add to the robustness of the short-term analysis by demonstrating that prices of targeted firms become more informed over longer periods following targeting.

[8] To ensure that first-quarter earnings announcements that coincide with the proxy mailing date are not influencing our results, we perform a robustness check by restricting the sample to proposals with mailing dates following the first quarter. The number of firm-year observations decreases to 400, however the results remain qualitatively similar to those presented here: Mean (median) CAARV for the 200-observation above-median institutional ownership subset for the [0,20] and [21,40] windows are positive and significant at the 5% level, while all of the event windows for the below-median subset are insignificant.

[9] In unreported cross-sectional regressions, we control for additional firm characteristics (e.g. size and the unionization rate at targeted companies) and proposal characteristics (e.g. proxies with multiple proposals and proposals that achieve a majority vote at the annual meeting) that may also impact abnormal trading volume. The coefficient estimate for institutional ownership is significant in these regressions at the 5% level. We also find that proposal-specific indicator variables including the type of issue addressed by the proposal (antitakeover, board-related, or compensation related) are not significantly related to abnormal trading volume.

Table 2
Labor union proposal mean cumulative average abnormal relative volume.

Panel A: all firm-year proposals						
Event window	**No. obs.**	**Mean CAARV**	**Median CAARV**	**Positive: negative**	**Cross-sectional Z**	**Generalized rank t**
[0,20]	893	0.5335	0.1423	450:443**	2.835***	1.776*
[21,40]	893	0.5148	0.1795	456:437**	2.561**	1.848*
[41,60]	892	−0.2034	−0.6283	404:488	−0.170	−0.239
[0,60]	893	0.8451	−0.2556	435:458	1.954*	1.312
Panel B: institutional ownership > median						
Event window	**No. obs.**	**Mean CAARV**	**Median CAARV**	**Positive: negative**	**Cross-sectional Z**	**Generalized rank t**
[0,20]	447	0.8750	0.7717	244:203***	3.299***	2.588**
[21,40]	447	0.6379	0.1795	229:218*	1.877*	1.768*
[41,60]	447	0.1554	−0.2888	211:236	0.171	0.416
[0,60]	447	1.6682	0.5018	230:217**	1.972**	1.933*
Panel C: institutional ownership ≤ median						
Event window	**No. Obs.**	**Mean CAARV**	**Median CAARV**	**Positive: negative**	**Cross-sectional Z**	**Generalized rank t**
[0,20]	446	0.1963	−0.2491	215:231	0.954	0.589
[21,40]	446	0.2773	−0.1895	216:230	1.160	1.119
[41,60]	446	−0.6100	−0.7824	196:249	−1.233	−0.932
[0,60]	446	−0.1350	−1.1969	206:240	0.212	0.251

Table 2 presents cumulative average abnormal relative trading volume (CAARV) following the mailing of proxies containing union-sponsored shareholder proposals during the 1989–2010 period. The first column identifies the event window where day 0 is the proxy mailing date. Abnormal relative trading volume for each day is the difference between actual and predicted log-transformed percentage of shares traded on day *t*. The standardized cross-sectional z-statistic tests for the significance of each CAARV using a two-tail test, and the generalized sign test indicates if the percentage of positive abnormal volume in the event period is significantly different than in the estimation period. Panel A represents the full sample of the mailing dates of proxies containing union-sponsored proposals, while Panel B (C) covers subsample bifurcated by high (low) institutional equity ownership. *, **, and * correspond to significance at the 1, 5, and 10% levels, respectively.

We examine differences in yield spread between targeted and non-targeted control firms using the propensity score matching methodology (Rosenbaum and Rubin, 1983 & 1985) which mitigates selection bias by matching sample firms with control firms with similar characteristics according to a function of covariates.[10] Our sample of treated and untreated at-issue yield spread observations is drawn from the SDC Platinum database, covering the sample period 1989–2010. We limit the sample to “plain vanilla” fixed rate bonds by eliminating exotic structures such as asset-backed bonds and extendible notes. Since information flow may be different in regulated industries, we exclude issuers classified as financial (6000 < = SIC < = 6999) and utilities (4900 < = SIC < = 4999).[11]

The outcome variable is yield spread, which is calculated as (i_{Corp}- i_{Govt}) where i_{Corp} is the at-issue yield-to-maturity of the sample corporate bond and i_{Govt} is the interpolated yield-to-maturity for the point on the Treasury yield curve corresponding to the same time to maturity as the sample corporate bond. We obtain monthly constant-maturity Treasury bond indices to calculate the interpolated yield curve from the Federal Reserve of St. Louis Economic Data (*FRED*).

Targeted firms are matched to control firms using propensity scores using the following logit model:

$$\begin{aligned} Union\,target = {} & \alpha_0 + \alpha_1 One\text{ - }year\ stock\ return + \alpha_2 Three \\ & - year\ sales\ growth + \alpha_3 Insider\ ownership \\ & + \alpha_4 Institutional\ ownership + \alpha_5 Firm\ size \\ & + \alpha_6 Unionization\,rate \end{aligned} \tag{1}$$

To maintain consistency with the time frame used in the trading activity analysis, we define *Union target* as a union targeting event that occurs within a total of three months (60 trading days) following the issue offering date. The covariates are based on the specification of Karpoff et al. (1996), who investigate the causes and consequences of shareholder proposal activism over the 1986–1990 sample period. They model the likelihood of targeting with alternative performance measures, firm size, financial leverage, and equity ownership by institutions and corporate insiders. Karpoff et al. (1996) provide evidence that performance is inversely related to the likelihood of targeting, larger firms are more likely to be targeted, and institutional and corporate insider equity ownership is significantly correlated with the targeting choice. We measure performance alternatively with *One-year stock return* (cumulated return over the year prior to the mailing date), and *Three-year sales growth* (geometric growth in net sales over the three years prior to the year of the proxy mailing). *Insider ownership* is the proportion of equity held by corporate insiders, and *Firm size* is logged total assets in the year of the proxy mailing converted to constant 2000 dollars. Finally, we include *Unionization rate* to test if union presence at targeted firms provides a motivation for union activism.

