The impact of labor unionization on CSR reporting

Abstract

Corporate social responsibility (CSR) reporting has been theorized as a key communication device and an integral part of a broader stakeholder integration management strategy. This paper examines the relationship between CSR disclosures and organized labor, an important internal stakeholder, whose institutional role in dynamically advancing employee interests creates opportunities and challenges for strategic management and firm sustainability. By employing a sample of 2,526 US firm-year observations for the period 2002–2015, we demonstrate that managers in unionized contexts are more likely to issue CSR reports than managers in firms where labor is not organized. Considering stakeholder theory, we argue that, in unionized contexts, managers more intensively resort to CSR disclosures to align interests, develop collaborative bonds with unions and smoothen relationships with external financial stakeholders. This effect is more prominent in areas where the geographical concentration of organized labor and the prevailing political ideology facilitate the role of unions. We offer implications for managers, labor unionists and market participants.

Keywords: Corporate social responsibility, sustainability reporting, stakeholder theory, unionization.

1 Introduction

Corporate social responsibility (CSR) reporting constitutes a discretionary business practice that has gained great momentum in recent decades (Arena et al., 2018, Chantziaras et al., 2020a, Chong and Rahman, 2020, Krasodomska and Cho, 2017). CSR reporting landscape has changed significantly during the last decades (Dobbs and van Staden, 2016, Vishwanathan et al., 2020), evolving from the mere issuance of environmental reports, as of early 80's, to the recognition and implementation of CSR reporting in the early 00's (Latapí Agudelo et al., 2019). The KPMG (2017) survey of corporate social responsibility reporting shows that more than 75% of the world's largest firms now issue CSR reports. According to Corporate Register[1], US firms published 909 standalone CSR reports in 2018 compared to just 170 in 2002. Literature has extensively examined the factors that affect the propensity of companies to initiate voluntary standalone CSR reports (see Dobbs and van Staden, 2016, Griffin and Sun, 2018, Grougiou et al., 2016, Hinze and Sump, 2019). It has been substantiated that CSR reporting is, *inter alia*, driven by salient stakeholders' expectations and pressures and constitutes an effective mechanism for managing the diverse and often contradictory stakeholder demands (Amini et al., 2018, Garriga and Melé, 2004, Lee et al., 2013, Romero et al., 2019).

Prior literature has drawn attention to the relationship between CSR communication initiatives and employees as a prominent internal stakeholder whose current concerns and future needs are important for corporate sustainability (Brunton et al., 2017, Wolf, 2013). It is, however, surprising that very limited academic attention has been directed towards understanding whether the intensity of issuing CSR disclosures is affected by the organization of employees in unions. Such an examination is required since the institutionalized role of

[1] Corporate Register (CorporateRegister.com) is the world's largest online directory of corporate responsibility (CR) reports.

unions in advancing employee claims, often by employing dynamic means, such as lobbying, industrial action and activism (Mitchell et al., 1997, Pendleton and Gospel, 2013, Preuss, 2008), can entail severe difficulties for management. Certain unions' demands and actions may contradict other stakeholders' interests, such as enhanced short-term profitability, corporate value and moderated cost of capital; thereby, undermining business sustainability (Chantziaras et al., 2020b, Doucouliagos and Laroche, 2009, Faleye et al., 2006).

In light of stakeholder theory (Garriga and Melé, 2004, Wolf, 2013), we investigate whether, in the presence of unions, organizations more frequently employ CSR reporting to facilitate an alignment of interests among salient stakeholders. To do so, we focus on a single geographical context, the United States, to obtain a homogenous sample in terms of the underlying institutional settings. The importance of the US setting resides in that it attracts great international attention for CSR practices and, additionally, represents a half-way house between countries where unionization has not been institutionalized and countries where the influence of unions in shaping corporate strategies is dominant.

Against this background, we employ a sample of 2,526 US firm-year observations for the estimation window 2002–2015. We show that managers in unionized contexts more intensively resort to CSR reporting than their counterparts in non-unionized corporations. Furthermore, we demonstrate that the propensity for CSR reporting in unionized firms increases significantly in states where the geographical concentration of organized labor and political ideology, i.e., the dominance of the Democratic Party, empowers the position and enhances the influence of unions.

Our contribution to the literature is twofold. First, we contribute to CSR literature by providing insights into the importance of the structure of industrial relations as a key driver of core stakeholder integration management mechanisms. Second, we contribute by shedding light on the role of institutional parameters, geographical concentration of organized labor and dominant political ideology, which either increase or limit unions' power to influence management's propensity for issuing CSR disclosures.

2 Background and Development of Hypotheses

2.1 Stakeholder theory and CSR reporting

CSR reporting has been approached through various theoretical perspectives (see Brooks and Oikonomou, 2018, Stubbs et al., 2013, Xu et al., 2019). We employ stakeholder theory as the most relevant theoretical background since it brings to the fore the role of salient internal and external stakeholders in shaping fundamental communications and integration management strategies (see Brooks and Oikonomou, 2018).

Central to stakeholder theory are the challenges managers encounter when they deal with important non-homogenous groups and individual stakeholders characterized by varying incentives, motivations, goals, risk preferences, power and influence (Chyz et al., 2013, Doh and Guay, 2006, Freeman, 1984) that can exert conflicting pressures on managers (Chen et al., 2011, Freeman, 1984, Mitchell et al., 1997, Reverte, 2009). Inevitably, managers prioritize the establishment of good relationships with prominent stakeholders and make systematic efforts to maintain reliable communication channels with them (Mitchell et al., 1997).

Previous studies have theorized stakeholder integration management as an important driver for CSR reporting (Pérez-López et al., 2015, Romero et al., 2019). The latter is seen as a key mechanism of corporate communication, capable of initiating dialogues with a wide spectrum of influential parties, including: employees, shareholders, investors, consumers, public authorities, environmental groups and NGOs (Doh and Guay, 2006, Reverte, 2009). Through CSR reports, managers signal out community, health and safety, diversity and other social and environmental concerns and ethical messages (Holder-Webb et al., 2009, Leventis et al., 2013); thus, addressing wider informational needs (Reverte, 2009) which are positively discounted by most prominent stakeholders (Akisik and Gal, 2017, Romero et al., 2019).

Interestingly, not all organizations employ such mechanisms due to several, non-mutually exclusive, reasons (Pedersen et al., 2013). Firstly, smaller firms and/or firms in less labor-intensive and visible industries may experience less diverse stakeholder demands, as well as limited sustained pressure regarding their social and environmental performance (Stubbs et al., 2013). Secondly, some non-reporting firms are likely to prioritize shareholder interests, placing predominant emphasis on economic growth, profitability and return on investment; thus, marginalizing broader societal and environmental concerns (Stubbs et al., 2013). Thirdly, issuing standalone CSR reports is both costly and time-consuming (Mahoney et al., 2013). Hence, adopting a cost-benefit analysis (Pérez-López et al., 2015), managers may resort to disclosing CSR only when salient stakeholder pressures create such complexities that the employment of costly but effective communication mechanisms is considered necessary to facilitate the alignment of conflicting interests (Mahoney et al., 2013).

2.2 Unionized firms and CSR reporting

Unions are institutionalized bodies established to advance employee interests. At the epicentre of the unions' role are employee working conditions, remuneration and other benefits, and job security (Kaufman, 2004). They often have a strong say in core corporate matters. For instance, about 47% of collective agreements involve some form of organized labor participation in committees and on corporate boards that advances labor's interests in relation to corporate welfare (Faleye et al., 2006).

Unions may co-operate productively with management and facilitate corporate operations and goals, when the interests of both are aligned i.e., unions can generate productivity-enhancing ideas, reduce cost associated with turnover, dissolve workplace disputes, shape employees' work perceptions, mobilize media and politicians to bail out unized firms (Cardador et al., 2017, Chen et al., 2011, Freeman, 1984, Kaufman, 2004, Vishwanathan et al., 2020, Xing et al., 2017).

However, in advancing employee claims, it is highly likely that unions resort to various tactics. Collective bargaining, lobbying, industrial action and activism (through strikes, work stoppages or deliberate decreases in production) are used to make the employee voice heard (Chyz et al., 2013, Faleye et al., 2006). Such actions undermine the collaborative and trust-based relationships with firm management since they can trigger corporate reputation loss, impact upon organizational profitability[2] and cause negative publicity to interested third parties, including investors and other capital market participants (Cardador et al., 2017, Chantziaras et al., 2020b, Chyz et al., 2013).

By disclosing CSR information, managers reduce information asymmetries (Bravo and Reguera-Alvarado, 2019). Managers may be reluctant to disclose sensitive financial information such as earnings forecasts (Kleiner and Bouillon, 1988) and good news about corporate performance (Chung et al., 2016), with a view to discouraging unions claiming additional rents and benefits (Cheng, 2017, Hilary, 2006, Kleiner and Bouillon, 1988). However, managers will be very much willing to disclose information on the organization's attitudes, motives, strategies, policies and actions (Flammer and Luo, 2017), since such information makes it possible for unions to identify and form mutual interests and ties with the organization (Boodoo, in press, Bravo and Reguera-Alvarado, 2019, Colombo et al., 2019, Glavas and Kelley, 2014, McKersie et al., 2008). Most unions acknowledge the broader

[2] For instance, see the case of the Hostess bankruptcy: http://goo.gl/Dx1C1r; http://goo.gl/umhBmM; http://goo.gl/cZYEqi; http://goo.gl/C9y4Pq (Accessed 26 September, 2020).

benefits of CSR reporting and have, therefore, argued for such disclosures to become compulsory in practice (Preuss, 2008).

