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Magnus Schauf
Eline Schoonjans

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Better Safe than Sorry: Toxic Waste Management after Unionization

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Abstract

Do unions protect workers from releasing or handling toxic waste? This paper studies the impact of organized labor on toxic waste management at US facilities between 1991 and 2020. If unions, as collective voice, bargain for worker benefits such as workplace safety and member health, their effect on toxic releases remains unclear due to a tradeoff. Reducing toxic waste releases has positive health and environmental effects but requires more and dangerous activities to handle waste after production. Using a regression discontinuity design on close-call union elections, we find a significantly negative effect of unionization on the sum of toxic waste recycling, energy recovery, and treatment at the facility site. In contrast, total toxic releases to air, land, and water increase after unionization. These effects are more pronounced in states without right-to-work laws, for less toxic chemicals, and for non-heavy industries. Finally, we show that unionized facilities increase waste prevention activities through innovative product and process modifications and have less catastrophic releases. However, these effects cannot offset the reduction in waste handling, resulting in more waste releases. Our findings suggest that unions prioritize safety over sustainability and call upon managerial and governmental action to better align these two objectives.

*Schauf: Technical University of Munich, School of Management. Email: magnus.schauf@tum.de. Schoonjans: Technical University of Munich, School of Management. Email: eline.schoonjans@tum.de. We thank Stefan Bauernschuster, Matthias Doepke, Gunther Friedl, Hanna Hottenrott, Regina Riphahn, Peter Schäfer, Hannes Schwandt, Sebastian Schwenen, and seminar participants at the 33rd BGPE workshop, Northwestern University, and Technical University of Munich for helpful comments.

1 Introduction

Each year, US manufacturing facilities produce about 30 billion pounds of hazardous chemical waste, 10% of which end up in the environment through air, land, and water releases (EPA, 2022). Such toxic chemical releases have a multitude of negative effects on human health, our planet, and the economy.

Workers represent a stakeholder group highly affected by toxic waste (Dietz et al., 2015). On the one hand, workers are exposed to toxic chemical releases at the workplace and possibly as members of local communities. On the other hand, workers are also most affected by performing dangerous and costly tasks that avoid releases of these chemicals. Hence, workers face a sustainability-safety tradeoff between protecting themselves, the community, and the society at a larger scale from releasing toxic waste and protecting themselves from handling toxic waste.

While mechanisms are often not well understood (Ertugrul and Marciukaityte, 2021), recent research suggests that workers' interests and the interests of other stakeholders like society or nature diverge (Ertugrul and Marciukaityte, 2021; Heitz et al., 2021). Yet, little research causally examines whether pollution that is particularly detrimental on a local level (Currie et al., 2015) leads to better alignment of interests between workers and society. Toxic waste management provides a context to disentangle mechanisms by directly comparing qualitatively different worker benefits. Although some benefits also improve welfare for other, especially local, stakeholders, it is unclear which objectives unions prioritize.

This paper empirically studies how a shift in stakeholders' bargaining power toward workers following unionization affects the sustainability-safety tradeoff and associated toxic waste externalities. Specifically, we test whether facilities rather release (less sustainability, more safety) or "cure", i.e. handle, (more sustainability, less safety) their toxic waste. Toxic waste cure refers to end-of-production procedures, which include recycling, energy recovery or other waste treatment. Releases entail chemical emissions and disposals to the environment, i.e. air, land, and water.

We formulate the sustainability-safety tradeoff as a bidirectional hypothesis. Unionization could lead to less releases and more cure or more releases and less cure.

The *collective voice* view stipulates that unionized workers bargain more easily for their interests than non-unionized workers (Freeman and Medoff, 1979). Basic interests are securing labor and reasonable pay (McDonald and Solow, 1981). Both translate into a distaste for waste cure which is often costly (King and Lenox, 2002; Frondel et al., 2007). Saving these costs could permit unionized workers to extract part of the rent surplus in the form of higher pay or better employment conditions. Managers also have cost saving incentives if labor costs increase after unionization.

Furthermore, labor unions' interests extend to additional dimensions of member welfare, such as workplace safety and worker health (Freeman and Medoff, 1979). Given the high risk for fatal injuries when performing waste cure activities (BLS, 2021), pursuing workplace safety could also align with decreasing waste cure. On the contrary, welfare maximization arguably implies protection from pollution and its negative effects at the workplace and in local communities, be it for selfish or altruistic motivations. Moreover, since environmental protection can represent an investment in human capital, which is stickier following unionization (McDonald and Solow, 1981), managers might strive to reduce toxic releases to increase their workers' health and thereby productivity (Graff Zivin and Neidell, 2012). This reduction can be achieved with more cure but also by means of other pollution prevention activity.

Consequently, we argue that innovation activities and better training could align interests between different stakeholder groups and relax workers' sustainability-safety tradeoff. In light of this multi-win character, we hypothesize that such activities should increase following unionization. Stimulating innovation environments require short-term failure tolerance and long-term reward mechanisms (Manso, 2011). Unionization facilitates the former through better employment protection while the reduction of negative waste effects might represent sufficient long-term motivation. Whether and to what extent unions can realize all of these interests depends on their bargaining power which can be restricted by, i.a., right-to-work (RTW) legislation (Chava et al., 2020).

For estimation, we rely on a local regression discontinuity design (RDD) using union elections that narrowly pass or don't pass the majority cutoff of 50 percent plus one vote. The local RDD establishes causality under verifiable assumptions, addressing endogeneity concerns of unobservable differences between unionized and non-unionized facilities (Hahn et al., 2001; Lee and Lemieux, 2010). We consider chemical-facility-level data reported annually by manufacturing facilities to the Toxic Release Inventory (TRI), a program of the US Environmental Protection Agency (EPA). By means of thoroughly cleaned facility names, we match this detailed source of toxic waste management data to US union elections between 1990 and 2017. Our final sample contains 5,583 chemical-facility-year observations related to 605 union elections.

Our main result provides evidence of a shift in waste management practices after a shift in stakeholder power toward unionized workers. Specifically, we find that, on average over three years after unionization, releases increase and cure decreases at facilities with close-call union wins. The effect is statistically significant and economically large for *on-site* release and cure in particular, i.e. at the location where unions have most interest in. Changes in on-site releases (cure) are 14.7 (59.3) percentage points higher (lower) in facilities with a marginal union victory compared to facilities with marginal union losses, on average. We do not find significant evidence of "outsourcing" releases or cure to other facilities (off-site). As such, our findings indicate unions' preference for safety when facing a sustainability-safety tradeoff.

This main effect survives a battery of contextual and RDD-specific robustness tests and remains significant in a global RDD using all elections. In cross-sectional analyses, we examine whether union power as determined by the presence of RTW laws, chemical toxicity, and industry affiliation play a role for the unionization effect. We find that facilities reduce cure especially in non-RTW states where unions have higher bargaining power and for less toxic chemicals. Similarly, facilities from non-heavy industries primarily drive the observed decrease in waste cure. We argue that operative flexibility and worker expectations might explain this heterogeneity.

Next, we probe into two possible explanations behind our main effect. Specifically, we investigate the impact of production output and financial constraints. First, consistent with DiNardo and Lee (2004), we do not find a significant effect of unions on productivity and rule out that changes in production output explain increasing releases or decreasing cure. Second, since waste cure is costly, financially constrained facilities might not be able to afford waste cure. However, we also rule out that financial constraints drive our results. This finding emphasizes the sustainability-safety tradeoff and unions' welfare extraction rather than purely monetary motivations as drivers behind the observed changes in waste management practices.