Descriptive statistics for the variables used in Eq. (1) are provided in Table 3, Panel A. After deleting missing observations, there are a total of 101 *Union target* events and 4106 non-targeted controls for a total of 4207 observations from which the propensity scores are estimated. Panel B provides logit coefficient estimates using Eq. (1). As expected, *One-year stock return* and *Three-year sales growth* are negative and significant at the 5% (1%) levels, respectively. Higher *Institutional equity percent* is positively related to the likelihood of targeting albeit not at conventional significance levels. *Firm size* is positively related to targeting as expected at the 1% level, and *Unionization rate* is statistically unrelated. The latter result is consistent with Ertimur et al. (2011) finding that union activism is not necessarily focused on highly unionized firms.[12]

In Table 3 Panel C1, we employ the nearest-neighbor approach with replacement, using a caliper of 0.1, which identifies a single

[10] We thank our anonymous referee for this comment.

[11] Following Güntay et al. (2004), we also exclude issuer names that include the words “Acquisition”, “Capital”, “Financial”, “Finance”, “Funding”, “Leasing”, and “Security.”

[12] We test if the balancing property holds, i.e. if the treatment and comparison observations have identical mean propensity scores within blocks of the propensity score. The data is balanced if the means of each of the explanatory variables are equal within each of these blocks. Eq. (1) satisfies the balancing property after splitting the sample into eight equally spaced intervals of the propensity score. The balancing properties are obtained with the Stata routine PTEST, written by Edwin Leuven and Barbara Sianesi.

Table 3
Impact of union shareholder activism on yield spread: propensity score matching.

Panel A: summary statistics for variables used in logit model					
	Mean	**St dev.**	**25th Quartile**	**Median**	**75th Quartile**
Union target	0.024	0.153	0	0	0
One-year stock return	0.202	0.370	−0.009	0.170	0.376
Three-year sales growth	0.149	0.369	0.029	0.080	0.167
Insider equity	0.094	0.170	0.004	0.017	0.106
Institutional equity	0.625	0.357	0.502	0.642	0.789
Log (total assets)	8.694	1.441	7.703	8.705	9.747
Unionization rate	0.094	0.138	0.000	0.026	0.145
No. obs.	4207				

Panel B: logit model coefficient estimates	
Constant	−9.885***
	(0.000)
One-year stock return	−0.919**
	(0.010)
Three-year sales growth	−2.727***
	(0.002)
Insider equity percent	0.282
	(0.699)
Institutional equity percent	0.274
	(0.107)
Firm size	0.617***
	(0.000)
Unionization rate	−1.371
	(0.124)
No. obs.	4207
Pseudo R-squared	0.112
LR chi2	106.81
(p-value)	(0.000)
Balancing property satisfied?	Yes

Panel C: average treatment effect on the treated, nearest-neighbor matching							
	Treated	**No. obs.**	**Controls**	**No. obs.**	**ATET**	**Standard error**	**T-stat**
Panel C1: one-to-one matching							
Full sample	0.0125	101	0.0109	101	0.0016	0.0019	0.45
Inst. own. > median	0.0135	73	0.0229	73	−0.0094	0.0046	−2.04
Inst. own. ≤ median	0.0099	28	0.0122	28	−0.0022	0.0045	−0.51
Panel C2: one-to-five matching							
Full sample	0.0125	101	0.0128	505	−0.0003	0.0019	−0.14
Inst. own. > median	0.0135	73	0.0196	365	−0.0061	0.0026	−2.38
Inst. own. ≤ median	0.0135	28	0.0137	140	−0.0002	0.0044	−0.51

Table 3, Panel A provides descriptive statistics for variables used in the logit model, and Panel B provides coefficient estimates. The primary sample is based on 101 targeted observations and 4106 untargeted observations with a complete set of non-missing covariates. Panel C provides ATETs using one-to-one and one-to-five nearest-neighbor matching. Targeted firms are issuers whose offering date is within a three-month window following the proxy mailing date. Total assets are converted to constant 2000 dollars. Additional variable details are provided in the Appendix.

match for each treated firm according to the closest propensity score. Matching firms are identified using firm level characteristics following the specification of Eq. (1), however this analysis is conducted at the bond level. Because some targeted firms have more than one issue occurring in the three-month period following the proxy mailing date, the above- and below median institutional ownership subsets are imbalanced. For robustness, we repeat the analysis using five matched firms for each treated firms and present results in Panel C2. The key result of PSM is the average treatment effect on the treated (ATET), which is the average difference in yield spreads between treated and the propensity score-matched control firms. In Panel B1 using one-to-one matching, the average yield spread of the matched control sample insignificantly different from the treated (matched) sample. We conduct the PSM procedure on subsamples, using above- and below-median proportion of equity held by institutions. The two subsets are imbalanced, reflecting the intuition that firms with higher institutional ownership are larger, and larger firms are more frequent bond issuers. The subset of treated firms comprising the top half is associated with an average yield spread of 0.0135, while the matched sample is associated with an average of 0.0229. The difference in yield spread is −94 basis points which is significantly different from zero based on the z-statistic of −2.04. In contrast, the difference between the subset of treated firms comprising the bottom half and its matched sample is about 22 basis points which is not statistically different from zero. Panel C2 provides qualitatively similar results using one-to-five matching. Overall, these results provide preliminary support for Hypothesis *H3*.