However, unions may often be skeptical or negative against CSR reports. They may perceive CSR reporting as a marketing tool employed by firms to promote self-image and thereby avoid drastic reforms (Delmas and Burbano, 2011, Sobczak and Havard, 2015). Moreover, they may conceptualize such reports as a way to initiate a multi-stakeholder dialogue, which would inevitably undermine their power since they would then be treated as yet one more stakeholder among many (Sobczak and Havard, 2015). Unions' skepticism about CSR reporting may also be inflamed by voices relating such practices to a new type of greenwashing behavior, an irresponsible tactic, whose main aim is to deceptively manipulate the stakeholder views of the firm (Siano et al., 2017). Against this background, we hypothesize that:

H1a: *Ceteris paribus*, labor unionization is associated with corporate CSR reporting.

The impact of organized labor on the initiation of CSR reports may not be uniform across the board. Prior studies argue that the geographical concentration of organized labor significantly increases unions' power (e.g., Cox and Oaxaca, 1982). Indicatively, highly-unionized states have established institutionalized norms of wage equality by having higher minimum wages and more equal income distributions (Bassanini et al., 2017, Western and Rosenfeld, 2011). We expect that managers of unionized organizations whose headquarters are located in highly-unionized states would more intensively resort to CSR disclosure; this is due to a desire to communicate their social and management standing (Brady and Wallace, 2000) to attain and maintain a convergence of interest with unionists. Hence, we extend *H1a* by testing whether a unionized firm headquartered in one of the seven states where more than half of all union members live is likely to engage in CSR reporting. Thus, we hypothesize that:

H1b: *Ceteris paribus*, there is an association between unionized firms headquartered in US states with more intense unionist concentration and CSR reporting.

2.3 Unionized firms, political ideology and CSR reporting strategies

Unionized firms' stakeholder integration strategies, and in particular CSR reporting, may also be influenced by the prevailing political ideology in a geographical area, which could substantially affect unions' position of influence and bargaining power (Kerrissey and Schofer, 2013). Indeed, depending on the political ideologies prevailing in their locations, unions may operate in more- or less-favorable contexts, which would empower or impede their role accordingly, also affecting management communication strategies (Doh and Guay, 2006, Tope and Jacobs, 2009).

Historically, the relationship between organized labor and the Democratic Party is marked by the long-term commitment of Democrats to support labor unions in their demands for fair pay and economic security (Beland and Unel, 2018). These ties date back to the 1930s, when Franklin D. Roosevelt passed the National Labor Relations Act (Wagner Act, 1935) and extended collective bargaining to public workers (Klein, 2017). Further, the adoption of public-sector collective bargaining laws by different states during the 1960s, 1970s, and early 1980s escalated the rise of public-sector unions, and sealed the advantage of Democrats over Republicans on labor issues (Anzia and Moe, 2016). This long-term support by Democrats for union expectations and demands (e.g., the 2016 Workplace Democracy Act and the 2009 Employee Free Choice Act, EFCA) has, in turn, provided tangible benefits for the Democratic Party (Chen et al., 2012). Indeed, unions systematically devote financial and organizational resources to persuading constituencies to vote for the Democratic Party. For instance, to influence the 2017-2018 Elections, union officers spent approximately $1.37 billion, donating an overwhelmingly greater part of those figures to the Democrats (NILRR, 2019).

In contrast, conservative political ideologies, which emphasize economic freedom and property rights, conceptualize collective bargaining as a restricting parameter that undermines long-term corporate economic prospects and prosperity (Bartels and Brady, 2003). Conservative politicians, generally members of the Republican Party, have become increasingly resistant to policies empowering labor (Anzia and Moe, 2016) and have paved the way for corporate employers to invest in an anti-labor management consulting industry to weaken union bargaining power (Klein, 2017).[3]

Against this background, we interpret the dominance of the Democratic Party in a state as an important parameter, which empowers the position of unions and their bargaining capability. Thus, operating in contexts where the union's views are politically well-received and supported, unionized firm managers are highly likely to be more susceptible to issuing standalone CSR reports. These reports can be used as a communication mechanism and integral part of the broader management integration strategy to establish links with unions, reduce information asymmetries, align interests and signal commitment to the concerns and needs of society and employees. Therefore, we hypothesize that:

H2: *Ceteris paribus*, there is an association between the US states' prevailing political ideology unionized firms are headquartered in and CSR reporting.

[3] For instance, during the Reagan/Bush era, a systematic attempt took place to reverse as much labor law doctrine as possible (Dark III, 1999). In 2011, in Wisconsin, conservative politicians passed a labor law that restricted collective bargaining rights for public employees (Anzia and Moe, 2016). Similar efforts took place in several other Republican-controlled states; *inter alia*, Idaho, Ohio, and Tennessee. Similarly, conservative politicians and lobby groups organized a very successful public campaign against unionization at Volkswagen in Chattanooga, Tennessee (Mueller and Stegmaier, 2017).

3 Research Design

3.1 Sample

We examine the period from 2002 onwards because of data restrictions for CSR reporting activity through the Thomson Reuters ASSET4 database. We construct our sample for the period 2002–2015, using the entire universe of firms in the Kinder, Lydenberg, and Domini (KLD) database. Accounting and ownership structure data are obtained from the Compustat and Thomson Reuters EIKON databases. We begin with 3,399 firms and then drop 18 of them due to a missing Central Index Key (CIK), since this is a prerequisite field for downloading the firm's 10-K report and identifying its unionization status (see section 3.2). Our data requirements on control variables for our main model (1) necessitate dropping a further 860 firms due to missing data and another 29 firms because of missing ownership-structure data. We further eliminate 172 firms with headquarters in foreign countries or outlying US territories (e.g., Puerto Rico, the Virgin Islands and Guam), using each firm's historical business address as extracted from its filings (see Marciukaityte, 2015).[4] Next, following established literature (Cheng, 2017), we exclude 505 firms which lack union-related expressions in their 10-K filings (see section 3.2), to avoid arbitrarily defining certain companies as non-unionized. Our sample thus far comprises 1,815 firms, which translates into 11,405 firm-years, for which we retrieve CSR reporting activity from the Thomson Reuters ASSET4, CorporateRegister.com and CSRwire databases.

Following studies which have argued that a firm's CSR orientation might suffer problems related to self-selection (i.e., Gao et al., 2014), we employ a propensity score-matched (PSM) approach and create matching pairs of firms with and without CSR reporting

[4] We obtain each firm's historical business address through 10-K filings, since databases tend to backfill business addresses (Marciukaityte, 2015). We download company filings, as available through the Securities Exchange Commission FTP server, and develop a PERL script that parses state code, state name, city, and zip code.

activity, as described in section 4.2. The necessity to engage a PSM sample stems from our observation that significant variations exist between CSR reporting firms and non-CSR reporting firms (see section 4.3). Our final sample for analysis comprises 747 firms, which translates into 2,526 firm-years (or 1,263 matched pairs).

3.2 Measuring labor unionization

We operationalize a firm-level unionization measure to indicate the existence of organized labor (Chantziaras et al., 2020b, Cheng, 2017); however, we further sensitivity test for alternative definitions (see section 5.1). We work at firm level since the measurement error is lower (similar to Cheng, 2017). In order to determine whether company employees are organized (i.e., covered by a collective bargaining agreement), we draw upon Item 1 (Business) of 10-K company filings. First, we download the company filings available from the Securities Exchange Commission FTP server, and develop a PERL Script (similar to Chantziaras et al., 2020b, Cheng, 2017). This allows us to parse the sentences related to union coverage. We employ a battery of keyword combinations in our code, such as: bargaining agreement(s), bargaining unit(s), collective agreement(s), collective bargain(ing), labo(u)r agreement(s), labo(u)r organization(s), labo(u)r union(s), organized labo(u)r, organiz(s)ed employee(s)/staff/personnel/workforce, work council(s), trade union(s), union('s) activity(ies), union('s) agreement(s), union contract(s), union organization(s), unioniz(s)ed and union(s) (similar to Chantziaras et al., 2020b, Cheng, 2017). Once our code parses all the union-related sentences, we manually verify and identify observations of companies disclosing the percentage of employees covered under collective bargaining agreements (*PCT_UNION*) and, for the control group, observations of companies that firmly report no union representation.

In order to align our findings with previous literature, and also to examine whether measurement errors between firm- and industry-level data seriously impede empirical findings, we estimate a unionization proxy by employing industry-level data. Using data from the Union Membership and Coverage Database (UMCD), we estimate *UNION_IND* by multiplying the percentage of employees covered by collective bargaining in a firm's primary Census Industry Classification (CIC) industry with the number of company employees over lagged total assets (Chen et al., 2011, 2012, Chyz et al., 2013, Hilary, 2006). Since UMCD data are available in CIC codes, we use a crosswalk list (retrieved from the US Census Bureau) and convert CIC to Standard Industrial Classification (SIC) codes (https://www.census.gov/people/io/methodology/).

3.3 Empirical model

We employ a logistic regression to examine the association between firm unionization and the probability of a company issuing a standalone CSR report. In order to address endogeneity concerns, we use a lead-lag approach and build our model specification similar to that of Dhaliwal et al. (2011):

$$\log\left[\text{prob(DCSR)}/\left(1-\text{prob(DCSR)}\right)\right] = \beta_0 + \beta_1\text{Union}_{t-1} + \beta_2\text{PERF}_{t-1} + \beta_3\text{GOV}_{t-1} + \beta_4\text{INS\_OWN}_{t-1} + \beta_5\text{INST\_OWN}_{t-1} + \beta_6\text{SIZE}_{t-1} + \beta_7\text{AGE}_{t-1} + \beta_8\text{TOBINQ}_{t-1} + \beta_9\text{ROA}_{t-1} + \beta_{10}\text{LEV}_{t-1} + \beta_{11}\text{LIT}_{t-1} + \beta_{12}\text{FIN}_{t-1} + \beta_{13}\text{COMP}_{t-1} + \beta_{14}\text{GLOBAL}_{t-1} + \beta_{15}\text{LIQUIDITY}_{t-1} + \sum\text{YEAR} + \sum\text{INDUSTRY} + \varepsilon \quad (1)$$

The dependent variable (*DCSR*) in our model incorporates data from three sources (namely the Thomson Reuters ASSET4, CorporateRegister.com and CSRwire databases) and indicates whether a firm discloses a standalone CSR report in year t or not. The main variables of interest are captured under the vector *Union*, which represents the union-related variables employed to capture the impact of unionization on CSR reporting. We include firm- (*PCT_UNION*) and industry-level unionization proxies (*UNION_IND*), as discussed in section

3.2. CSR performance (*PERF*) is included in the model since it has been found to be positively associated with CSR reporting (Dhaliwal et al., 2011). Following Chantziaras et al. (2020a), we construct a CSR index by totaling the positive (strengths) and negative (concerns) indicators of six KLD categories (i.e., community, diversity, employee relations, environment, human rights and product), while we exclude the category for corporate governance since it is regulated.