Lastly, we investigate two mechanisms, catastrophic releases and innovative pollution prevention activity, that could relax the sustainability-safety tradeoff. We show a decrease in catastrophic releases which better training might explain. Moreover, we show an increase in innovative pollution prevention activity. Together, these findings suggest that unionization motivates facilities to focus on multi-win outcomes.

Our paper contributes to the literature in several ways. First, we identify labor unions as significant drivers of facilities' environmental performance in general and waste management strategy in specific. Other important determinants of firms' environmental performance entail environmental and liability regulation (Shapiro and Walker, 2018; Akey and Appel, 2021), pressure by community stakeholders (Kassinis and Vafeas, 2006), financial constraints

(Dutt and King, 2014; Cohn and Deryugina, 2018; Xu and Kim, 2022), corporate governance (Kim et al., 2019; Shive and Forster, 2020), and innovation (King and Lenox, 2002). We add to this literature by highlighting that unions represent an obstacle in reducing environmental impact because of their lower willingness for curing toxic chemicals.

Second, we contribute to the literature on real effects of unionization. Previous literature reports unionization effects on investment (Connolly et al., 1986; Machin and Wadhwani, 1991), innovation (Haucap and Wey, 2004; Bradley et al., 2017), and corporate governance (DeAngelo and DeAngelo, 1991; Chyz et al., 2013; Chung et al., 2016). While the evidence is sometimes mixed, a commonly shared conclusion is that unions and their members extract welfare at the expense of shareholders. A particularly recent stream investigates the effect of unionization on other stakeholders, e.g. using environmental, social, and governance scores (Ertugrul and Marciukaityte, 2021; Heitz et al., 2021) or product recalls (Kini et al., 2021). These studies corroborate the general consensus of unionization being detrimental to other stakeholder groups. We add to this literature by exploiting the various advantages that the context of waste management offers. Most importantly, unionized workers in our setting face a tradeoff between being exposed to pollution or to dangerous jobs. As such, we can study a setting where negative externalities from pollution take place and partially realize at the same level as unionization and decisions on waste management: the facility. Moreover, we establish our effects using quantity-based measures rather than methodology-based and often controversial ESG scores (Chatterji et al., 2016; Berg et al., 2022).

Overall, we show that real effects of unionization extend to toxic waste management and provide new insights into how unionized workplaces operate on the local facility level. Our results highlight that the utility of being in a labor union goes beyond purely monetary incentives. Finally, the tendency for unions to value safety over sustainability highlights the need for policymakers and managers to increase standards of and trust in waste management practices while augmenting efforts toward multi-win outcomes, e.g. via green innovation and training.

The rest of this paper is organized as follows. Section 2 develops our hypothesis from reviewing the literature on unionization in the context of workplace and environmental protection. Section 3 describes our data and econometric approach. Section 4 presents main results, Section 5 shows cross-sectional heterogeneity, and Section 6 investigates possible underlying mechanisms. Section 7 concludes.

2 Hypothesis Development

This section discusses the role of labor unions in influencing environmental externalities generally and toxic waste management at manufacturing firms specifically. As the organized voice of individual workers, labor unions assist in internalizing employment-related external costs and reduce transaction costs through *collective bargaining* (Freeman, 1976). They directly lobby for their interests when bargaining with managers or when interacting with policymakers. Unions' main interests are employment and employment conditions, such as wages, job security, and safety (Freeman and Medoff, 1979; McDonald and Solow, 1981).

2.1 Sustainability-Safety Tradeoff

Empirical evidence on whether unions' interests align with those of other stakeholders beyond their bargaining unit, including the general workforce, communities, and society at large, is somewhat mixed. A particularly recent stream investigates the effect of unionization on non-shareholding stakeholders, e.g. by means of environmental, social, and governance scores (Ertugrul and Marciukaityte, 2021; Heitz et al., 2021) or product recalls (Kini et al., 2021). They find that unions extract welfare for their members at the expense of external stakeholders. However, anecdotal evidence shows union support for environmental protection policies, despite their potential negative impact on member jobs. In the 1950's and

1960's unions lobbied for environmental regulation to reduce air, land, and water pollution (Dewey, 1998). The passing of these regulations (Clean Air Act and Clean Water Act) led to significant costs for workers due to lower wages and unemployment (Walker, 2011, 2013). Yet, their positive impact on member and community health seemed to outweigh these costs, as workers and their families are arguably most exposed to detrimental effects of pollution on the local level (Dietz et al., 2015).

Health effects resulting from contamination with or exposure to pollution, as well as resulting economic effects, are empirically well-documented. Adverse inter-generational health impacts (for a review, see Graff Zivin and Neidell, 2013), including pronounced elderly mortality (Deryugina et al., 2019), higher respiratory and heart-related hospital admissions (Schlenker and Walker, 2016), and increased probability of low birthweight (Currie et al., 2015), lead to ripple effects throughout the whole economy: Housing prices decrease (Currie et al., 2015), workers lower labor supply and become less productive (Hanna and Oliva, 2015; He et al., 2019), investors trade worse (Huang et al., 2020), and managers and their human capital migrate away from polluting plants (Levine et al., 2019).

Consequently, pollution from releases of toxic chemical waste from manufacturing facilities directly impact worker well-being. Workers are exposed to pollution directly at the workplace and potentially also as inhabitants of neighboring communities (Currie et al., 2015; Dietz et al., 2015). Put differently, a firm's toxic releases essentially reduce workers' welfare non-monetarily through its negative health impacts and monetarily because workers might spend more on pharmaceuticals or other "defensive investments" to mitigate pollution effects (Deschênes et al., 2017). Therefore, a union that maximizes member welfare should also pursue sustainability through lower toxic releases. Moreover, if unionization leads to costlier and stickier human capital (McDonald and Solow, 1981), managers can have incentives to reduce releases and increase workers' productivity. As such, environmental protection is an investment in human capital (Graff Zivin and Neidell, 2012).

However, industrial ecology scholars argue that workers partially carry the burden of lower toxic releases to the environment through higher exposure and higher risks at their jobs, without reasonable accounting for associated health and safety effects (Ashford, 1997; Armenti et al., 2011). Workers perform waste management jobs, namely recycling, use for energy recovery and other treatment that typically destroys the chemical in order to reduce releases. These waste management practices are usually the final step in the production process and we refer to them collectively as waste "cure". Official statistics for workplace safety underscore that unions could directly bargain for less waste cure because it is dangerous. For instance, "refuse and recyclable material collector" is the sixth most fatal work-injury related occupation with 33 fatalities per 100,000 full-time equivalent workers, which is ten times higher than the average over all occupations (BLS, 2021). Recent empirical studies report a downward shift in the distribution of accident-case rates and injuries after unionization (Zoorob, 2018; Heitz et al., 2021; Li et al., 2022), supporting the argument that unions address workplace safety. Potential channels might be better training of employees (Heitz et al., 2021) and protection from dangerous tasks such as waste cure.

Besides concerns for member safety, unionization might lead to less waste cure because of cost saving incentives. Infrastructure to cure waste is costly to install and operate (Frondel et al., 2007). Further, King and Lenox (2002) argue that on-site waste cure often results in unexpected additional costs. Finally, unionization could indirectly lead to less waste cure through greater employment protection, inducing "shirking" or a less thorough work attitude (Bradley et al., 2017).