5.2. Regression analysis of at-issue yield spreads

The preceding PSM analysis provides evidence that union targeting causally affects yield spreads, particularly among firms where institutional equity ownership is higher. We proceed to a multivariate regression setting which controls for additional bond- and firm-level characteristics likely to be associated with at-issue yield spreads. As suggested above, we surmise that one channel by which targeting affects spread is through the information component of spread. Therefore, it is important to control for default and liquidity risk which comprise the remaining components of spread. As with the prior analyses, the cross-sectional yield spread analysis spans 1989–2010. We regress at-issue yield spreads on the *Union* *target* binary variable and independent variables that control for bond- and firm-level characteristics likely to be related to the risk premium on corporate bonds. The cross-sectional model is specified as follows:

$$\begin{aligned} Yield\,spread = {} & \alpha_0 + \alpha_1 Union\ target + \alpha_2 Institututional\,ownership \\ & + \alpha_3 Z - Score\ Dummy + \alpha_4 Unionization\ rate \\ & + \alpha_5 Residual\ bond\ rating + \alpha_6 Baa - Aaa\ spread \\ & + \alpha_7 Callable + \alpha_8 Putable + \alpha_9 Subordinate \\ & + \alpha_{10} Time - to - maturity + \alpha_{11} Coupon\ rate \\ & + \alpha_{12} Issue\,amount + \alpha_{13} Analyst\ forecast\ error \\ & + \alpha_{14} Analyst\ forecast\ dispersion \\ & + \alpha_{15} No.estimates + \alpha_{16} Financial\ leverage \\ & + \alpha_{17} Market - book\,ratio + \alpha_{18} ROA + \alpha_{19} Std.ROA \\ & + \alpha_{20} Firm\ size + Fama \\ & - French\,30\,industry\ fixed\ effects \\ & + Year\ fixed\,effects \end{aligned} \tag{2}$$

We include a series of variables that we interact with the *Union target* indicator variable. Following the prediction of Hypothesis *H2, Union target* × *Institutional ownership* tests if higher equity ownership by institutions elevates the activism effect on yield spreads. Ertimur et al. (2010) provide evidence that proposals are more likely to be implemented when the sponsor is a labor union. Therefore, it is possible that the impact of activism on yield spreads is directly related to reduced default risk to the extent implementation affects investor perceptions of default. We investigate this conjecture by interacting *Union target* with *Z-score dummy*: a significant negative coefficient would support the view that targeting has a direct incremental effect on target firms that are closer to bankruptcy. *Z-score dummy* is a direct measure of proximity to bankruptcy and is based on the formulation described by Altman (1968). We create a binary variable equal to one if *z*-score is less than 1.81, which indicates a high likelihood of financial distress. Finally, we examine if the union activism effect on yield spread varies with union intensity at targeted firms using *Unionization rate*.

Credit ratings are determined by variables that are also used to explain yield spreads; thus, in order to discern the impact these variables have on yield spreads independent of their effect on credit ratings, we follow Mansi et al. (2004) by creating a *Residual bond rating* variable that is purged of the information contained in the bond- and firm-specific control variables. *Residual bond rating* is the residual of a regression of Moody's bond ratings (converted to numerical equivalents ranging from 1 ("C") to 21 ("Aaa") on the right side variables specified in Eq. (1) and provides an overall measure of default risk independent of the direct effects the additional control variables may have on bond ratings. The additional variables are drawn from a large body of work on the determinants of yield spreads (e.g. Mansi et al., 2011; Ortiz-Molina, 2006; Klock, Mansi and Maxwell, 2005; Bhojraj and Sengupta, 2003). Exposure to systematic economic risk is captured by *Baa-Aaa spread. Callable* and *Putable* control for embedded call and put options, respectively. *Subordinate* is an indicator variable equal to one if the bond is subordinate to other debt issues. *Time to maturity* controls for the effects of bond term on yield spread and *Coupon rate* controls for positive coupon effects documented in prior work (e.g. Campbell and Taksler, 2003). *Issue amount* (converted to constant 2000 dollars) controls for liquidity: Larger issues are associated with economies of scale in underwriting and reduction in liquidity risk (Bhojraj and Sengupta, 2003).

Turning to the firm-level explanatory variables, we control for the extent of information asymmetry between corporate insiders and outside investors alternatively with *Analyst forecast dispersion, Analyst forecast error,* and *No. analyst estimates* (Mansi et al., 2011). *Financial leverage* measures default risk and *Market-book ratio* measures cash flow growth opportunities. Profitability is measured with *ROA*, while *Std. (ROA)* measures cash flow risk. Finally, *Firm size* (converted to constant 2000 dollars) is an alternative measure of liquidity as larger issuers are more likely to be known to market participants and therefore more likely to be heavily traded by institutional investors. We control for unobservable effects related to industry and time by including Fama–French 30 industry and year indicator variables.

Appendix B provides summary statistics of the bond- and firm-level variables and Table 4 provides regression coefficient estimates for Eq. (2). In Models (1) and (2), we provide coefficient estimates using a matched set of untargeted control firms following the one-to-five matching procedure employed in the PSM analysis above: Model (1) provides estimates for the *Union target* and *Institutional ownership* main effects, and Model (2) includes the *Union target* × *Institutional ownership* interaction to test if the effect of targeting on yield spread varies according to the level of institutional equity ownership. The *p*-values are based on robust standard errors that are clustered at the industry level. Consistent with the premise that the activism process reduces asymmetry among outside investor groups, Model (1) shows that *Union target* main effect is negatively, albeit insignificantly, associated with yield spreads. Model (2) includes the interaction *Union target* × *Institutional ownership*. Consistent with Hypothesis *H3* and our earlier results, the interaction coefficient estimate is negative and significant at the 5% level indicating that the marginal effect of targeting on yield spreads is increasing in higher levels of institutional ownership.

Because the number of observations used in Models (1) and (2) is relatively small, we re-estimate Eq. (2) using the full sample of industrial bond issues from the SDC dataset. This results in a sample size of 3716 observations with a complete set of non-missing control variables. In Model (3), the *Union target* main effect is negative and significantly different from zero at the 5% level, indicating that targeting is associated with systematically lower at-issue spreads. Similar to the matched sample findings in Model (2), Model (4) illustrates the *Union target* × *Institutional ownership* interaction is negative and significant at the 5% level demonstrating that the effect of targeting on yield spreads is stronger when institutional ownership is higher across the broader sample. In Models (5) and (6), we test if the *Union target* effect on yield spreads is related to proximity to default or to unionization intensity at targeted firms with the interactions *Union target* × *Z-score dummy* (Model 4) and *Union target* × *Unionization rate* (Model 5). These interaction terms are insignificantly different from zero, providing additional evidence that the union activism effect on yield spreads is related to information flow between large outside investors. The signs and significance of the remaining bond- and firm-level control variables are largely as expected.[13]

6. Effect of union activism on stock price informativeness

6.1. Methodology

The empirical results in previous sections demonstrate that activism is associated with greater trading activity and reduced information asymmetry between different groups of investors over the initial time period following the mailing of proxies containing union-sponsored proposals. These findings imply that stock prices impound private information through greater institutional trading, thereby becoming more informative, at least in the shorter run. We

[13] As an alternative approach to deal with sample selection bias, we perform a Heckman-type (1979) treatment effect model and obtain similar results. For brevity, we do not report the results here but they are available upon request.