We rely on the relevant literature and include indicators with a demonstrable influence on CSR reporting (see Appendix for variable definitions). Thus, in our model, we include corporate governance activity (*GOV*) using the KLD ratings (similar to Grougiou et al., 2016), since CSR disclosures are positively associated with strong corporate governance mechanisms (Jo and Harjoto, 2011). We also control for ownership structure, measured as the percentage of shares owned by insider (*INS_OWN*) and institutional investors (*INST_OWN*) (Grougiou et al., 2016).

Larger and older firms are more visible to the public and hence, face significant pressures from a wide group of stakeholders, which incentivizes them to behave in a socially responsible way (Brammer et al., 2009). Thus, we include firm size, calculated as the natural logarithm of the market value of common equity (*SIZE*) (Dhaliwal et al., 2011), and company age, expressed as the natural logarithm of years that the firm has appeared in Compustat (*AGE*) (Khan et al., 2013). Future growth opportunities are operationalized using *TOBINQ*, defined as the market value of common equity plus the book value of preferred stock, plus the book value of long-term debt and current liabilities, scaled by the book value of total assets (Dhaliwal et al., 2011). We also incorporate measures of financial performance (*ROA*) and leverage (*LEV*), calculated respectively as income before extraordinary items scaled by total assets, and total debt scaled by total assets. This is because firms with better financial performance are more likely to have the resources to practice CSR activities and produce CSR reports (Dhaliwal et al., 2011), while debt servicing plays a monitoring role and increases the overall demand for disclosure as a strategy to mitigate the risk of litigation (Skinner, 1997). We operationalize litigation risk (*LIT*) as a dichotomous variable indicating the existence/non-existence of a legal proceeding against the firm under SEC regulation S–K §229.103 (Grougiou et al., 2016).

Voluntary disclosure may occur as a part of a broader strategy to attract investors when seeking to raise capital on public markets (Dhaliwal et al., 2012). We consider a firm's financing activity (*FIN*) calculated as: [(sale of common and preferred shares - the purchase of common and preferred shares) + (the long-term debt issuance - the long-term debt reduction) over total assets] (Dhaliwal et al., 2011). Proprietary costs arising from product market competition have a negative impact on disclosure incentives (Dye, 1985). For this reason, we employ a product market competition proxy measured by the Herfindahl-Hirschman index (*COMP*), calculated as the sum of the squared fractions of sales in an industry (industries are defined according to the two-digit SIC codes) multiplied by minus one (Dhaliwal et al., 2011). This proxy has been suggested as a way to additionally capture corporate visibility and public pressure over CSR activity (Dhaliwal et al., 2012).

Influenced by studies reporting a positive relationship between corporate disclosure and a firm's global orientation (Dhaliwal et al., 2011), we construct an indicator variable for foreign income reporting (*GLOBAL*). Liquidity has also been found to be influential in CSR disclosures (Leventis and Weetman, 2004). Thus, we measure *LIQUIDITY* by relating the number of shares traded in the year to the total shares outstanding at the year-end (Dhaliwal et al., 2011). Finally, we consider industry and year fixed effects (Dhaliwal et al., 2011) by clustering standard errors at firm-level (Jo and Harjoto, 2011) and winsorizing all the continuous variables in the 1st and 99th percentile, to mitigate any effects from outliers.

4 Results

4.1 Univariate analysis

In Table I, we provide the descriptive statistics of the entire sample, both CSR reporting firms and non-CSR reporting firms, where CSR reporting firms (*DCSRREPORTER*) are defined as those disclosing at least one CSR report within the 14-year analysis window (2002–2015). We also compare the means and report the corresponding statistical significance between the two groups. The CSR reporting firms and the non-CSR reporting firms significantly differ across all control variables, indicating potential problems related to self-selection (i.e., Gao et al., 2014).

[Insert Table I about here]

As indicated, CSR reporting firms are more unionized (20.0%) than non-CSR reporting firms (5.7%), with the corresponding difference being statistically significant at 1%. With respect to other control variables, the mean values of all control variables (except *ROA* and *LEV*) are similar to those reported in Dhaliwal et al. (2011). CSR reporting firms are larger in size and older than their counterparts; whereas they exhibit a significantly lower level of financing (*FIN*), are more leveraged (*LEV*) and more liquid (*LIQUIDITY*), and exhibit higher levels of global operations (*GLOBAL*), similar to Dhaliwal et al. (2011). Approximately 20% of our sample firms are involved in a major litigation, with CSR reporting firms exhibiting a higher litigation probability, thus providing confirmation of the findings of Dhaliwal et al. (2011) for their industry-based litigation measure. Lastly, we observe no significant variations between CSR reporters and non-CSR reporters in terms of geographical dispersion across states where Democrats predominate (the mean values of *DEMOCRATS* are 0.593 and 0.621 respectively).

4.2 Self-selection and propensity score matching

Descriptive statistics illuminate significant differences between CSR reporting and non-CSR reporting firms, which raises concerns of self-selection; this is because the standardized differences lie outside the threshold of ±20, suggesting large differences (Rosenbaum and Rubin, 1983). While we rely on a lead-lag approach to mitigate potential reverse causality (Dhaliwal et al., 2011), we further follow Gao et al. (2014) to accommodate self-selection issues related to CSR. We employ a PSM technique to moderate problems related to differences between groups. We build our first-stage logistic model for CSR reporting firms following previous studies (e.g., Gao et al., 2014, Lys et al., 2015) as follows:

$$\mathrm{Prob}(\mathit{DCSRREPORTER} = 1) = \mathrm{logit}(\beta_0 + \beta_1 \mathrm{PERF}_{t-1} + \beta_2 \mathrm{GOV}_{t-1} + \beta_3 \mathrm{INS\_OWN}_{t-1} + \beta_4 \mathrm{INST\_OWN}_{t-1} + \beta_5 \mathrm{SIZE}_{t-1} + \beta_6 \mathrm{AGE}_{t-1} + \beta_7 \mathrm{TOBINQ}_{t-1} + \beta_8 \mathrm{ROA}_{t-1} + \beta_9 \mathrm{LEV}_{t-1} + \beta_{10} \mathrm{LIT}_{t-1} + \beta_{11} \mathrm{FIN}_{t-1} + \beta_{12} \mathrm{COMP}_{t-1} + \beta_{13} \mathrm{GLOBAL}_{t-1} + \beta_{14} \mathrm{LIQUIDITY}_{t-1} + \beta_{15} \mathrm{FRCF}_{t-1} + \beta_{16} \mathrm{SALES}_{t-1} + \beta_{17} \mathrm{PM}_{t-1} + \sum \mathrm{YEAR} + \sum \mathrm{INDUSTRY} + \varepsilon) \quad (2)$$

As previously defined, the dependent variable of our model is an indicator variable, *DCSRREPORTER*, which captures whether a firm has disclosed at least one CSR report within the 14-year analysis window (2002–2015). We include all the control variables from model (1), apart from unionization proxies and augment for free cash flow (*FRCF*), calculated as the difference between cash flow from operations and cash flow used in investing activities, since firms with freer cash flow have more resources to invest in CSR (Lys et al., 2015). Following Gao et al. (2014), we add *SALES* and *PM*, defined respectively as sales revenue divided by total assets, and income before extraordinary items divided by sales revenue.

First, we run a logit model (see Table II) in which the dependent variable is *DCSRREPORTER* over the aforementioned covariates, estimated for the 11,405 firm-year observations with data available for model (2); of which 3,372 are classified as CSR reporting firms. The overall model is significant (Wald χ^2=2,852.169, p<.001) and adequately fits the data (Hosmer-Lemeshow χ^2=11.745, p=0.163 and the area under the ROC curve is 0.902). The estimation results indicate that 11 out of the 17 covariates are significant, while the pseudo R^2 is 43.3%. Our results support the notion of larger and older firms, with better CSR performance and global orientation, being more likely to be CSR reporters.

[Insert Table II about here]

Next, we calculate the propensity scores using predicted probabilities from the logistic regression; we match each CSR reporter firm-year to a control firm-year with the closest propensity score, under a nearest-neighbor matching approach without replacement and a caliper constraint (δ=0.1) (as Gao et al., 2014). We ensure that all matched pairs belong in the same year and two-digit SIC industry, since CSR reporting patterns vary across industries (Reverte, 2009). This process yields 1,263 matching pairs (37.45% of the CSR-reporting firm-year observations being matched). In Table III, we tabulate the differences in means and standardized bias of the sample, both before and after the propensity score matching. We observe that none of the differences in means is significant, nor do the standardized differences exceed the threshold of ±20 for the sample subsequent to the propensity score matching, which clearly indicates that the matching was successful in achieving balance across all covariates.

[Insert Table III about here]

Finally, we report a Pearson's correlation matrix in Table IV, where the majority of variables are significantly correlated with *DCSR*. *GLOBAL*, *AGE*, and *PERF* exhibit the highest coefficients, with values of 0.16, 0.12, and 0.11 respectively. Firm-level labor unionization is significantly correlated with *DCSR* at 1%, while industry-level unionization is not. Other inferences suggest that multicollinearity is not a serious problem.