Manufacturing facilities can also cooperate with special third-party waste treatment plants to relocate waste before managing it. Since such off-site waste cure essentially avoids unionized workers on-site to cure waste, our reasoning above applies for on-site waste cure in particular.

Taken together, we posit that workers face a tradeoff between increasing sustainability (through lower toxic releases) and increasing safety (through less dangerous waste cure practices). Since unionization shifts the weights of stakeholder interests, workers should have more power to impact the sustainability-safety tradeoff according to their interests. We hypothesize that it is ex ante unclear whether these interests relate to less releases but more cure or more releases but less cure. While we so far assumed the decision at stake to be sustainability *versus* safety, we next argue how facilities could relax this tradeoff.

2.2 Relaxing the Sustainability-Safety Tradeoff

Several mechanisms exist to reduce both the need for waste release and cure. First, a mechanical lever for unions to decrease waste that needs to be managed, is by decreasing production output. In light of stronger employment protection, previously mentioned "shirking" behavior might not only refer to unpleasant waste cure but also extend to other, production-related, job duties (Bradley et al., 2017). On the contrary, unionization can enhance productivity through its bundling of worker interests and reduction of transaction costs (Freeman, 1976). Empirical evidence reports little changes in productivity due to unionization (DiNardo and Lee, 2004)

Second, a theoretically more appealing mechanism is to increase investment in source reduction or pollution prevention activities such as training, innovation, and improvement of operations. Traditional hold-up theory would predict investments to decline after unionization because managers anticipate excessive rent capturing by unions (Grout, 1984; Connolly et al., 1986). However, in our setting the character of these investments is different. Firm investment in pollution prevention implies positive externalities (or reduced negative externalities) that directly benefit workers through higher workplace safety, less need for working on waste cure, and lower chemical releases. As such, firms can share non-monetary benefits from these investments with their workforce. These benefits could represent long-term rewards which, combined with short-term failure tolerance of union contracts, often stimulate innovation activity (Manso, 2011).

Third, pressure due to the unionization shock on labor costs can also play a role. Higher labor costs might induce managers to save costs. Reportedly, cost savings are the major influencing factor for waste prevention earlier during the production process, rather than end-of-pipe waste cure (Frondel et al., 2007). Moreover, waste prevention encourages the development of worker skills, which in turn leads to higher financial performance (King, 1999; King and Lenox, 2002). Nevertheless, King and Lenox (2002) argue that managers have systematically overlooked innovative waste prevention as a green and profitable opportunity. Taken together, the multi-win character of pollution prevention investment should dominate managerial hold-up concerns. Consequently, we expect that unionization relaxes the sustainability-safety tradeoff through innovative pollution prevention policies and less failures in operations.

3 Methodology and Data

3.1 Data and Sample

Our sample consists of industrial facilities that are mandated to report toxic waste by chemical to the US Environmental Protection Agency (EPA) and had a union election between 1990 and 2017. Specifically, we use data of hazardous chemical usage and waste submitted to EPA's Toxic Release Inventory (TRI) under the Emergency Planning and Community Right to Know Act (EPCRA) of 1986. The EPCRA and subsequent amendments require a facility to report chemical-level data for currently about 600 chemicals. TRI data items include chemical information, production waste quantities, waste management practices, and the geographic location of these management practices (on-site versus off-site).

We add toxicity data, namely reportable quantities for chemicals according to the Comprehensive Environmental Response, Compensation, and Liability Act, using the Chemical Abstracts Service Registry Number (CAS) for merging (King and Lenox, 2000, 2002; Kim et al., 2019). Reportable quantities represent the threshold value in pounds that determines necessary emergency reporting and action in case of an accidental spill. These quantities are in a range of 1 to 5,000, with the lowest value of 1 indicating a highly toxic chemical for which spilling one pound or more already requires emergency action. We also merge in alternative chemical-level toxicity data, namely the human toxicity potential published by Hertwich et al. (2001). By providing separate values for air versus water exposure routes, these data enable multiple robustness tests with toxicity explicitly focusing on human health effects.[1]

Although not free of flaws,[2] the TRI does not seem to be subject to systematically manipulated reporting (Bui and Mayer, 2003). Moreover, the EPA checks submitted reports and can fine misreporting facilities. Consequently, academic literature exploits the rich TRI data in many different applications (e.g. King and Lenox, 2002; Currie et al., 2015; Shapiro and Walker, 2018; Akey and Appel, 2021). Following previous literature (Dutt and King, 2014; Akey and Appel, 2021), we exploit the empirical advantages of this detailed chemical-level data, e.g. more precise estimates of toxic waste quantities, more precise controls for changes in production output at multi-product facilities, and accounting for chemical fixed effects in further analyses.

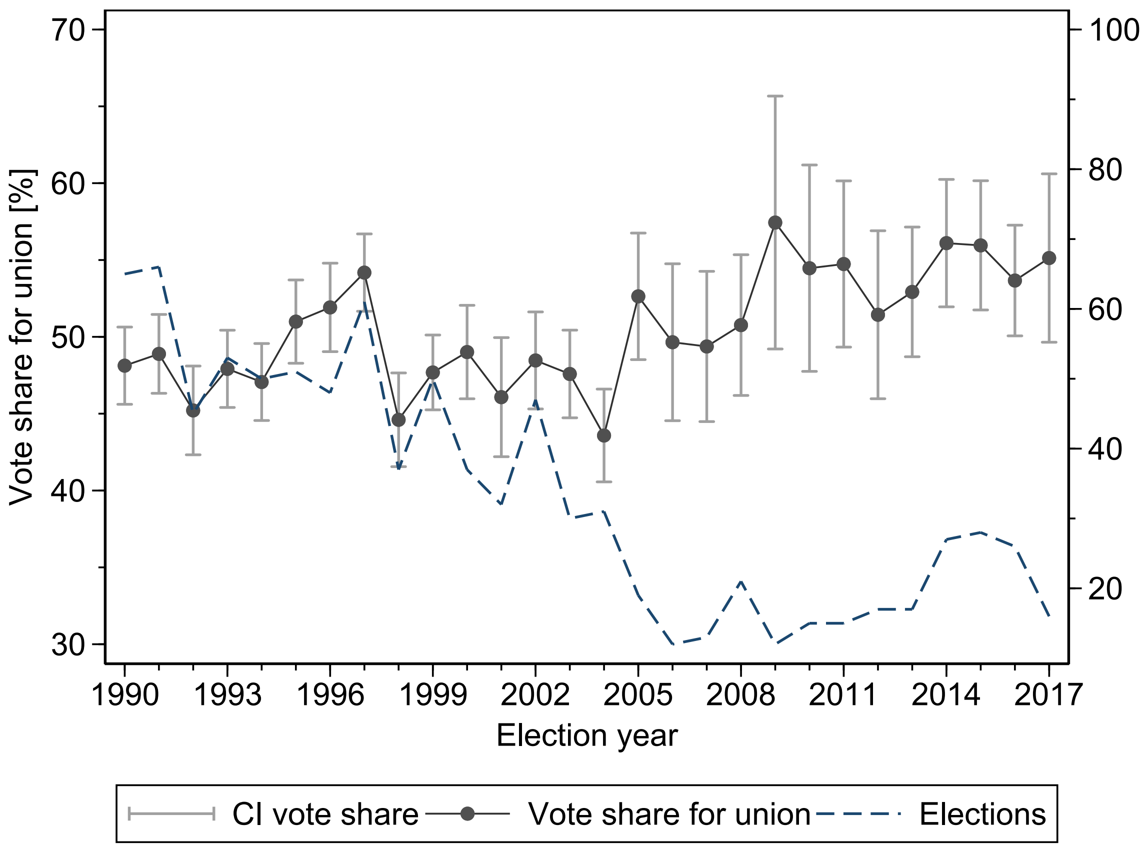
The US authority supervising and validating facility-level union elections, the National Labor Relations Board (NLRB), provides union election data. Per election case, the data contain facility and union information as well as election characteristics, most importantly the number of eligible voters and the valid votes pro and contra joining the respective union.