Table 4
Regressions of at-issue yield spreads on union targeting, interactions with institutional equity ownership, and other control variables.

	Matched control firms			Unrestricted control firms		
	Model (1)	**Model (2)**	**Model (3)**	**Model (4)**	**Model (5)**	**Model (6)**
Union target	−0.0017	0.0063	−0.0019**	0.0019	−0.0014	−0.0021**
	(0.365)	(0.193)	(0.044)	(0.330)	(0.166)	(0.042)
Institutional ownership	0.0021	0.0047	−0.0004	−0.0004	−0.0004	−0.0004
	(0.643)	(0.283)	(0.333)	(0.392)	(0.348)	(0.334)
Union target × institutional ownership		−0.0119**		−0.0056**		
		(0.047)		(0.038)		
Z-score dummy	0.0041**	0.0042**	0.0024***	0.0024***	0.0024***	0.0024***
	(0.040)	(0.040)	(0.000)	(0.000)	(0.000)	(0.000)
Union target × Z-score dummy					−0.0012	
					(0.538)	
Unionization rate	0.0020	0.0026	0.0008	0.0008	0.0007	0.0007
	(0.524)	(0.416)	(0.279)	(0.265)	(0.315)	(0.348)
Union target × unionization rate						0.0026
						(0.398)
Residual bond rating	−0.0008**	−0.0008**	−0.0012***	−0.0012***	−0.0012***	−0.0012***
	(0.028)	(0.025)	(0.000)	(0.000)	(0.000)	(0.000)
Baa- Aaa spread	0.6736	0.8132*	0.8160***	0.8194***	0.8157***	0.8161***
	(0.172)	(0.053)	(0.000)	(0.000)	(0.000)	(0.000)
Callable	0.0004	0.0006	0.0018***	0.0019***	0.0018***	0.0018***
	(0.886)	(0.834)	(0.000)	(0.000)	(0.000)	(0.000)
Putable	−0.0029	−0.0034	−0.0015*	−0.0015*	−0.0015*	−0.0015*
	(0.202)	(0.122)	(0.053)	(0.053)	(0.053)	(0.054)
Subordinate	0.0058**	0.0060**	0.0016***	0.0016***	0.0016***	0.0016***
	(0.029)	(0.016)	(0.001)	(0.002)	(0.001)	(0.001)
Time to maturity	−0.0001	−0.0001	−0.0002***	−0.0002***	−0.0002***	−0.0002***
	(0.545)	(0.602)	(0.000)	(0.000)	(0.000)	(0.000)
Coupon rate	0.4432*	0.4395*	0.6343***	0.6338***	0.6342***	0.6344***
	(0.063)	(0.063)	(0.000)	(0.000)	(0.000)	(0.000)
Issue amount	−0.0002	−0.0002	0.0001	0.0001	0.0001	0.0001
	(0.682)	(0.691)	(0.228)	(0.260)	(0.237)	(0.223)
Analyst forecast error	0.0641	0.0773	0.0039**	0.0039**	0.0039**	0.0039**
	(0.399)	(0.247)	(0.032)	(0.032)	(0.032)	(0.032)
Analyst forecast dispersion	0.8462***	0.9013***	−0.0010	−0.0010	−0.0011	−0.0011
	(0.002)	(0.000)	(0.744)	(0.744)	(0.741)	(0.744)
No. analyst estimates	−0.0024	−0.0027	−0.0015***	−0.0014***	−0.0015***	−0.0015***
	(0.196)	(0.115)	(0.000)	(0.000)	(0.000)	(0.000)
Financial leverage	0.0091	0.0103	0.0037***	0.0037***	0.0037***	0.0037***
	(0.141)	(0.110)	(0.006)	(0.006)	(0.006)	(0.006)
Market-book ratio	0.0006	−0.0001	−0.0001	−0.0001	−0.0001	−0.0001
	(0.749)	(0.948)	(0.848)	(0.840)	(0.851)	(0.846)
ROA	0.0158	0.0247	−0.0123**	−0.0123**	−0.0123**	−0.0123**
	(0.456)	(0.264)	(0.030)	(0.031)	(0.029)	(0.030)
Std. (ROA)	0.0224	0.0271	0.0189***	0.0189***	0.0189***	0.0189***
	(0.609)	(0.575)	(0.000)	(0.000)	(0.000)	(0.000)
Firm size	−0.0002	−0.0002	−0.0011***	−0.0011***	−0.0011***	−0.0011***
	(0.736)	(0.725)	(0.000)	(0.000)	(0.000)	(0.000)
Fama–French 30 Industry fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
No. obs.	567	567	3716	3716	3716	3716
R-squared	0.948	0.949	0.886	0.886	0.886	0.886
F-statistic	59.85	62.59	750.8	755.8	810.1	771.5

Table 4 presents the coefficient estimates of yield spread regressed on union targeting indicators and other control variables. Models (1) and (2) present results using sample and matched firms using nearest-neighbor one-to-five matching based on propensity scores drawn from Eq. (1). Models (3)–(6) present results using the full sample of at-issue yield spreads. Variable descriptions are provided in Appendix A. *P*-values are provided in parentheses and are based on standard errors that are clustered at the industry level. Total assets and issue size are converted to constant 2000 dollars. ***, **, and * correspond to significance at the 1, 5, and 10% levels, respectively.