[Insert Table IV about here]

4.3 Multivariate analysis

The results of our analyses are presented in Table V. Both models, using firm- and industry-level unionization proxies (Columns 1 and 2 respectively), are significant (Wald χ^2 p<.001) and adequately fit the data (Hosmer-Lemeshow χ^2=6.662 (6.083), p=0.57 (0.63)), while the area under the ROC curve is 0.799 for model (1) and 0.792 for model (2). Firm-level unionization is significant at 1%, while industry-level unionization is insignificant, with the explanatory power of our models being approximately 18%. The coefficient of *PCT_UNION* is positive and statistically significant at 1% (Column 1, β=1.862, z-stat=3.41). This lends support to *Hypothesis 1a*. Based on the marginal probability effect of unionization (0.182), a one standard deviation (0.193) increase in the level of unionization leads to a 3.51% (0.182 × 0.193) increase in the probability of a firm issuing a standalone CSR report. *UNION_IND* is positive but non-significant (Column 2, β=13.950, z-stat=1.23), which implies a possible measurement error of unionization using industry- rather than firm-level data.

[Insert Table V about here]

Drawing upon Column 1 in Table V, most coefficients across control variables have the predicted sign (except for *ROA* and *FIN*). *PERF* is positive and significant at 5%, suggesting that superior CSR performance encourages companies to issue CSR reports voluntarily (Grougiou et al., 2016). Firms larger in size, exhibiting lower profitability and growth opportunities, and raising less capital in the previous year (*SIZE*, *ROA*, *TOBINQ* and *FIN*) are more likely to engage in CSR reporting. The mean Variance Inflation Factor (VIF) has values less than 1.383 across all model specifications, lower than the conservative cut-off value of 5 (e.g., Studenmund, 2016), and thus implies no multicollinearity. Overall, the results support the previous literature and additionally demonstrate that unionization loads positively to CSR reporting activity.

We rely on the annual Union Members Summary data to test *H1b*. This is issued by the Bureau of Labor Statistics (BLS) (http://www.bls.gov/) and documents that approximately half of the total unionized labor force across the US resided in just seven states during the period of our investigation.[5] With this in mind, we use an indicator variable to capture whether a firm's headquarters is located in one of the seven aforementioned, heavily-unionized states (*UNIONHALF*); we also augment our *H1a* with an interaction term between *PCT_UNION* and *UNIONHALF*. Column 3 in Table V indicates that the effect of unionization on CSR reporting increases when the firm headquarters are located in one of the states where more than half of all union members live. The coefficient of *PCT_UNION*×*UNIONHALF* is positive and statistically significant at 5% (Column 3, β=1.829, z-stat=2.02), while the model is significant (Wald χ^2 p<.001) and adequately fits the data (Hosmer-Lemeshow χ^2=9.456, p=0.305; the area under the ROC curve is 0.801). This provides support for *H1b*.

To test our second hypothesis, we augment our model with an interaction term between *PCT_UNION* and an indicator variable *DEMOCRATS* (Column 4 in Table V). We collect presidential election data (http://www.uselectionatlas.org) for each election cycle separately and create an indicator (*DEMOCRATS*) that equals one if the Democratic Party received the most votes in the state (i.e., the number of seats in the electoral college) in the presidential elections of 2000, 2004, 2008 and 2012. Column 4 in Table V indicates that the dominance of the Democratic Party magnifies the positive impact of unionization on CSR reporting activity. The coefficient of the interaction term *PCT_UNION*×*DEMOCRATS* is positive and

[5] These states are, in rank order: California, New York, Illinois, Michigan, Pennsylvania, Ohio and New Jersey; with the last two switching rankings during the period 2006–2009. Notably, California and New York together account for approximately 4.5 million out of the 15.5 million union members (on average) across the US within our 14-year window.

statistically significant at 1% (Column 4 in Table V, β=2.626, z-stat=3.04), while the model is significant (Wald χ^2 p<.001) and adequately fits the data (Hosmer-Lemeshow χ^2=9.958, p=0.268; the area under the ROC curve is 0.804). Thus, we accept *Hypothesis 2*.

We further plot the associations of the interaction terms and we mean-center *PCT_UNION* to ease interpretation (e.g., Burks et al., 2019).[6] Figure I illustrates that the geographical concentration of organized labor adds to union influence which, in turn, significantly affects the propensity for CSR reporting activity.[7] Figure II demonstrates how states governed by Democrats are associated with higher levels of unionization, which then increases the positive effect of unionization on CSR reporting engagement.

[Insert Figure I about here]

[Insert Figure II about here]

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5 Sensitivity Testing

We probe the sensitivity of our findings by employing alternative definitions of: a) unionization (5.1), b) CSR reporting (5.2), and c) CSR performance (5.3). We verify that they are all persistent. We further account for endogeneity (5.4) and specification issues related to variable omission (5.5).

[6] Burks et al. (2019, p. 72) argue that mean-centering the continuous constituents of the interaction term does not change the coefficient on the interaction effect; rather, it enables the researcher to "*meaningfully interpret*" the main effect.

[7] Both California and New York favor stricter controls on environmental, fair business practice, and employee safety issues as compared to most other states. For example, California has been a leader in requiring supply chain disclosures related to human trafficking (Birkey et al., 2016a) and introduced the Transparency in Supply Chains Act (effective as of 1 January, 2012). To mitigate any concern of these locales driving our results due to social and political exposures related to other CSR issues, we conduct sensitivity analyses and exclude firms with corporate headquarters in a) California, b) New York, and c) both sates.

5.1 Alternative definitions of unionization

We substitute our main unionization proxy, namely the percentage of unionized employees (*PCT_UNION*), with two alternative specifications. First, we employ an indicator variable signaling the existence of collective bargaining coverage (*D_UNION*). Second, following Cheng (2017) and Chantziaras et al. (2020b), we estimate a comprehensive measure of the overall influence of organized labor measures (*CB*).[8] Running the models, all coefficients of alternative unionization proxies remain positive and statistically significant at 5% for *D_UNION* and at 1% for *CB*.

5.2 Alternative measures of CSR reporting

In addition to our CSR disclosure proxy (*DCSR*) we employ three alternative specifications to test our hypotheses. First, we create an indicator variable which captures first-time CSR reporters (*DCSR_FIRST*) (Dhaliwal et al., 2011). Second, we employ an indicator for the CSR report being compliant with GRI standards (*GRI_COMP*).[9] Third, we approximate CSR disclosure by using the actual number of CSR reports (*NR_DCSR*) issued by the company in year t. We perform a logistic regression using *DCSR_FIRST* and *GRI_COMP* as dependent variables and an ordered logistic regression using *NR_DCSR*. We obtain results qualitatively similar to our main analyses.

[8] *CB* is calculated as the first principal component of the following: the percentage of unionized employees (*PCT_UNION*), the coverage dummy (*D_UNION*) and a union-related risk dummy (*UNION_RISK*). Similar to Cheng (2017) and Chantziaras et al. (2020b), we parse union-risk related expressions included in Item 1A (Risk Factors), since firms report idiosyncratic factors (e.g., labor relations, labor union activity) under this item. We construct an indicator variable (*UNION_RISK*) that is set to 1 if the company discloses risks related to: 1) union presence, 2) union organizing activity, 3) expiry of bargaining agreements (either in current or next year), 4) work stoppages or 5) negative impact on firm performance and/or profitability. We identify 728 instances of union-related risk. This variable is only available for the years 2005 and onwards since the SEC mandated the disclosure of Item 1A Risk Factors.

[9] We collected data on the GRI compliance of CSR reports through two sources: a) the Thomson Reuters EIKON database, and b) the Sustainability Disclosure Database.

5.3 Alternative measures of CSR performance

We test for CSR performance by using only "strengths" (*PERFS*) or only "concerns" (*PERFC*), similar to Grougiou et al. (2016). We further substitute the *PERF* and *GOV* variables with CSR data from the ASSET4 database (e.g., Ioannou and Serafeim, 2012) and create the measures *PERF4* and *PERF_CG*. We also test for alternative definitions of CSR performance, as presented in Dhaliwal et al. (2011). We create an indicator *CRO* that equals 1 if a firm was on the "100 Best Corporate Citizens" list (by Corporate Responsibility Officer) in year t, and 0 otherwise (Dhaliwal et al., 2011). We also measure CSR performance by using an indicator variable (*DJSI*) that signals for membership of the Dow Jones World Sustainability Index in year t. When we run the regressions again employing these proxies our inferences remain unchanged.

5.4 Endogeneity tests

Our analyses suggest a positive association between employee unionization and firm CSR reporting propensity. However, it is difficult to establish causal inferences due to potential endogeneity concerns, as "*labor union existence or the unionization rate is determined by a firm's CSR activity*" (Chun and Shin, 2018, p. 12). To mitigate such concerns, we conduct tests that account for potential endogeneity, such as a) the use of instrumental variables approach (e.g., Chen et al., 2011, 2012, Chino, 2016, Chun and Shin, 2018) and b) the implementation of a difference-in-difference approach (D-i-D) (e.g., Arslan-Ayaydin et al., in press).