[1]Further, by means of the CAS, we supplement our TRI data with a chemical's listing period in the TRI, obtained from the EPA. In robustness tests, we exclude chemicals based on their listing period not covering our entire sample period to control for regulatory changes (see Section 4.3).

[2]These flaws relate to the self-reported nature of the data and inconsistency over time due to frequent regulatory changes (e.g. Currie et al., 2015).

We only keep certification cases, i.e. elections for union representation at a facility. Figure 1 shows the trends in union elections over our sample period. Consistent with other studies (Bradley et al., 2017; Campello et al., 2018; Heitz et al., 2021; Kini et al., 2021) the number of elections gradually declines while the vote share, i.e. the success rate of unionization, fluctuates more, with a tendency to increase in more recent years.

Figure 1: Number of Union Elections and Vote Shares by Election Year



Notes: This figure shows trends in union elections from 1990 till 2017, certified by the NLRB and matched to the TRI data. The dashed line represents the number of elections certified by year. The line with dots depicts the average vote share in a year and the gray vertical lines indicate the 95% confidence interval for this average vote share.

Since NLRB and TRI data do not share a common identifier, we must string-match facility names. Before, we perform a meticulous string cleaning procedure in order to increase the number and quality of exact and fuzzy matches. Our string cleaning algorithm harmonizes facility names across both data sets for instance by adjusting abbreviations or removing legal forms.[3] We are careful to ensure accurate matches by requiring that the facilities' states and cities always match. We complement all exact matches with manually checked fuzzy matches of distance 1 or 2 as calculated by the Optimal String Alignment method.

In order to arrive at our main sample, some additional cleaning steps are necessary. First, as small union elections are less suitable for our research design (DiNardo and Lee, 2004; Frandsen, 2017), we exclude elections with less than 50 eligible voters.[4] Second, we only include observations which we can merge to the National Establishment Time Series (NETS) using the TRI facility identifier. NETS, a commercial database by consultancy Walls & Associates based on data from Dun & Bradstreet, provides facility-level sales, employees, and the "Paydex" business credit score, amongst others. These variables allow for testing one of the key identifying assumptions of our empirical design more thoroughly (see Section 3.3). Our main sample contains 605 unique union elections and 5,583 chemical-facility-year observations consisting of 283 different chemicals.

3.2 Variables

Having introduced our data sources and sample, we next briefly describe variables relevant for estimating our effects and for testing the validity of our design.

In our main analysis, we estimate direct and medium-term effects of unionization on several measures of waste management at the facility-chemical-year level.[5] We consider waste management variables up to three years post union election, because the certification and bargaining process as well as potentially resulting real effects are not immediate. First, we calculate releases as the sum of chemical releases to air, land, and water, on-site, off-site,

[3] Similar, though sometimes more compact string cleaning algorithms are commonly used in the literature, including for TRI and unionization data (Lee and Mas, 2012; Akey and Appel, 2021; Xu and Kim, 2022).

[4] Smaller elections can be more easily manipulated during the election and seem to occur in a more selective way given their notably higher vote share.

[5] We eliminate some rare cases of partial facility reports where two or more entries per chemical exist for a given year and facility by aggregating the individually reported quantities.

and in total. Second, we construct waste cured as the sum of waste that is recycled, used for energy recovery, or otherwise treated. Again, this measure can be a total as well as separated for on-site versus off-site. Importantly, we process waste releases and waste cure to ratios that indicate the change in the outcome variable on a year-by-year basis. The benefits of using ratios instead of level variables are threefold. First, we deskew our outcome variables. Second, we improve the accuracy and efficiency of our regression estimates, compared to alternative log-transformations (Cohn et al., 2022). Third, we facilitate the comparability of otherwise very heterogeneous chemical quantities and facility sizes (Dutt and King, 2014).

We formally define the change in the outcome variable as

$$\Delta Y_{i,j,t} = Y_{i,j,t}/Y_{i,j,t-1}, \tag{1}$$

where Y represents waste cure or releases, on-site, off-site, or in total, in pounds of chemical i at facility j in year t (Dutt and King, 2014). We exclude observations with ratios exceeding a threshold of three as these are either data errors or evidence of extraordinary transformations at the facility (Akey and Appel, 2021). In robustness checks, we adjust this threshold to either two or four and also take the natural log to further reduce the impact of extreme values (Dutt and King, 2014).

Throughout our analysis, the vote share, i.e. the number of votes in favor of the union divided by the total number of valid votes cast, constitutes our main independent or "running" variable. A vote share of 50% plus one vote implies a successful union election and thus establishes treatment in our analysis. Since vote share has finite, discrete, and asymmetric support, which changes with the number of votes cast, we adjust the vote share such that 50% constitutes the homogeneous and symmetric cutoff. For all even numbers of votes cast, we subtract an amount equal to ${}^{0.5}/_{\text{number of votes cast}}$ from the vote share. Cases with odd numbers of votes cast are not adjusted (DiNardo and Lee, 2004).

In our further analyses and for testing one identifying assumption of our empirical approach, we require additional facility-level and chemical-level variables. Facility-level variables encompass sales, number of employees, and the Paydex score from NETS (Akey and Appel, 2021). The Paydex score is a business credit score between 0 and 100, calculated from previous transaction data. A high score indicates a high probability for the focal firm to pay back its debt on time. Chemical-facility-level variables include information on (i) production output, (ii) pollution prevention measures, also termed "source reduction activities", and (iii) catastrophic releases from TRI.

First, we proxy production output with the production ratio that facilities report to the TRI. The production ratio indicates changes in the output of a manufactured product or a supporting operational procedure related to the use of a specific chemical. Following Akey and Appel (2021) and the construction of our dependent variables, we exclude observations with ratios larger than three. Second, to measure facilities' engagement in chemical-level pollution prevention, we construct an indicator variable from the approximately 50 codes that facilities can report as performed pollution prevention activity. This variable, "innovative prevention", is equal to one if a facility, for a given chemical, reports prevention activities related to product and process innovations and zero otherwise.[6] Finally, we examine catastrophic releases because we assume that better trained employees commit less mistakes leading to such releases.[7] Given the rarity of catastrophic releases, we construct an indicator variable equal to one if a facility reported catastrophic releases of a specific chemical and zero if not.

We describe all variables in Appendix Table 1. Appendix Table 2 contains summary statistics and Appendix Figure 1 illustrates spatial heterogeneity across US states for several

[6] This innovation measure captures TRI reporting codes W50, W51, W59, W80, W82, W83, W84, W89.

[7] Catastrophic releases refer to a "major uncontrolled emission, fire, or explosion, involving one or more highly hazardous chemicals, that presents serious danger to employees in the workplace" (OSHA, 2012). In about 2% of facility-chemical-year observations, facilities report a catastrophic release.

variables characterizing our sample.