further test if stock prices become more informative as a result of shareholder activism using a more direct measure of stock price informativeness. We use a measure suggested by Llorente et al. (2002, LMSW hereafter) which is based on stock return autocorrelation conditional on trading volume as a proxy for stock price informativeness.[14] This measure is constructed from the following regression estimated for each firm-year:

$$r_{j,t} = \alpha_j + \gamma_j r_{j,t-1} + \Theta_y r_{j,t-1} V_{j,t-1} + e_{j,t}, \tag{3}$$

where $r_{j,t}$ and $r_{j,t-1}$ are contemporaneous and lagged weekly stock returns, respectively, and $V_{j,t-1}$ is lagged log turnover detrended by subtracting the moving average of logged turnover over the prior 26 weeks. The key estimate is the coefficient of the interaction term Θ_y, which reflects the amount of information-based trading. As Fernandes and Ferreira (2008) explain, the intuition is that stocks with a high degree of information-based trading will exhibit positive return autocorrelation (i.e., higher values of Θ) in periods when volume is higher. Accordingly, we surmise that if activism results in more informed trading, Θ should increase for targeted firms and the effect should be pronounced for the subset of firms with higher institutional ownership.

We assess the impact of activism with difference-in-differences (DiD) analyses. Using propensity scores estimated from Eq. (1), in

[14] We thank our anonymous referee for the suggestions to use difference-in-differences analyses and the Llorente et al (2002) information based trading measure.

each year we identify five matched firms for each sample firm using the intersection of the Compustat, CRSP, Compact Disclosure / GMI Ratings, and Unionstats.com datasets. Using each treated (i.e., targeted) firm and the five control firms, we construct a two-year panel comprised of one year prior to- and following the year of the union-sponsored proposal for each targeted firm. We create a *Treated* dummy variable equal to one for treated firms and zero for matched firms. To test if the effect of activism on information-based trading (and thus stock price informativeness) is prevalent, for each treated and control firm we create a *Post* dummy variable equal to one (zero) if the year is one year following (one year prior to) the year of targeting, We test the impact of targeting on the outcome variables by interacting *Treated* with *Post* using the following regression model:

$$Outcome\,variable = \alpha_0 + \alpha_1 Treated + \alpha_2 Post + \alpha_3 Treated \times Post + Controls + e_{j,t} \qquad (4)$$

6.2. Results

In addition to *Treated, Post*, and *Treated*×*Post*, we include industry- and year fixed effects. The average value of the outcome variable for the matched firms during the pre-activism period is obtained from the intercept term (α_0). The corresponding value for matched firms in the 'post' period is the sum of regression coefficients $\alpha_0 + \alpha_2$. Similarly, the average value of the variable of interest for union-targeted (treated) firm during the 'pre' period is $\alpha_0 + \alpha_1$. The corresponding value for treated firms in the 'post' period is the sum of all four parameter estimates $\alpha_0 - \alpha_3$. The net difference in the dependent variable for treated firms in the 'post' period relative to matched firms is the regression coefficient α_3. We focus on alternative outcome variables including institutional equity ownership percentage, the LMSW Θ measure of informed-based trading, and analyst-based variables including forecast dispersion, forecast error, and number of analyst estimates.

First, we examine if the proportion of institutional equity ownership increases subsequent to union targeting. Following Hypothesis *H1*, if institutional ownership is the main channel through which information flows between labor unions and other groups of institutional investors, there should be an increase in institutional ownership among union-targeted firms. The results in Table 5 Panel A support this view as there is a statistically significant increase in the average level of institutional ownership between pre- and post-activism periods. On average, the percentage of institutional ownership of union-targeted firms increases by 0.36% across the year prior to and following the year of activism. In contrast, the matched set of firms experience an average decrease of 1.27%. This results in a net change of 1.64% in institutional ownership for union-targeted firms relative to their peer firms, and is statistically significant at the 10% level. A potential explanation for this 'flight-to-quality' effect is that institutions may be attracted to firms with activist union shareholders because they foresee activism leading to relative improvements in the diversity of information surrounding the firm, from which all shareholders will benefit (e.g., Goldstein and Yang, 2015).[15]

We now turn to our primary research question of whether stock prices become more informative as a result of union activism. According to Llorente et al. (2002), an increase in Θ indicates more information-based trading and thus more informative stock prices. As reported in Panel B of Table 5, the Θ measure indicates that stock prices of union-targeted firms are generally less informative than those of their matched counterparts prior to the year of targeting (−0.0451 vs. 0.0119). However, in the year following targeting, Θ increases significantly from −0.0451 to 0.0178 for firms that received union-sponsored proposals while the matched firms experience a decrease from 0.0119 to −0.0045. The net change is 0.0794 which is statistically significant at the 1% level. This provides strong evidence in further support of *H1*.

The focus of our study is on the improvement of stock price informativeness based on increased collaboration between different groups of investors; improvements in firms' information environment can also be the result of improvement in firms' transparency (e.g. improved financial disclosure). As a result of improved dissemination of managerial private information to stock market participants, the quality of firms' public information improves thereby leading to better earnings predictability and greater coverage by equity analysts. To ensure that the increase in Θ documented in Panel B does not merely reflect a reduction in the informational mismatch between managers and outsiders, we examine relative changes in analyst-based measures (e.g. Diether et al., 2002).[16] Panels C–E report univariate DiD results using analyst forecast dispersion, analyst forecast accuracy, and the logged number of analyst estimates, respectively. We do not detect statistically significant changes in any of these measures, thereby providing evidence that the increase in Θ is not driven by improvements in transparency following union activism.[17]

Table 5
Univariate difference-in-differences analysis.