Following Chen et al. (2011) and Chun and Shin (2018), we utilize the natural logarithm of the average age of workers in each industry (*WORKREAGE*) every year, employing data from the NBER CPS Merged Outgoing Rotation Groups File (http://www.nber.org/morg/annual/).[10] We expect a positive association between *WORKREAGE* and *PCT_UNION*, since senior workers have a relatively strong job attachment and low mobility, as compared to young workers, and their expected benefits from unionization are likely to be high (Chen et al., 2011, Chun and Shin, 2018). On the other hand, we do not expect *WORKREAGE* to directly affect a firm's CSR engagement (i.e., be correlated with the error term in the second stage regression). Thus, this could be a reasonable instrument for identifying the effects of unionization on CSR reporting. Our results indicate that *WORKREAGE* has a positive and significant coefficient, in the first-stage regression, and its *F*-statistic (18.015) is above the threshold of 10 (Staiger and Stock, 1997), which suggests a strong instrument. In the second stage of the instrumental variables (IV) regressions of the probit model (the "ivprobit" procedure in Stata), the coefficient of *PCT_UNION* remains positive and statistically significant at 1%, suggesting that the effect of unionization on CSR reporting remains significant, after controlling for potential endogeneity.

Beyond the IV estimates, we enrich our sensitivity tests through a natural experiment design and the implementation of a D-i-D approach. Following Arslan-Ayaydin et al. (in press), we exploit the natural experiments in Indiana and Michigan, which adopted right-to-work (RTW) laws in 2012. We select the RTW legislation as an exogenous variation in the power of labor unions, since its enactment prevents unions from requiring membership, or payment of union fees, as a precondition of employment, either before or after an employee is hired (Marciukaityte, 2015, 2018). This legislative framework reduces unions' density due to its reduced symbolic effect on employees, resources, activism(strikes) and bargaining power in advancing employees' interests and collectively regulating work (Ellwood and Fine, 1987,

[10] The NBER CPS Merged Outgoing Rotation Groups File data are available in Census Industry Classification (CIC) codes. We follow the methodology described in Section 3.3 and transform each firm's primary CIC into Standard Industrial Classification (SIC) codes, using a crosswalk list available through the U.S. Census Bureau (https://www.census.gov/people/io/methodology/).

Marciukaityte, 2015, 2018) and hence; has received fierce reactions from unions (Marciukaityte, 2015).[11]

We design the D-i-D as follows. First, we create an indicator variable (*ChRTW*) that equals 1 for firms located in one of the two aforementioned states (Indiana and Michigan). Second, we use a dummy (*POST*) that takes the value of 1 for observations in 2012 or after. Third, we include in the model with the triple interaction *PCT_UNION*×*ChRTW*×*POST*, alongside with the relevant interactions between the three terms. Running the model reveals that *PCT_UNION*×*ChRTW*×*POST* attains a negative and significant coefficient at 5%, while the coefficient of *PCT_UNION* remains positive and statistically significant at 1%.[12] Therefore, we conclude that the positive effect of employee unionization on CSR reporting is mitigated once the state enacts RTW laws.

5.5 Variable omission

We sensitivity test for a battery of variables that have been found or suggested to be influential in CSR reporting but are not included in our full model because of data and/or specification reasons. First, we employ alternative measures of firm size: 1) total assets (Khan et al., 2013), 2) sales (Grougiou et al., 2016), and 3) number of employees (Lau et al., 2016); all transformed into natural logarithms. The results from all different measures of firm size remain positive and statistically significant. Second, we account for enhanced CSR activity which is attributable to company diversification (Ioannou and Serafeim, 2012). To this end, we include the natural logarithm of the number of business segments the company operates

[11] It is demonstrated that, after the enactment of RTW laws, the number of employees participating in unions is substantially reduced by five to eight percent, while there are six to ten percent more "free riders", i.e., employees who are covered by collective bargaining agreements but are not union members, in RTW states when compared with their non-RTW counterparts (Moore, 1998).

[12] In untabulated tests, we further examine the parallel trend assumption of our D-i-D setting. We verify that it is not violated, since the difference between CSR reporting propensity remains constant over time either when considering firms headquartered in states that adopted RTW laws (treatment) or their counterparts in states that did not adopt RTW laws (control).

(*OPSEG*); which positively loads to CSR reporting but lacks statistical significance. Despite the majority of previous studies approximating firm performance by employing *ROA*, some other studies control for performance using the ratio of income before extraordinary items over net assets (*ROE*) (Grougiou et al., 2016). Including *ROE* in our model does not affect our results, since the *ROE* coefficient remains statistically significant and in the same direction with *ROA*. We also incorporate an alternative approximation of growth opportunities (different to *TOBINQ*), i.e., the market-to-book ratio (*MB*) (Grougiou et al., 2016), which lacks statistical significance. The incorporation of all the above variables does not change our inferences.

Liquidity, share performance and market risk have also been suggested as influential to CSR disclosures (Leventis and Weetman, 2004). Thus, we augment our model with: current ratio (*CR*), measured as current assets to current liabilities (Leventis and Weetman, 2004); share return (*SHR*), estimated as [(market price year end + dividends per share + quarterly special dividends) / (previous year's market price year end - 1)] *100 (Grougiou et al., 2016); and the daily stock return volatility over the fiscal year (*VOL*) (Ioannou and Serafeim, 2012). None of these proxies is statistically significant, nor does the direction or magnitude of the *PCT_UNION* coefficient change. Considering prior literature (e.g., Brammer et al., 2009, McWilliams and Siegel, 2001), we control for advertising (*ADV*) and research and development (*RD*) intensities, both scaled by net sales (McWilliams and Siegel, 2001); however, we do not observe any statistical significance in these proxies, though they do positively load to CSR reporting.[13] Since CSR disclosures can be correlated with the overall corporate financial transparency and quality of a firm (Dhaliwal et al., 2011), we control for financial disclosure quality using the absolute value of abnormal accruals (*ABS_JDA*), based on Dechow et al. (1995). Again, our results remain unchanged.

[13] Given the high number of missing values for *ADV* and *RD* variables, we use industry averages as a proxy for missing observations (similar to Marano and Kostova, 2016).

In addition, we examine alternative specifications of ownership structure. Jain and Jamali (2016) argue that block ownership (*BLOCK_OWN),* measured as the percentage owned by investors with at least a 5% stake (Jo and Harjoto, 2011)) tends to discourage proactive CSR. Khan et al. (2013) report a positive association between the percentage owned by foreign investors (*FOR_OWN*) and CSR disclosure. We also incorporate into our model indicator variables signaling that the firm operates in environmentally-sensitive industries (*ESI* – i.e., chemical, paper, metals, petroleum, mining and extractive, or utility industries (Birkey et al., 2016b)) or in controversial industries (*SIN* – i.e., alcohol, firearms, gambling, military, nuclear power or tobacco operations (Grougiou et al., 2016)). Our inferences remain unchanged after controlling for these proxies.

Further, prior literature indicates that firm location is influential to corporate CSR engagement (e.g., Boeprasert, 2012, Ding et al., 2019, Husted et al., 2016, Jiraporn et al., 2014). More specifically, there is evidence that a) firm location (Boeprasert, 2012, Ding et al., 2019), b) influence by geographic peers (Jiraporn et al., 2014), and c) high levels of local CSR density (Husted et al., 2016) are influential on a firm's CSR policy and engagement. We capture such effects by the following: first, we create an indicator (*URB*) that equals 1 for firms headquartered in MSAs with at least 1 million residents (as defined by the U.S. Census), and 0 otherwise (Francis et al., 2016); second, we use a dichotomous variable (*UAGG*) for firm headquarters located in an urban agglomerate area, namely one of the following MSAs: New York City, Los Angeles, Chicago, Washington, Baltimore, San Francisco, Philadelphia, Boston, Detroit, Dallas, and Houston (Francis et al., 2016); third, we operationalize local CSR density (*CSRDENS*) as the spatial distribution of CSR engagement by firms surrounding the focal firm, as determined by the location of its headquarters (Husted et al., 2016).[14] Our inferences remain unchanged.

As additional robustness exercises, we examine whether the significantly positive effect of unionization on CSR reporting persists, after controlling for employee welfare, CSR reporting quality, firm CSR experience, and other external factors influencing CSR reporting propensity. Following Arslan-Ayaydin et al. (in press), we use a proxy that captures the quality of employee treatment, namely employee welfare (*EMPWELFARE*). We operationalize *EMPWELFARE* as the difference between the number of strengths and number of concerns on the employee relationship dimension as per KLD data. Additionally, we include measures of CSR reporting quality and for firm CSR experience, through the incorporation of a) an indicator variable that equals one if the company's CSR report is compliant with GRI standards (*GRI_COMP*), and b) the number of years the company issues a CSR report (*YRSCSR)* by relying upon data through Thomson Reuters ASSET4, CorporateRegister.com and CSRwire databases. We additionally adjust our model to consider for influential external factors affecting a firms CSR propensity, such as: a) the number of analysts following the firm (*ANALYSTS)* obtained through Thomson Reuters EIKON and b) an indicator that equals one if the company is named as a defendant in a lawsuit, attributable to employee-related issues such as violations of one of the following: 1) Employment Law, 2) Fair Labor Standards Act, 3) Labor - Management Relations, 4) and Labor Law (*LIT_LABOR)*, obtained through Audit Analytics. The incorporation of all the aforementioned variables leaves our inferences unaffected.

[14] CSR engagement is proxied by aggregating five positive (strengths) indicators from the five KLD categories: community, diversity, employee relations, environment and human rights; while we exclude the category for corporate governance and product, similar to Husted et al. (2016). We compute CSR density using the following formula: $\mathrm{CSRDENS_i} = \sum_j \frac{\mathrm{KLDStr_{jt}}}{(1+\mathrm{d_{ij}})}$, where: *i* refers to the focal firm; *j* to all other firms at year *t*; and *d* is the distance between firm *i* and firm *j*.

Finally, we test for additional demographic controls. Recourse to the annual Union Members Summary (issued by the BLS) for the period 2002–2015 demonstrates how the propensity to unionize relates to gender, race and age. Furthermore, previous studies have emphasized that education is influential to unionization propensity (Eren, 2009). Thus, we control for: the percentage of male (*MALE*) and black workers (*BLACK*); the percentage of workers between the ages of 45 and 64 (*EMP45_64*); and the percentage of workers having attended college for at least four years (*COLLEGE*); all measured at the industry level.[15] Once again, our results are not affected.