3.3 Empirical Strategy

We identify the causal effect of unionization on waste management practices at manufacturing firms by means of a regression discontinuity design (RDD). Due to endogeneity, a simple OLS regression could provide inconsistent and biased estimates. Endogeneity arises when unionization depends on factors that also influence our variables of interest. For example, workers at facilities that have unsatisfying and dangerous waste management practices or other unobservable grievances may be more likely to unionize, which could imply over- or underestimation of the unionization impact.

The RDD establishes causality and exploits the election character of unionization, where a simple majority separates the treatment (i.e., successful union certification) from the control group (DiNardo and Lee, 2004). Specifically, considering close-call union elections around the majority cutoff represents a quasi-experimental approach which supports causal interpretation of the local treatment effect.

Formally, in our main empirical analysis, we estimate the following non-parametric local-linear equation (Lee and Lemieux, 2010):

$$\Delta Y_{i,j,t+n} = \alpha + \tau D_{j,t} + \beta_l (X - c) + (\beta_r - \beta_l) D_{j,t} (X - c) + \varepsilon_{i,j,t+n}, \tag{2}$$

where Y represents waste cured or released in pounds of chemical i at facility j in year $t+n$, with $n \in [1, 2, 3]$. Our running variable X, adjusted vote share, splits our observations into treatment ($D = 1$) and control ($D = 0$) groups at the cutoff c of 50%. We include observations, where X is within the mean-squared-error optimizing (MSE-optimal) bandwidth h on each side of the cutoff (Imbens and Kalyanaraman, 2012). For statistical inference, we correct for the estimation bias with bias-corrected confidence intervals and robust standard errors clustered at the facility level (Calonico et al., 2014, 2019). We choose a triangular kernel for weighting observations, but test robustness of our results to alternative choices.

In order to qualify as an econometric method estimating causal effects, an RDD requires two identifying assumptions: (i) ex ante comparability between treatment and control group observations as well as (ii) exogeneous election outcomes that no party can manipulate (Hahn et al., 2001; DiNardo and Lee, 2004; Lee and Lemieux, 2010). Table 1 shows the results for testing the balance of dependent variables and covariates in the vote share bandwidth of 9% around the cutoff. This bandwidth corresponds to the average MSE-optimal bandwidths of our main analyses. Most importantly, our main dependent variables, changes in (on-site) waste releases and cure are balanced in treatment and control group prior to the union election. Both treatment and control group are also comparable in terms of adoption of pollution prevention activities, sales, employees, and financial constraints, amongst others. Prior to unionization, we observe a significantly lower production ratio among facilities that ultimately unionize. Consequently, we explicitly control for the production ratio in further analyses.

Next, we perform diagnostic tests for the second identifying assumption, absence of manipulation. First, Figure 2 illustrates the distribution of vote share in favor of joining the union in 2% steps, i.e. 50 bins. As can be seen, the density of unions that fail is generally larger but decreases closer to the cutoff. This decline continues further on the treatment side of the cutoff. Overall, vote share is fairly distributed across the range of 0% to 100%.

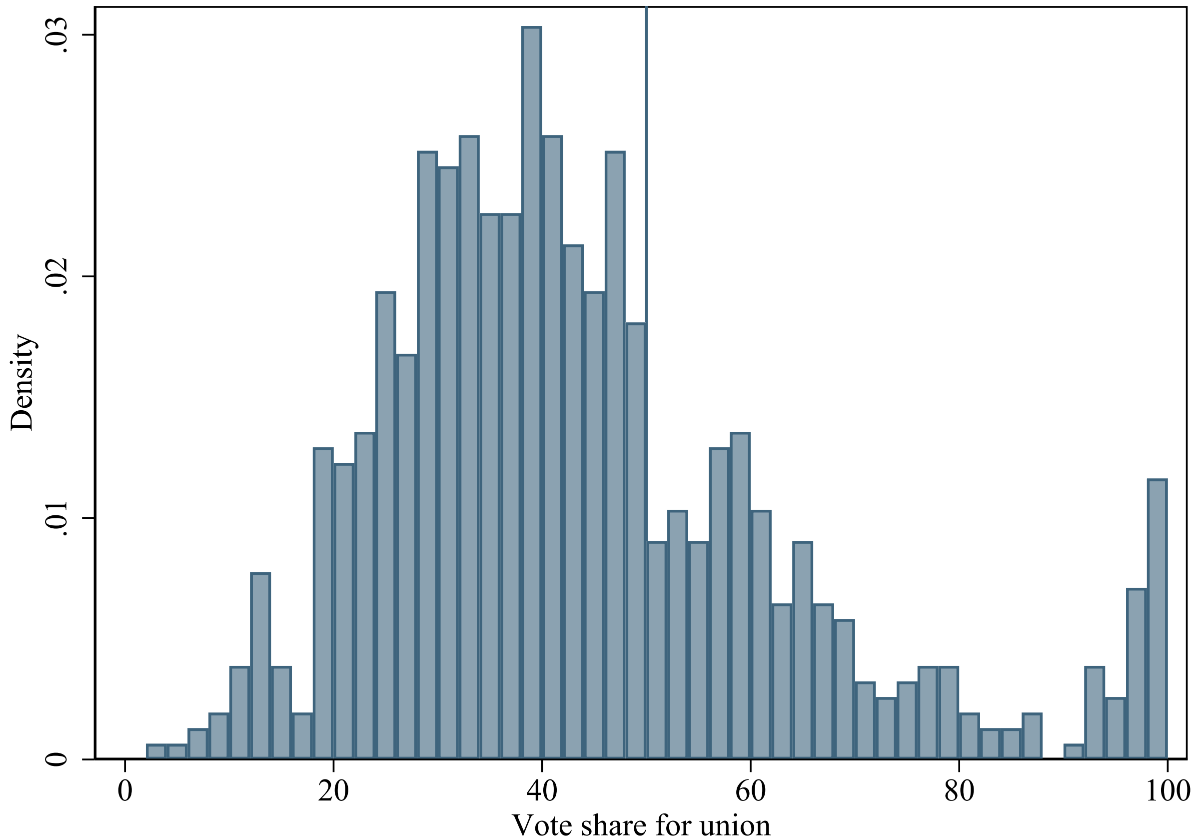
Second, we quantitatively test whether there is a discontinuity in the density of our running variable, vote share. To this end, we present results of the recently proposed test by Cattaneo et al. (2020) in Figure 3 and the McCrary (2008) density test in Appendix Figure 2. While we validate the absence of manipulation under the former test, the discontinuity estimate of the latter test marginally rejects the null hypothesis of no discontinuity. Several additional considerations dispel concerns that the discontinuity could imply problematic

Table 1: Pre-Treatment Balance

	No Union	**Unionization**	**Difference**		
	Mean	Mean	Diff	T-stat	p
Total waste release ratio	0.912	0.958	-0.046	-0.819	0.414
On-site release ratio	0.887	0.962	-0.075	-1.434	0.153
Off-site release ratio	0.884	0.788	0.096	0.846	0.400
Total waste cure ratio	0.934	0.923	0.011	0.162	0.871
On-site cure ratio	0.998	0.946	0.052	0.507	0.613
Off-site cure ratio	0.865	0.734	0.130	1.599	0.112
Catastrophic releases	0.018	0.019	-0.001	-0.037	0.970
Prevention count	0.201	0.281	-0.081	-1.304	0.193
Innovative prevention	0.022	0.019	0.003	0.222	0.825
Production ratio	1.031	0.927	0.104	2.810	0.005
Ln (Sales)	17.045	17.026	0.020	0.110	0.913
Ln (Employees)	5.014	5.155	-0.140	-0.878	0.380
Paydex score	67.047	68.274	-1.227	-1.030	0.304

Notes: This table presents the pre-election variable balance between the treatment and control group in the MSE-optimal bandwidth of our main specification in the year before unionization. All variables are defined in Appendix Table 1.