	Treated	Match	Difference, treated-match
Panel A: institutional equity ownership			
Pre	0.7228	0.7594	−0.0366
Post	0.7265	0.7467	−0.0202
Change	0.0036	−0.0127	0.0164*
Panel B: LMSW measure			
Pre	−0.0451	0.0119	−0.0570
Post	0.0178	−0.0045	0.0224
Change	0.0630	−0.0164	0.0794***
Panel C: analyst forecast dispersion			
Pre	0.3013	1.0184	−0.7171
Post	−0.3570	0.7437	−1.1007
Change	−0.6583	−0.2747	−0.3836
Panel D: analyst forecast error			
Pre	−0.1606	−0.7079	0.5473
Post	0.2148	−0.4617	0.6765
Change	0.3754	0.2462	0.1292
Panel E: no. analyst estimates			
Pre	2.7582	2.6106	0.1476
Post	2.7615	2.6229	0.1386
Change	0.0033	0.0123	−0.0090

Table 5 reports difference-in-differences analysis on alternative outcome variables using a matched sample comprised of five untargeted firms for each targeted (i.e. treated) firm). The matched sample is identified using propensity scores drawn from Eq. (1). *Post* is defined as the year following the year of targeting, and *Pre* is the year prior to targeting. The change from *Pre* to *Post* for treated and matched firms, and the differences between treated and matched firms, are based on the regression parameters specified by Eq. (4). The LMSW measure is stock return autocorrelation conditional on trading volume as defined by Eq. (1). The other variables are defined in Appendix A. ***, **, and * correspond to significance at the 1, 5, and 10% levels, respectively.

[15] While we do not test this explanation directly, an interesting question left for future research is whether this effect is stronger among passive institutional investors (i.e. non-active, as defined by Ferreira and Matos, 2008) such as banks, insurance companies, and pension funds. We thank our anonymous referee for this additional insight.

[16] If firms become more transparent following union activism we would expect to see an increase in analyst coverage and forecast accuracy and a reduction in forecast dispersion. However, our results are not consistent with this. Rather, they relate to the aspect of information asymmetry which stems from differential information between outsiders. We are grateful to an anonymous referee for highlighting this.

[17] We also conduct a two-step Heckman (1979) model on the dependent variables as an alternative way to deal with sample selection bias. The untabulated results are qualitatively similar to those reported here.

Table 6
Difference-in-differences regression analysis.

	Treated	Post	Treated × Post	Adj. R^2	No. obs.
Panel A: institutional ownership					
Full sample	−0.0098	−0.0052	0.0178**	0.2690	6628
	(0.356)	(0.348)	(0.049)		
Institutional ownership > median	−0.0400**	−0.0106	0.0268**	0.4930	3281
	(0.011)	(0.151)	(0.050)		
Institutional ownership ≤ median	0.0127	−0.0132**	0.0259**	0.3040	3347
	(0.362)	(0.019)	(0.028)		
Panel B: LMSW measure					
Full sample	−0.0605***	−0.0169	0.0811***	0.1340	6934
	(0.001)	(0.152)	(0.002)		
Institutional ownership > median	−0.0592*	−0.0054	0.0994**	0.1370	3281
	(0.052)	(0.752)	(0.016)		
Institutional ownership ≤ median	−0.0547**	−0.0130	0.0373	0.1900	3347
	(0.049)	(0.445)	(0.318)		
Panel C: analyst forecast dispersion					
Full sample	−0.6387	−0.3456	−0.3041	0.0540	6772
	(0.329)	(0.356)	(0.645)		
Institutional ownership > median	−1.0750	−1.2579	−0.7379	0.1190	3225
	(0.325)	(0.264)	(0.761)		
Institutional ownership ≤ median	0.0079	0.0619	−0.2926	0.0710	3277
	(0.772)	(0.314)	(0.335)		
Panel D: analyst forecast error					
Full sample	0.4448	0.2763	0.0894	0.0670	6820
	(0.279)	(0.334)	(0.723)		
Institutional ownership > median	1.0960	0.6587	−0.5621	0.0860	3234
	(0.379)	(0.295)	(0.418)		
Institutional ownership ≤ median	0.0729	−0.1447	0.8626	0.0730	3316
	(0.684)	(0.472)	(0.411)		
Panel E : no. analyst estimates					
Full sample	0.0762***	0.0237*	−0.0159	0.4780	693
	(0.008)	(0.057)	(0.513)		
Institutional ownership > median	0.0492	0.0201	−0.0438	0.4410	328
	(0.196)	(0.248)	(0.285)		
Institutional ownership ≤ median	0.0291	−0.0258	0.0778*	0.5450	3347
	(0.508)	(0.213)	(0.069)		

Table 6 reports difference-in-difference regression analysis. The LMSW measure is stock return autocorrelation conditional on trading volume. The difference-in-differences analysis is based on five matched firms for each sample firm. We include firm size (logged total assets converted to constant year 2000 dollars), financial leverage, future growth opportunities (market-book ratio), and profitability (ROA) as control variables. Each regression model also includes firm level industry and year fixed effects. The standard errors are clustered at the firm level. The ***, **, and * correspond to significance at the 1, 5, and 10% levels, respectively.

We conduct a series of robustness tests for our primary Θ DiD results. First, to ensure that our results are not dependent on the choice of comparison time periods, we examine four alternative 'pre' and 'post' time period specifications. These results are presented in Appendix C, Panels A1–A3. In the first specification (A1), we construct a panel of four years for each firm where *Post* is equal to one (zero) if the year is two years following (two years prior to) the year of targeting. In the second specification (A2), each panel is comprised of three years where *Pre* is the year prior- and year of activism and *Post* is the year following. In Panel A3, we extend the panel for each firm to four years where *Pre* is the year prior- and the year of activism and *Post* is the two years following the activism year. In each case, the Θ measure in the *Post* period continues to be positive and highly statistically significant.

In our second robustness test, it is possible that the documented increase in information-based trading following the union activism maybe caused by simultaneous public pension fund activism. Public employee pension plans are among the largest institutional investors in the marketplace and, like labor unions, have a lengthy track record of shareholder proposal activism (e.g. Del Guercio and Woidtke, 2014). Using the GMI dataset which contains all shareholder proposals over the 2003–2010 period, we exclude from our dataset all union proposal observations that coincide with public fund-sponsored corporate governance proposals repeat the DiD analysis. We present the results for this subset in Appendix C Panels B1–B4. Panel B1 reports results using the two-year panel for each firm where POST equals one for the year following the activism year and zero otherwise, and Panels B2-B4 provide results for the alternative *Post* specifications used in Panels A1–A3. Collectively, these results suggest that our primary DiD results are not driven by activism by other prominent institutional investors.[18]