6 Conclusion

We embarked upon an examination of the impact of labor unions, a salient internal stakeholder, on CSR reporting, a commonly-employed communication device. We developed empirical models to investigate whether managers in unionized firms are more active in initiating CSR reports than their counterparts in non-unionized firms. We additionally examined whether institutional characteristics, such as the geographical concentration of organized labor and the prevailing political ideology in a geographical area, drive unionized firm managers to disclose discretionary CSR reports more intensively. By employing a sample of 2,526 US firm-year observations for the estimation window of 2002–2015, we document that, in the presence of unions, managers tend to more intensively incorporate CSR reporting into their stakeholder integration and communications strategies. We also demonstrate that managers' propensity for CSR reporting significantly increases in states with a high concentration of unionized citizens and in geographical areas where the Democratic Party is at the helm.

[15] The data source is the NBER CPS Merged Outgoing Rotation Groups File, available on the NBER website: http://www.nber.org/morg/annual/ (Accessed 26 September, 2020).

We frame our results theoretically by drawing upon stakeholder theory and argue that perceptions of the management perplexities associated with unions become more emphatic when the contextual backdrop fosters the organizational and institutional role of unions. Thus, in unionized contexts, managers more intensively resort to CSR reports in order to: disseminate signals of commitment to ethical organizational values, reduce information asymmetries and facilitate the exchange of information; thereby identifying mutual interests. Such efforts on the part of managers are positively discounted by unions, which usually place high value on corporate devotion to social and environmental practices. Moreover, through CSR reporting, managers promote a more investment-friendly profile and attract the interest of financial stakeholders, including market participants hitherto hesitant to include unionized firms in their portfolios.

Our contribution to the literature is twofold. First, we extend CSR literature by demonstrating that CSR reporting operates as an important communication device, essentially related to a broader integration management strategy in the presence of organized labor. Second, we enrich current understandings by demonstrating that the role of organized labor on CSR reporting is moderated in the presence of institutional factors that empower the role and position of unions, i.e., the geographical concentration of organized labor and political ideology.

The implications of our study are important for managers, union leaders and market participants. Managers who operate under increased stakeholder pressure and complexity could benefit from our empirical analysis, which suggests that the initiation of CSR reports is a central, long-term strategic stakeholder integration and management communication device through which unobserved CSR qualities are publicly disclosed; in this way, dialogue with salient stakeholders, including unions, is greatly facilitated. Managers should also consider that voluntary CSR reports reduce information asymmetries with a wide range of stakeholders and may attract the interest of investors and other market participants who had not included unionized firms in their portfolios due to the perceived complexities associated with them.

Union leaders should view CSR disclosures as an integral part of a broader integration management strategy and communication tool, intended to smooth out relationships with salient stakeholders and, in particular, as an indication of management's effort to alleviate tensions. Moreover, unionists should be aware that CSR reports constitute an opportunity to identify mutual interests and align goals. Business analysts, investors and shareholders should be aware that standalone CSR reports are, *inter alia,* employed by managers to reduce information asymmetries and disparities with unions, and to communicate an investment-friendly context. To the extent that CSR reporting is associated with the prospect of financial benefits, market participants should factor such policies by unionized firms into their investment analyses.

We note some limitations to this study which, however, provide opportunities for further research. First, our data refer to the US context, which may limit the generalization of our results. Hence, researchers could employ cross-country datasets to overcome this limitation. However, they should also bear in mind that institutional and cultural differences in relation to CSR and unionization could bring further difficulties to deal with. Second, it would be important to know what benefits are enjoyed by the unionized companies that issue CSR reports. Relevant areas for such an investigation might include the cost of capital, union negotiation outcomes, corporate image and media exposure. Third, while we have followed an established method for measuring CSR disclosure employing CSR standalone CSR reports, we acknowledge that there is useful qualitative information we do not analyze. This analysis could potentially relate specific CSR information to unions' needs and demands. Further, there are alternative channels through which companies disclose relevant information such as 10-K filings, annual reports, firm websites, media, public announcements and so on. These are not captured by our data. Future studies could therefore extend our findings by considering the examination of qualitative information and alternative CSR communication channels. Finally, current understandings of CSR disclosure practices by unionized firms could be enriched by employing behavioral and organizational frameworks.

Appendix - Variable definitions

Variable	Definition
Panel A: CSR determinants for main model	
Dependent variable:	
DCSR	1 if the company discloses a CSR report according to Thomson Reuters ASSET4, CorporateRegister.com and CSRwire databases, and 0 otherwise.
Main variables of interest:	
PCT_UNION	Percentage of a company's unionized employees, as extracted from company filings.
UNION_IND	Industry level unionization, calculated as the product of the percentage of unionized employees (from Union Membership and Coverage Database – UMCD) in the industry with the number of the company's employees, scaled by lagged total assets, as in Hilary (2006).
UNIONHALF	1 if firm headquarters are located in one of the states where more than half of the total unionized labor force across the US live, 0 otherwise. The seven states are California, New York, Illinois, Michigan, Pennsylvania, Ohio and New Jersey (source: Bureau of Labor Statistics).
DEMOCRATS	1 if the Democratic Party won the most votes in the recent presidential elections of 2000, 2004, 2008, and 2012, 0 otherwise (source: http://www.uselectionatlas.org).
Control variables:	
PERF	CSR performance defined as the total positive (strengths) and negative (concerns) of six CSR rating categories (i.e., community, diversity, employee relations, environment, human rights and product) (source: KLD).
GOV	Corporate governance performance defined as the total of positive (strengths) and negative (concerns) features (source: KLD).
INS_OWN	Percentage of shares owned by insider investors (source: Thomson Reuters EIKON).
INST_OWN	Percentage of shares owned by institutional investors (source: Thomson Reuters EIKON).
SIZE	Natural logarithm of the market value of equity (source: Compustat).
AGE	Firm age, measured as the natural logarithm of the number of years the firm has been in Compustat.
TOBINQ	Tobin's Q defined as the sum of the market value of common equity plus the book value of preferred stock, book value of long-term debt and current liabilities over the book value of total assets (source: Compustat).
ROA	Return on assets, measured as the ratio of income before extraordinary items over total assets (source: Compustat).
LEV	Leverage ratio, measured as total debt over total assets (source: Compustat).
LIT	1 if the company is named as a defendant in a lawsuit, and 0 otherwise (source: Audit Analytics).
FIN	The amount of debt or equity capital raised by the firm over total assets. It is measured as the issuance of common stock and preferred shares minus the purchase of common stock and preferred shares plus the long-term debt issuance minus the long-term debt reduction (source: Compustat).
COMP	Herfindahl-Hirschman Index (HHI) multiplied by –1. HHI is calculated by summing the squares of the market shares of all companies in an industry. We calculate a firm's market share by dividing the sales of a firm by the total sales of all companies in an industry in that year, where industries are proxied using the two-digit SIC codes (source: Compustat).
GLOBAL	1 if the firm reports non-zero foreign income, 0 otherwise (source: Compustat).
LIQUIDITY	Ratio of the number of shares traded in the year to the total shares outstanding at the end of the year (source: Compustat).
Panel B: Additional covariates for PSM	
DCSRREPORTER	1 if a firm disclosed at least one CSR report according to Thomson Reuters ASSET4, CorporateRegister.com and CSRwire within our 14-year analysis window (2002–2015); 0 otherwise.
FRCF	Cash flow from operations minus cash flow used in investing activities. (source: Compustat).
SALES	Sales revenue over total assets. (source: Compustat).
PM	Income before extraordinary items over sales revenue. (source: Compustat).

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Table I Descriptive statistics of CSR reporting firms and non-CSR reporting firms prior to propensity score matching.

Variable	Entire Sample (N = 11,405)					CSR reporting firms (N = 3,372)					Non-CSR reporting firms (N = 8,033)					Std diffs (%)	Mean diff.
	25th	**Mean**	**Median**	**75th**	**StDev**	**25th**	**Mean**	**Median**	**75th**	**StDev**	**25th**	**Mean**	**Median**	**75th**	**StDev**		
DCSR	0	0.14	0	0	0.347	0	0.472	0	1	0.499	0	0	0	0	0	133.700	-0.472***
PCT_UNION_{t-1}	0	0.099	0	0.103	0.189	0	0.2	0.099	0.359	0.235	0	0.057	0	0.003	0.148	72.600	-0.142***
$UNION\_IND_{t-1}$	0.001	0.005	0.002	0.005	0.008	0.001	0.005	0.002	0.006	0.008	0.001	0.004	0.002	0.004	0.009	9.600	-0.001***
$PERF_{t-1}$	-1	-0.219	0	1	2.137	-1	0.488	0	2	3.046	-2	-0.516	-1	0	1.514	41.700	-1.004***
GOV_{t-1}	-1	-0.223	0	0	0.672	-1	-0.307	0	0	0.751	-1	-0.187	0	0	0.632	-17.300	0.120***
INS_OWN_{t-1}	0	0.006	0	0	0.036	0	0.003	0	0	0.023	0	0.007	0	0	0.04	-11.900	0.004***
$INST\_OWN_{t-1}$	0.683	0.788	0.847	0.943	0.194	0.715	0.808	0.853	0.935	0.165	0.668	0.78	0.843	0.946	0.204	14.800	-0.027***
$SIZE_{t-1}$	6.152	7.154	7.049	8.041	1.416	7.459	8.333	8.316	9.221	1.321	5.878	6.659	6.628	7.41	1.134	136.000	-1.674***
AGE_{t-1}	2.398	2.928	2.89	3.434	0.67	2.708	3.228	3.219	3.932	0.707	2.398	2.802	2.773	3.178	0.611	64.400	-0.426***
$TOBINQ_{t-1}$	1.106	2.019	1.545	2.401	1.422	1.012	1.768	1.397	2.11	1.16	1.15	2.124	1.612	2.535	1.506	-26.500	0.356***
ROA_{t-1}	0.009	0.023	0.045	0.085	0.152	0.024	0.05	0.048	0.085	0.075	-0.002	0.012	0.044	0.085	0.173	28.300	-0.038***
LEV_{t-1}	0.008	0.216	0.185	0.342	0.21	0.135	0.259	0.255	0.368	0.171	0	0.198	0.137	0.321	0.222	30.800	-0.061***
LIT_{t-1}	0	0.205	0	0	0.403	0	0.304	0	1	0.46	0	0.163	0	0	0.369	33.700	-0.141***
FIN_{t-1}	-0.039	0.018	0	0.034	0.139	-0.043	-0.001	-0.006	0.025	0.093	-0.038	0.026	0	0.039	0.154	-21.400	0.027***
$COMP_{t-1}$	-0.075	-0.067	-0.043	-0.032	0.064	-0.086	-0.073	-0.044	-0.031	0.071	-0.072	-0.065	-0.041	-0.032	0.06	-12.400	0.008***
$GLOBAL_{t-1}$	0	0.385	0	1	0.487	0	0.413	0	1	0.493	0	0.372	0	1	0.483	8.400	-0.041***
$LIQUIDITY_{t-1}$	1.245	2.551	2.026	3.239	1.939	1.386	2.729	2.134	3.428	2.034	1.185	2.476	1.982	3.184	1.893	12.900	-0.253***
$UNIONHALF_{t-1}$	0	0.423	0	1	0.494	0	0.417	0	1	0.493	0	0.426	0	1	0.495	-2.000	0.010
$DEMOCRATS_{t-1}$	0	0.613	1	1	0.487	0	0.593	1	1	0.491	0	0.621	1	1	0.485	-5.900	0.029*