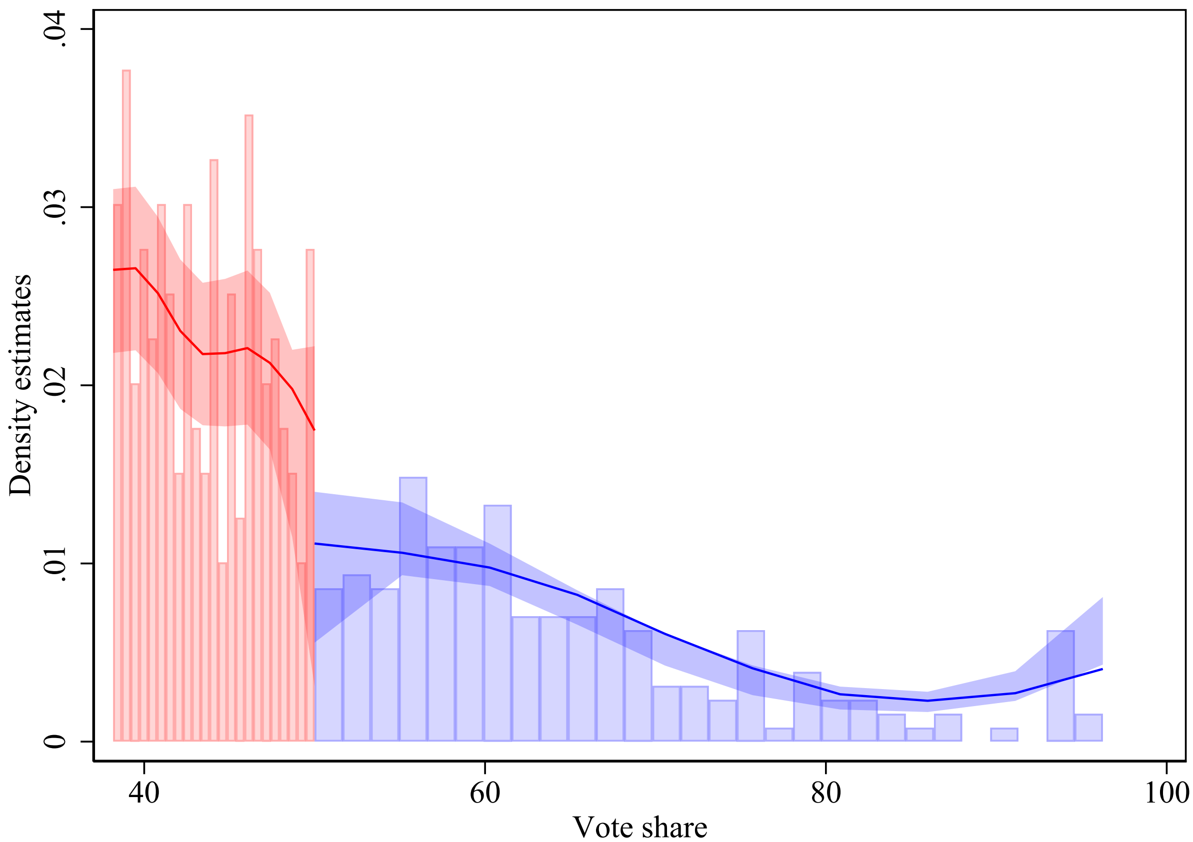
Figure 2: Distribution of Vote Shares



Notes: This figure plots a histogram of the distribution of the adjusted vote shares across 50 equally spaced bins. Each observation represents a unique election in our main sample. Union election results are from the NLRB over 1990–2017.

manipulation in our setting. The McCrary test is not designed for discrete variables like vote share. Therefore, we perform a formal discontinuity test explicitly designed for discrete variables using vote share rounded to the nearest integer (Frandsen, 2017). With a p-value of 0.23, this test again does not indicate the presence of manipulation. Moreover, we argue that with an increasing number of voters, less precise majority manipulation in the secret ballots is possible (DiNardo and Lee, 2004; McCrary, 2008). This argument additionally supports the restriction of our sample to elections with at least 50 voters. In sum, we conclude that both identifying assumptions of the RDD are sufficiently satisfied.

Figure 3: Cattaneo (2020) Discontinuity Test



Notes: This figure plots the Cattaneo et al. (2020) discontinuity test for the density of the adjusted vote share variable. The bins plot the distribution of the adjusted vote shares in a histogram. The solid red and blue line estimate the local polynomial density and the shaded red and blue compute bias corrected confidence intervals on each side of the cutoff. Each observation represents a unique election in our main sample.

4 Main Results

This section presents our main empirical results from applying the RDD in Equation 2. We discuss our baseline estimates, external validity, and a battery of robustness checks.

4.1 Effects on the Sustainability-Safety Tradeoff

In our main analysis, we estimate the effect of unionization on changes in toxic chemical releases and cure. We consider average effects up to three years after the election as well as year-by-year effects to investigate both short-term and medium-term effects.

Table 2: Main Analysis

	Changes in waste releases			Changes in waste cure		
	Total	On-site	Off-site	Total	On-site	Off-site
	(1)	(2)	(3)	(4)	(5)	(6)
Unionization	0.096	0.147**	−0.172	−0.236	−0.593**	−0.085
	(0.085)	(0.071)	(0.171)	(0.179)	(0.244)	(0.125)
Mean	0.941	0.942	0.880	0.943	0.948	0.909

Notes: This table presents the unionization effect on changes in waste releases and cure up to three years post election. We report bias-corrected local regression discontinuity estimates using the MSE-optimal bandwidth and triangular kernel. The mean reports the control group's average dependent variable in the MSE-optimal bandwidth below the cutoff of 50%. Union election results are from the NLRB over 1990–2017. Toxic waste data are from EPA's TRI over the 1991-2020 time period. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

Table 2 reports the corresponding regression discontinuity estimates using local polynomial estimation of order one. We find a positive effect of unionization on changes in releases, i.e. increasing releases. This effect is not statistically significant for the total release ratio in column (1). In columns (2) and (3), we decompose waste releases into on-site and off-site releases. On-site, where unions arguably have greater incentives to influence managerial decisions, the change in waste releases increases significantly by 0.147 for union wins compared to union losses. At the 50% cutoff, this estimate reads as follows: facilities with a union loss reduce their on-site releases by approximately 12% whereas facilities with union wins increase on-site releases by 3%, compared to the previous year. Off-site changes in releases are statistically indistinguishable from zero.