Third, the preceding discussion implies that there should be distinctive differences in information effects for shareholder proposals sponsored by proponents that do not have access to institutional communication networks. We examine proposals sponsored by individual 'corporate gadflies' as a benchmark for non-institutional comparison group.[19] Using the 2003–2010 Corporate Library dataset which includes all proposals targeting S&P 1500 firms, we search for proposals sponsored by Kenneth Steiner, Emil Rossi, John Chevedden, Evelyn Y. Davis, or Gerald Armstrong. We identify 535 proxy-year observations that contain proposals sponsored by only by unions or union-affiliated funds, and 185 proxies that only contain proposals sponsored by individual proponents. Following the DiD process discussed above, we examine the post-targeting DiD effect for these 185 firm-year observations. In untabulated results, the change in Θ is negative (−0.092) but insignificantly different from zero. These numbers are also insignificantly negative for both high and low institutional owner-

[18] We note that our empirical results do not rule out the possible informational impact of other institutional activists. Following Goldstein and Yang (2015), the complementarities in trading and thus the improvement in information production can improve as long as institutional outsiders with different information set interact.

[19] http://proxymonitor.org/forms/pmr_02.aspx/.

ship groups. Overall we do not find any evidence that individually-sponsored proposals are related to significant information effects.

Other explanatory variables may impact the outcome variables specified in Table 5. To ensure that our univariate DiD results in Table 5 are robust to the inclusion of these variables, we re-estimate Eq. (4) with additional firm-level explanatory variables that may be correlated with the outcome variables. These include: firm size (logged total assets converted to constant 2000 dollars), financial leverage, future growth opportunities (market-book ratio), and profitability (ROA). As in prior regressions, we also include industry- and year fixed effects. Following prior analyses, we present the DiD regression coefficients for the overall sample and for subsets bifurcated by institutional equity ownership. Viewed collectively, the results in Table 6 reflect the univariate findings presented in Table 5. Specifically, Panel A shows that the *Treated*×*Post* estimate is positive and statistically significant in the full sample using institutional equity percent as the dependent variable. In Panel B using the Θ measure as the outcome variable, *Treated*×*Post* is positive and significant at the 1% level for the full sample. Consistent with Hypothesis *H2*, this result is driven by the above-median institutional equity ownership subset where *Treated*×*Post* is significant at the 5% level ($p = 0.016$). In Panels C-E using the analyst measures as outcome measures, the general insignificance of *Treated*×*Post* suggests that targeting does not impact the extent of information asymmetry between corporate insiders and outside market participants.

7. Conclusion

We investigate if shareholder activism affects the information flow of targeted firms. Our focus is on information asymmetry which arises from differential information between outside investors. Viewed collectively, our evidence supports the view that information-based trading increases following labor union activism. There is no change in analyst coverage, or analyst forecast accuracy or dispersion, which suggests that firm transparency or the information mismatch between corporate insiders and outsiders does not change. The improvement in price informativeness is more pronounced in firms with more institutional ownership, which indicates that the flow of information between unions and institutional investors is an important aspect of the reduction in information asymmetry. Our findings are in line with the theoretical findings of Goldstein and Yang (2015), which points to the benefit of greater diversity of information among different groups of significant investors in the stock market in improving the overall amount of information revealed in stock prices. In their model, different groups of investors possess different sets of information about firms' fundamentals. When other groups of investors help reduce the uncertainty of information on aspects they do not possess, it will reduce their overall uncertainty. As a result, they are more motivated to trade and in the process, disseminate their own information. The overall stock market becomes more informative as a consequence due to synergies in information acquisition.

While our empirical findings establish that there is an increase in information flow and information based trading among outsiders upon the union-sponsored activism on average, we do not distinguish between the different level of potential knowledge and expertise that unions can offer in this process. In an extreme scenario, certain unions in particular industries may possess external private information that is not even available to corporate insiders. When these unions share this information during their activism, managers of peer firms might gain additional information (e.g. managerial learning in the context of Foucault and Fresard (2014)) through these revelations. While beyond the scope of this paper, these additional aspects are left for future research.

Appendix A. Description of variables used in the study

Variable name	Description and source
Panel A: variables used in trading volume analyses	
Institutional ownership	Percentage of equity held by institutional shareholders. *Sources*: Compact Disclosure (1989–2005); The Corporate Library (2006–2010)
Mean trading volume	Logged mean daily trading volume over the year of the proposal. *Source*: CRSP
Trading volume volatility	Logged standard deviation of daily trading volume over the year of the proposal *Source*: CRSP
CII member	Indicator variable if the union is a member of the Council of Institutional Investors in 2013 *Source*: http://www.cii.org/ciigeneralmembers
Unionization rate	Proportion of workers that are members of a labor union in the issuer's 3-digit SIC code in a given year, *Source*: Union Membership and Coverage Database (www.unionstats.com)
Firm size	Logged total assets, converted to constant year 2000 dollars. *Source*: Compustat
Panel B: additional variables used in PSM analysis	
Union target	Indicator variable equal to 1 if the issue occurs within a maximum of 90 days following the proxy mailing date. *Source*: Investor Responsibility Research Center (1989–2002); GMI Ratings (2003–2010)
One-year stock return	Cumulated stock return over the 225 days prior to the issue date, winsorized at the 1% tails. *Source*: CRSP
Three-year sales growth	Geometric growth in sales (SALE) over the three years prior to the year of the issue date. *Source*: Compustat
Insider ownership	Percentage of equity held by corporate insiders *Sources*: Compact Disclosure (1989–2005); The Corporate Library (2006–2010)
Panel C: additional variables used in at-issue yield spread analysis	
Panel C1: bond-level variables	
Yield to maturity	Yield-to-maturity (*YTM*) is calculated using inputs provided by SDC Platinum (time to maturity, coupon rate, and offer price).
Yield spread	The bond's YTM minus the interpolated monthly Treasury bond yield. Winsorized at the 1% tails. *Source*: SDC Platinum (bond prices), St. Louis Federal Reserve (Treasury Note and Bond yields)
Callable	Binary variable = 1 if the bond is callable (i.e. if Call Protection = 'Non-Call Life'.) *Source*: SDC Platinum
Putable	Binary variable = 1 if the bond is putable. *Source*: SDC Platinum
Subordinate	Binary variable =1 if the bond issue is subordinate or senior subordinate. *Source*: SDC Platinum
Time to maturity	Number of years to final maturity. *Source*: SDC Platinum
Coupon rate	Annual coupon payment per one dollar of par value. *Source* SDC Platinum
Issue amount	Logged global USD proceeds of the issue. Source: SDC Platinum
Panel C2: firm-level variables	
Analyst forecast error	Absolute value of the analyst forecast error (the actual EPS minus the median forecast deflated by the fiscal-year-end stock price) *Source*: IBES via Datastream
Analyst forecast dispersion	Standard deviation of the inter-analyst forecast divided by the fiscal-year-end stock price. *Source*: IBES via Datastream
No. analyst estimates	Number of analyst estimates for the issuer's stock. *Source*: IBES via Datastream
Financial leverage	Interest-bearing debt (sum of DLC and DLTT) divided by total assets (AT). *Source*: Compustat
Market-book ratio	Book value of assets net of book equity (AT – CEQ) plus market value of equity (PRCC_F * CSHO), divided by total assets (AT)
ROA	Income before extraordinary items (IB) divided by mean total assets (AT) for the current and prior year. *Source*: Compustat
Std (ROA)	Standard deviation of *ROA* for the prior 5 years. *Source*: Compustat
Z-score dummy	*Z*-score is calculated following Altman (1968). *Z-score dummy* equals one if *z*-score $<$ 1.81 and zero otherwise. *Source*: Compustat
Fama–French 30 industry dummies	Industry classifications based on 30 industry definitions. *Source*: Kenneth R French Data Library (http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html)