This table presents the descriptive statistics for the sample prior to propensity score matching. Splits the sample into CSR reporting and non-CSR reporting groups. The last Column compares the differences in mean values of each variable across groups and the statistical significance of differences reported which are based on t-tests for continuous variables and chi-square tests for dummy variables. See Appendix for variable definitions. The standardized difference in percent is: $100(\bar{x}_{gr1} - \bar{x}_{gr0}/\sqrt{(s^2_{gr1} - s^2_{gr0})/2}$. Where: $\bar{x}_{gr1}$ and $\bar{x}_{gr0}$ $(s^2_{gr1} - s^2_{gr0})$ are the sample mean (variance) in the CSR-reporting and non-CSR reporting groups.

Note: Values with asterisks *, ** and *** indicate significance at the 10%, 5% and 1% levels respectively (2-tailed).

Table II Results of the first-stage logistic regression to model CSR reporting propensity.

	(1)
$PERF_{t-1}$	0.121***
	(8.05)
GOV_{t-1}	0.033
	(0.67)
INS_OWN_{t-1}	-0.596
	(-0.60)
$INST\_OWN_{t-1}$	-0.294
	(-1.49)
$SIZE_{t-1}$	1.111***
	(26.13)
AGE_{t-1}	0.209***
	(4.05)
$TOBINQ_{t-1}$	-0.372***
	(-10.77)
ROA_{t-1}	-0.461
	(-1.27)
LEV_{t-1}	0.600***
	(3.31)
LIT_{t-1}	0.065
	(0.89)
FIN_{t-1}	1.053***
	(3.44)
$COMP_{t-1}$	-0.614
	(-0.42)
$GLOBAL_{t-1}$	0.505***
	(5.68)
$LIQUIDITY_{t-1}$	0.101***
	(6.05)
$FRCF_{t-1}$	0. 00006***
	(6.77)
$SALES_{t-1}$	0.401***
	(6.44)
PM_{t-1}	0.193**
	(2.25)
(intercept)	-8.600***
	(-10.74)
Industry & Year Effects	Included
Wald χ^2	2,852.169
Pseudo R^2	0.433
Hosmer-Lemeshow χ^2	11.745
Area under the ROC curve	0.902
Mean VIF	1.369
Observations	11,405

Note: This table presents the logistic regression estimates of the first-stage regression of the propensity score matching approach, where the dependent variable is an indicator signaling that the firm issues at least one CSR report within our 14-year window (*DCSRREPORTER*). Standard errors are clustered at firm level with z-statistics presented in parentheses. Values with asterisks *, ** and *** indicate significance at the 10%, 5% and 1% levels respectively (2-tailed). See Appendix for variable definitions.

Table III Covariate balance of CSR-reporting firms and non-CSR reporting firms, prior and subsequent to propensity score matching.

Variable	Prior to propensity score matching						Subsequent to propensity score matching					
	CSR reporter (N = 3,372)		Non-CSR reporter (N = 8,033)		Std diffs (%)	Mean diff.	CSR reporter (N = 1,263)		Non-CSR reporter (N = 1,263)		Std diffs (%)	Mean diff.
	Mean	Median	Mean	Median			Mean	Median	Mean	Median		
DCSR	0.472	0	0	0	133.700	-0.472***	0.352	0	0	0	104.300	-0.352***
$PERF_{t-1}$	0.488	0	-0.516	-1	41.700	-1.004***	-0.267	0	-0.181	0	-4.300	0.086
GOV_{t-1}	-0.307	0	-0.187	0	-17.300	0.120***	-0.323	0	-0.321	0	-0.200	0.002
INS_OWN_{t-1}	0.003	0	0.007	0	-11.900	0.004***	0.006	0	0.005	0	4.800	-0.001
$INST\_OWN_{t-1}$	0.808	0.853	0.78	0.843	14.800	-0.027***	0.814	0.875	0.817	0.873	-2.100	0.004
$SIZE_{t-1}$	8.333	8.316	6.659	6.628	136.000	-1.674***	7.477	7.523	7.482	7.483	-0.500	0.005
AGE_{t-1}	3.228	3.219	2.802	2.773	64.400	-0.426***	3	2.944	2.999	2.944	0.100	-0.001
$TOBINQ_{t-1}$	1.768	1.397	2.124	1.612	-26.500	0.356***	1.936	1.437	1.947	1.554	-0.800	0.011
ROA_{t-1}	0.05	0.048	0.012	0.044	28.300	-0.038***	0.041	0.045	0.041	0.05	-0.600	0.001
LEV_{t-1}	0.259	0.255	0.198	0.137	30.800	-0.061***	0.229	0.221	0.234	0.221	-2.600	0.005
LIT_{t-1}	0.304	0	0.163	0	33.700	-0.141***	0.199	0	0.205	0	-1.600	0.006
FIN_{t-1}	-0.001	-0.006	0.026	0	-21.400	0.027***	0.007	-0.001	0.007	-0.002	0.100	0.000
$COMP_{t-1}$	-0.073	-0.044	-0.065	-0.041	-12.400	0.008***	-0.059	-0.042	-0.06	-0.041	0.600	0.000
$GLOBAL_{t-1}$	0.413	0	0.372	0	8.400	-0.041***	0.425	0	0.426	0	-0.200	0.001
$LIQUIDITY_{t-1}$	2.729	2.134	2.476	1.982	12.900	-0.253***	2.776	2.2	2.769	2.147	0.400	-0.008
$FRCF_{t-1}$	1202.18	461.034	129.95	56.115	70.800	-1072.228***	255.734	161.134	301.6	176.279	-8.800	45.860*
$SALES_{t-1}$	0.986	0.812	1.052	0.864	-9.000	0.065***	1.021	0.894	1.042	0.865	-3.100	0.021
PM_{t-1}	0.118	0.117	-0.085	0.084	26.500	-0.203***	0.086	0.103	0.082	0.106	1.100	-0.004
$UNIONHALF_{t-1}$	0.417	0	0.426	0	-2.000	0.010	0.426	0	0.412	0	2.900	-0.014
$DEMOCRATS_{t-1}$	0.593	1	0.621	1	-5.900	0.029*	0.587	1	0.593	1	-1.100	0.006

This table presents the descriptive statistics for the samples prior and subsequent to propensity score matching. It splits the sample into CSR-reporting and non-CSR reporting groups. The differences in mean values of each variable across groups and statistical significance of differences reported are based on t-tests for continuous variables and chi-square tests for dummy variables. See Appendix for variable definitions. The standardized difference in percent is: $100(\bar{x}_{gr1} - \bar{x}_{gr0}/\sqrt{(s^2_{gr1} - s^2_{gr0})/2}$. Where: $\bar{x}_{gr1}$ and $\bar{x}_{gr0}$ $(s^2_{gr1} - s^2_{gr0})$ are the sample mean (variance) in the CSR-reporting and non-CSR reporting groups.

Note: Values with asterisks *, ** and *** indicate significance at the 10%, 5% and 1% levels respectively (2-tailed).

Table IV Pearson correlation matrix (N = 2,526).