Since we conjecture that unions are particularly concerned with workplace safety and might therefore push for less waste cure, increasing releases can be a consequence of this preference. We explicitly test this hypothesis in columns (4) to (6) of Table 2. In column (4), we report a negative, albeit not statistically significant effect of unionization on the change in waste cured. Successful unionization implies an, on average, 0.24 drop in the waste cure ratio compared to facilities where a union lost in a close-call election. Again, we further validate this finding by decomposing total waste cure into on-site and off-site cure. As expected, unionization significantly and strongly impacts on-site waste cure. On a year-by-year average, unionized facilities reduce on-site waste cure by 59 percentage points compared to non-unionized facilities. This effect is statistically significant marginally above the 1%-level (column (5)). The economic magnitude of the effect is quite large but should be interpreted with caution considering that it is the local treatment effect at the cutoff. In contrast, column (6) shows that there is hardly any detectable effect on off-site waste cure where the elected union has less incentive to take action. This finding is in line with our theory because influences and safety concerns of a union should predominantly affect workers' exposure to waste cure at their facility, i.e. on-site.

To further understand the timing of the apparent negative waste cure effect coinciding with higher on-site releases over three years after a successful election, we rerun our analysis for each year separately. Table 3 contains the results. As can be seen in panel A, changes in total releases increase strongest in the first year but are never statistically significant. In panel B, we show that the on-site release ratio is positive in all years and statistically significant in the year directly following unionization. The results for waste cured in panel C and on-site waste cured in panel D mirror this strong immediate reaction and somewhat lower lagged effects. Specifically, we observe a large reduction in the on-site waste cure ratio in the year following the election, significant at the 1%-level. Again, the economic magnitude of the effect is quite large.

Overall, our results suggest that newly unionized firms have a strong distaste for waste cure, especially on-site, at the expense of higher on-site releases to the environment. This effect generally is strongest in the year following unionization, but unions' rejections of on-site waste cure in particular persist in the second and third year as well. Hence, unions seem to prioritize workplace safety or cost-savings over potentially detrimental effects through

Table 3: Waste Release and Cure by Year

	(1) t+1	(2) t+2	(3) t+3
Panel A: Changes in waste releases			
Unionization	0.134	−0.022	0.096
	(0.135)	(0.169)	(0.161)
Panel B: Changes in on-site waste releases			
Unionization	0.231*	0.103	0.172
	(0.134)	(0.115)	(0.162)
Panel C: Changes in waste cure			
Unionization	−0.353*	−0.102	−0.134
	(0.205)	(0.159)	(0.349)
Panel D: Changes in on-site waste cure			
Unionization	−0.981***	−0.886**	−0.493*
	(0.302)	(0.390)	(0.293)

Notes: This table presents the unionization effect on changes in waste releases and cure for each of the three years post election. We report bias-corrected local regression discontinuity estimates using the MSE-optimal bandwidth and triangular kernel. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

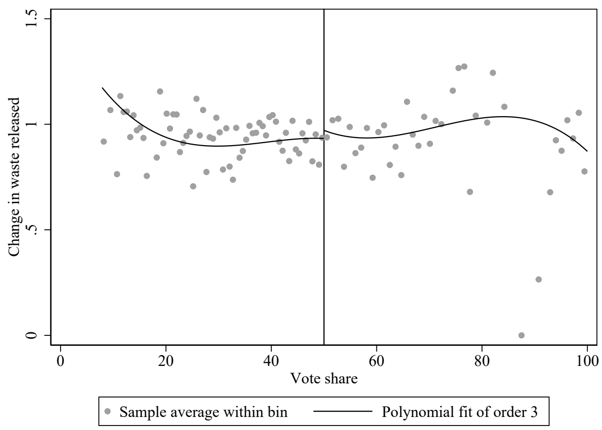
increased exposure to chemicals in the environment. In other words, safety dominates sustainability. Moreover, the lower magnitude of the release effect might indicate that facilities find other ways than releases to facilitate less cure.

4.2 External Validity

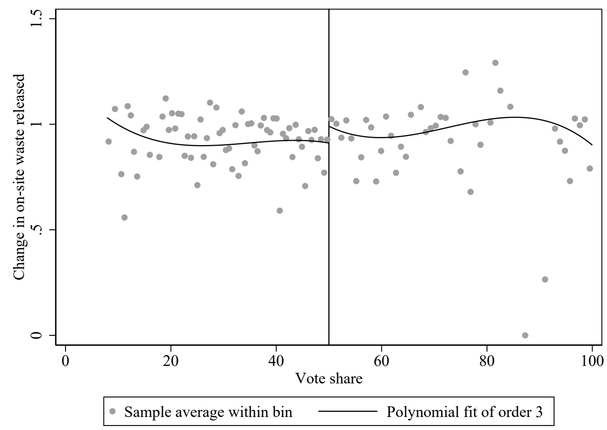
An inherent feature of a local RDD is strong internal validity and limited external validity because only close-call elections are considered. Hence, the local RDD excludes many observations when estimating the treatment effect but for these observations with clearer election outcomes, the relationship between waste management practices and unionization might differ. To probe into the external validity of our results, we estimate a global polynomial RDD using all observations, i.e. including clear wins and losses. Since this global RDD introduces more bias in coefficients, literature suggests using a higher-order polynomial (Dittmar et al., 2020).

Figure 4: Global Polynomial Discontinuity Estimates

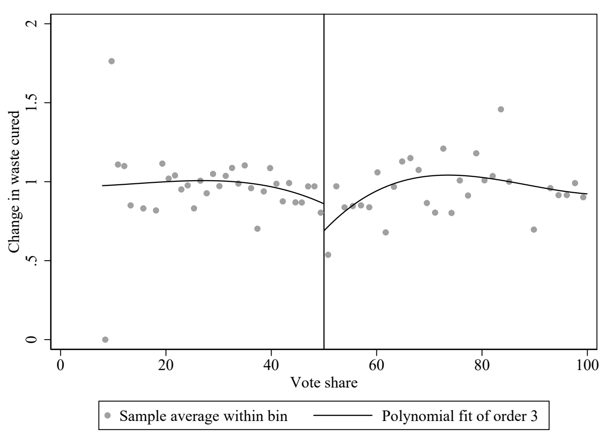
A Changes in total waste releases



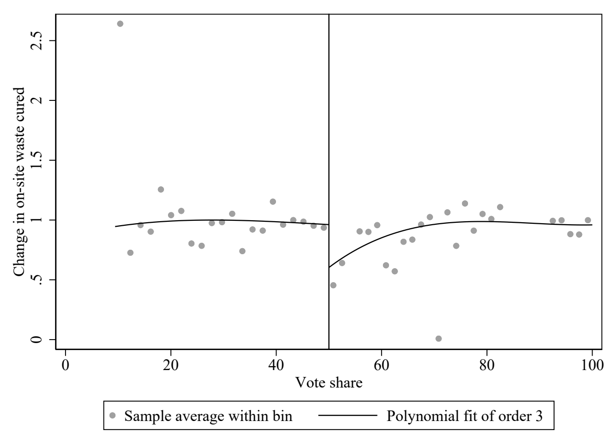
B Changes in on-site waste releases



C Changes in total waste cure



D Changes in on-site waste cure



Notes: This figure presents the global unionization effect on waste management practices up to three years post election. Figure A presents changes in total waste releases, Figure B presents changes in on-site waste releases, Figure C presents changes in total waste cure, and Figure D presents changes in on-site waste cure. We show third-order global regression discontinuity functions.

Figure 4 plots the global third-order polynomial (Dittmar et al., 2020) from a regression of waste release and cure ratios on vote share at both sides of the unionization cutoff. The effect of unionization on changes in waste releases appears relatively small in this global estimation. On the contrary, there is a drop in the waste cure ratio at the right side of the cutoff, i.e. when the union election passes with a union win. For both releases and cure, the effects are stronger on-site than in total.