Appendix B. Summary statistics for variables used in yield spread regressions

The following table provides the summary statistics for the 3716 bond-year observations with a complete set of control variables, based on 784 unique firm issuers spread over 2093 firm-years. The variable descriptions are provided in Appendix A.

Panel A: pooled sample bond characteristics					
	Mean	St dev.	25th Quartile	Median	75th Quartile
Yield to maturity	0.0726	0.0182	0.0619	0.0701	0.0830
Yield spread	0.0182	0.0175	0.0068	0.0119	0.0255
Time to maturity (years)	11.5997	8.5777	6.0904	10.0110	10.8753
Coupon rate	0.0719	0.0187	0.0613	0.0700	0.0825
Proceeds amount ($MM)	452	701	132	249	473
No. obs.	3716				

Panel B: issue characteristics		
	No. issues	Proportion of sample
Embedded options		
Callable	1888	0.5081
Putable	65	0.0175
Security and collateral:		
Senior secured	62	0.0167
Senior	932	0.2508
Senior subordinate	293	0.0788
Subordinate	17	0.0046
Unclassified	2412	0.6491
Credit (Moody) rating:		
Aaa	29	0.0078
Aa1-Aa3	369	0.0993
A1-A3	1236	0.3326
Baa1-Baa3	1035	0.2785
Ba1-Ba3	464	0.1249
B1-B3	532	0.1432
Caa1-Ca	51	0.0137
High yield	1047	0.2817

Panel C: pooled issuer characteristics					
	Mean	Std. dev.	25th Quartile	Median	75th Quartile
Panel C1: Firm-level control variables					
Institutional equity ownership	0.6562	0.3925	0.5196	0.6697	0.8080
Unionization rate	0.1539	0.1514	0.0370	0.0090	0.2240
Z-score dummy	0.3426	0.4747	0	0	1
Analyst forecast error	0.0157	0.2000	0.0003	0.0010	0.0036
Analyst forecast dispersion	0.0068	0.0636	0.0004	0.0010	0.0027
Number of analyst estimates	14.5069	8.5885	8	13	20
Financial leverage	0.3684	0.1853	0.2456	0.3386	0.4499
Market-book ratio	1.7169	0.8132	1.1911	1.4518	1.9440
ROA	0.0422	0.0794	0.0179	0.0467	0.0780
Std. ROA	0.0477	0.0536	0.0206	0.0323	0.0548
Total assets ($MM)	11,294	23,870	1,4712	4341	11,138
No. obs.	2093				

Appendix C. DiD tests of robustness using the LMSW measure as the outcome variable

Appendix C reports difference-in-differences analysis for the LMSW measure using alternative comparison time periods defining the year of targeting as year-zero. ***, **, and * correspond to significance at the 1, 5, and 10% levels, respectively.

	Treated	Match	Difference, treated-match
Panel A: alternative comparison periods			
Panel A1: PRE = (−2, −1), POST = (1, 2)			
Pre	−0.0289	0.0067	−0.0356
Post	0.0148	−0.0021	0.0169
Change	0.0437	−0.0088	0.0525***
Panel A2: PRE = (−1, 0), POST = (1)			
Pre	−0.0300	0.0087	−0.0387
Post	0.0146	−0.0025	0.0171
Change	0.0446	−0.0111	0.0558***
Panel A3: PRE = (−1, 0), POST = (1, 2)			
Pre	−0.0251	0.0123	−0.0374
Post	0.0194	0.0040	0.0154
Change	0.0445	−0.0083	0.0528**
Panel B: alternative comparison periods and no public pension funds			
Panel B1: PRE = (−1), POST = (1)			
Pre	−0.0377	0.0289	−0.0667
Post	0.0416	−0.0028	0.0388
Change	0.0793	−0.0262	0.1055***
Panel B2: PRE = (−2, −1), POST = (1, 2)			
Pre	−0.0203	0.0173	−0.0376
Post	0.0320	−0.0206	0.0526
Change	0.0523	−0.0379	0.0902***
Panel B3: PRE = (−1, 0), POST = (1)			
Pre	−0.0155	0.0209	−0.0364
Post	0.0450	0.0041	0.0409
Change	0.0605	−0.0168	0.0773***
Panel B4: PRE = (−1, 0), POST = (1, 2)			
Pre	−0.0180	0.0165	−0.0345
Post	0.0356	−0.0142	0.0498
Change	0.0536	−0.0307	0.0843**

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