Variable	DCSR	1	2	3	4	5	6	7	8
1. UNION_PCT$_{t-1}$	0.09***	1.00							
2. UNION_IND$_{t-1}$	-0.02	0.10***	1.00						
3. PERF$_{t-1}$	0.11***	-0.17***	-0.14***	1.00					
4. GOV$_{t-1}$	0.03*	0.06***	0.06***	0.11***	1.00				
5. INS_OWN$_{t-1}$	0.00	-0.02	0.08***	-0.01	0.04**	1.00			
6. INST_OWN$_{t-1}$	0.04*	-0.18***	0.00	0.00	-0.19***	-0.13***	1.00		
7. SIZE$_{t-1}$	0.09***	-0.07***	-0.12***	0.15***	-0.20***	-0.04**	0.31***	1.00	
8. AGE$_{t-1}$	0.12***	0.30***	0.12***	-0.05***	0.02	-0.03*	-0.07***	0.04**	1.00
9. TOBINQ$_{t-1}$	-0.08***	-0.32***	-0.08***	0.16***	0.03	0.04**	0.12***	0.36***	-0.29***
10. ROA$_{t-1}$	-0.02	-0.05**	0.06***	0.01	-0.03	-0.03*	0.13***	0.26***	0.07***
11. LEV$_{t-1}$	0.03	0.27***	-0.04*	-0.08***	-0.01	-0.08***	-0.01	0.07***	0.07***
12. LIT$_{t-1}$	-0.01	-0.03	-0.01	0.02	-0.16***	0.00	0.06***	0.11***	-0.04**
13. FIN$_{t-1}$	-0.05**	0.02	-0.01	-0.03	0.10***	0.02	-0.05***	-0.07***	-0.14***
14. COMP$_{t-1}$	-0.05**	0.15***	-0.16***	-0.02	-0.02	0.01	-0.09***	0.06***	0.05***
15. GLOBAL$_{t-1}$	0.16***	-0.12***	-0.15***	0.09***	-0.06***	0.00	0.24***	0.17***	0.03
16. LIQUIDITY$_{t-1}$	0.00	-0.11***	-0.08***	0.05***	-0.23***	0.01	0.28***	0.14***	-0.16***
17. UNIONHALF$_{t-1}$	0.00	0.04**	0.01	0.04*	-0.08***	0.02	-0.03	0.06***	0.01
18. DEMOCRATS$_{t-1}$	0.08***	-0.02	-0.05**	0.00	-0.01	0.04**	0.02	0.03	0.05**
Variable	**9**	**10**	**11**	**12**	**13**	**14**	**15**	**16**	**17**
9. TOBINQ$_{t-1}$	1.00								
10. ROA$_{t-1}$	0.22***	1.00							
11. LEV$_{t-1}$	-0.23***	-0.12***	1.00						
12. LIT$_{t-1}$	0.00	-0.06***	0.00	1.00					
13. FIN$_{t-1}$	0.00	-0.24***	0.15***	-0.05**	1.00				
14. COMP$_{t-1}$	0.01	-0.05**	0.06***	-0.01	0.10***	1.00			
15. GLOBAL$_{t-1}$	0.02	0.03	-0.06***	-0.04*	-0.04**	0.02	1.00		
16. LIQUIDITY$_{t-1}$	0.13***	-0.10***	-0.07***	0.19***	0.01	-0.04**	0.14***	1.00	
17. UNIONHALF$_{t-1}$	0.01	-0.01	-0.07***	0.00	-0.02	0.06***	0.08***	0.04**	1.00
18. DEMOCRATS$_{t-1}$	-0.03	0.04*	-0.01	-0.01	-0.03	-0.02	0.22***	-0.04**	0.04*

Note: This table presents the Pearson correlation coefficients of the variables used in our analyses. Values with asterisks *, ** and *** indicate significance at the 10%, 5% and 1% levels respectively (2-tailed). See Appendix for variable definitions.

Table V Impact of unionization and the joint effect of unionization, geographical concentration of organized labor and political ideology on CSR reporting, logit analysis.

	(1) Firm-level unionization	(2) Industry-level unionization	(3) Interaction with UNIONHALF	(4) Interaction with DEMOCRATS
PCT_UNION_{t-1}	1.862***	.	1.042	0.185
	(3.41)		(1.53)	(0.24)
$UNION\_IND_{t-1}$	.	13.950	.	.
		(1.23)		
$PCT\_UNION_{t-1} \times UNIONHALF_{t-1}$	.	.	1.829**	.
			(2.02)	
$UNIONHALF_{t-1}$	.	.	-0.277	.
			(-1.08)	
$PCT\_UNION_{t-1} \times DEMOCRATS_{t-1}$	.	.	.	2.626***
				(3.04)
$DEMOCRATS_{t-1}$	.	.	.	-0.434*
				(-1.92)
$PERF_{t-1}$	0.104**	0.093*	0.111**	0.108**
	(2.12)	(1.90)	(2.20)	(2.15)
GOV_{t-1}	0.115	0.119	0.102	0.108
	(0.91)	(0.93)	(0.80)	(0.85)
INS_OWN_{t-1}	1.104	0.919	1.236	1.303
	(0.49)	(0.42)	(0.54)	(0.56)
$INST\_OWN_{t-1}$	-0.114	-0.257	-0.064	0.021
	(-0.19)	(-0.43)	(-0.11)	(0.03)
$SIZE_{t-1}$	0.258**	0.284**	0.267**	0.244**
	(2.22)	(2.48)	(2.29)	(2.07)
AGE_{t-1}	0.267	0.311*	0.266	0.259
	(1.61)	(1.84)	(1.61)	(1.53)
$TOBINQ_{t-1}$	-0.223**	-0.272***	-0.221**	-0.215**
	(-2.17)	(-2.58)	(-2.16)	(-2.06)
ROA_{t-1}	-0.868**	-0.880**	-0.905**	-0.853*
	(-2.03)	(-2.13)	(-2.04)	(-1.90)
LEV_{t-1}	-0.444	-0.262	-0.446	-0.436
	(-0.84)	(-0.51)	(-0.85)	(-0.82)
LIT_{t-1}	0.137	0.114	0.129	0.130
	(0.84)	(0.70)	(0.79)	(0.81)
FIN_{t-1}	-1.811***	-1.772***	-1.831***	-1.837***
	(-2.86)	(-2.81)	(-2.88)	(-2.85)
$COMP_{t-1}$	-2.754	-2.636	-2.489	-2.877
	(-0.93)	(-0.91)	(-0.81)	(-0.94)
$GLOBAL_{t-1}$	-0.104	-0.109	-0.095	-0.022
	(-0.42)	(-0.43)	(-0.38)	(-0.09)
$LIQUIDITY_{t-1}$	0.010	0.029	0.004	0.001
	(0.24)	(0.61)	(0.09)	(0.02)
(intercept)	-4.628***	-4.901***	-4.647***	-4.417***
	(-3.76)	(-4.07)	(-3.76)	(-3.53)
Industry & Year Effects	Included	Included	Included	Included
Wald χ^2	262.245	268.882	264.101	264.015
Pseudo R^2	0.188	0.178	0.192	0.195
Hosmer-Lemeshow χ^2	6.662	6.083	9.456	9.958
Area under the ROC curve	0.799	0.792	0.801	0.804
Mean VIF	1.212	1.190	1.338	1.383
Observations	2,526	2,526	2,526	2,526

Note: This table presents the estimates of a logistic regression, where the dependent variable is *DCSR*. Columns 1 and 2, respectively, provide estimates for firm- and industry-level unionization proxies. In Column 3 we interact *PCT_UNION* with an indicator signaling that firm headquarters in a state, where more than half of all union members live (*UNIONHALF*). In Column 4, we interact *PCT_UNION* with an indicator variable that equals one if the Democratic Party won the most votes in the presidential elections of 2000, 2004, 2008, and 2012 (*DEMOCRATS*). Standard errors are clustered at firm level with z-statistics presented in parentheses. Values with asterisks *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively (2-tailed). See Appendix for variable definitions.

Figure I Interaction effect of unionization and states where half of the total unionized labor force of the US resides on CSR reporting

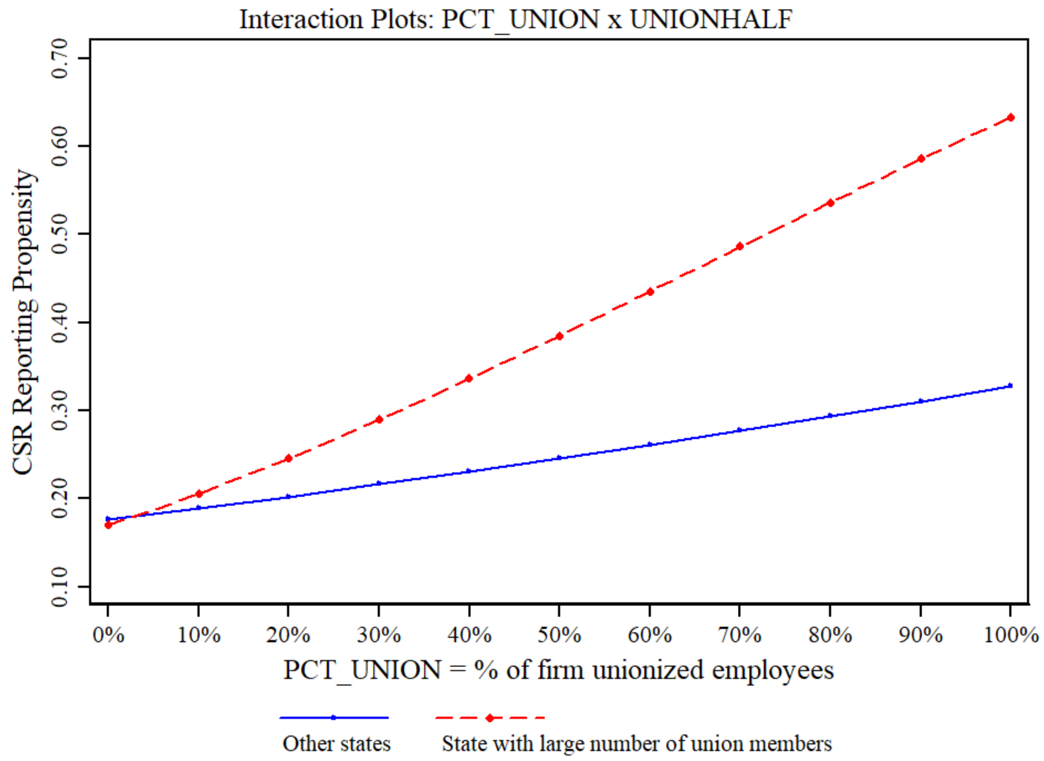

Figure II Interaction effect of unionization and dominance of the Democratic Party at presidential elections on CSR reporting