Table 4 reports the conventional estimates of unionization in a global third-order polynomial on our main dependent variables, total waste release ratio (columns (1)-(2)), on-site waste release ratio (columns (3)-(4)), total waste cure ratio (columns (5)-(6)), and on-site waste cure ratio (columns (7)-(8)). In columns (2),(4), (6), and (8), we include year and chemical fixed effects to control for unobservable variations across chemicals and over time.

Table 4: Global Regression Discontinuity (Third-Order Polynomial)

	Changes in waste releases				Changes in waste cure			
	Total		On-site		Total		On-site	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Unionization	0.098	0.074	0.118*	0.098*	-0.193	-0.265**	-0.399**	-0.479**
	(0.071)	(0.069)	(0.063)	(0.059)	(0.136)	(0.122)	(0.201)	(0.198)
Chemical FE	No	Yes	No	Yes	No	Yes	No	Yes
Year FE	No	Yes	No	Yes	No	Yes	No	Yes
N	4044	4015	3789	3752	3335	3292	1712	1677

Notes: This table presents the unionization effect on changes in waste releases and cure up to three years post election. In columns (2),(4), (6), and (8), we include year and chemical fixed effects to control for unobservable variations across chemicals and over time. For estimations with fixed effects, we drop singleton chemical or year observations (Correia, 2017). We report conventional regression discontinuity estimates using the global bandwidth, triangular kernel, and a third-order polynomial. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

As can be seen, the direction and size of our estimates are consistent when using a global polynomial setting.[8] Consequently, we show that our results can be extrapolated to the complete sample of union elections despite our main results from the local RDD having primarily internal validity. The effect sizes of changes in on-site waste releases and cure remain meaningful, but slightly smaller for all elections compared to close-call elections in our main analysis.

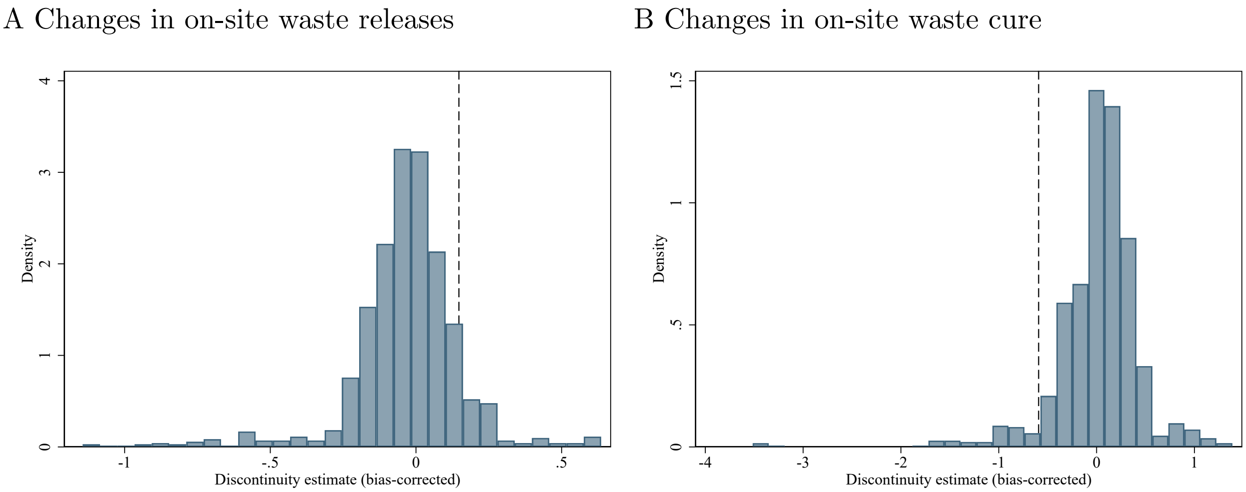
[8]We verify that our results are not driven by particular estimation specifications. Specifically, our results remain virtually unchanged when estimating a second-order or fourth-order polynomial and when using a kernel function (epanechnikov or uniform) that more equally weights observations across the range of our running variable.

4.3 Robustness

This section examines the robustness of our local RDD findings to assumptions and default econometric model specifications. To this end, we perform several contextual and local RDD-specific robustness checks. First, we change the threshold for including ratios in our analysis from three to two and four (Akey and Appel, 2021). Second, we use the natural logarithm of the ratios, thereby mitigating the effects of outliers while excluding zeros, i.e. observations where facilities do not release or cure a chemical anymore (Dutt and King, 2014). Third, we rerun our analysis using a multiplicative version of the individual ratios in year t+3. We construct this measure by multiplying the ratios in the three years following the election which represents an alternative for testing the medium-term, compound effect of unionization on chemical-level waste management practices (Akey and Appel, 2021). Fourth, we add year and chemical fixed effects as covariates. Fifth, we adjust the number of eligible voters to 25 and 75 (DiNardo and Lee, 2004). Lastly, we only include chemicals for which the EPA consistently mandated reporting during our sample period. Our main result of increasing on-site waste releases and decreasing on-site cure is robust in all specifications and significant in the vast majority of alternative specifications. Appendix Table 3 contains corresponding estimation results.

While these robustness tests target contextual assumptions, variables, or sample definitions, a local RDD requires discretionary specifications for estimation. Although we use standard specifications, we still test the sensitivity of our estimates towards alternative specifications. These tests include using a second-order local polynomial to estimate the discontinuity and using an epanechnikov kernel function. Moreover, we also run a "donut" RDD where we exclude observations directly at the cutoff, i.e. those with a vote share or adjusted vote share of 50% (Barreca et al., 2011). As shown in Appendix Table 4, our results remain virtually unchanged across these robustness checks.

Figure 5: Density of Discontinuity Estimates at Placebo Cutoffs



Notes: This figure plots a histogram of the distribution of the discontinuity estimates from placebo tests with artificial cutoffs between 20% and 80% of the vote share. The dashed vertical line represents the “true” discontinuity estimates from Table 2.

Finally, we perform placebo tests using alternative cutoffs. Thus, by artificially assigning treatment observations to the control group or vice versa, we expect to not find an effect in most cases, implying that our estimated effects are not random. Specifically, we alter the cutoff from vote shares of 20% to 80% in steps of 0.05%, hence running 1,200 iterations of our RDD with 1,200 different cutoffs (Bordignon et al., 2016). Figure 5 compares the discontinuity estimates from these false cutoffs to the true estimate of .147 for the on-site release and -.593 for the on-site cure changes, depicted by the dashed lines. The large majority of these RDD runs produces an estimate of zero or close to zero, i.e. does not find a significant effect of unionization on on-site waste releases and cure. This finding further supports our identification and attribution of our results to the actual treatment, unionization.

5 Cross-Sectional Heterogeneity

Our main analysis shows that unionization results in increasing on-site releases and decreasing waste cure, especially on-site. Next, we investigate whether (i) union power as determined by legislation, (ii) chemical toxicity, and (iii) industry affiliation moderate our effects.

5.1 Union Power

As of 2022, 27 states in the US have right-to-work (RTW) laws in place. Under RTW legislation, workers at a facility are not obliged to join or pay for the union that represents them. Recent empirical evidence underscores that RTW laws essentially restrict the power of unions with real effects on, i.a., investment and innovation (Bradley et al., 2017; Chava et al., 2020). Consequently, if unions actively bargain for changes in waste management practices, we expect these changes to be stronger for facilities located in a non-RTW law state where unions have more bargaining power. For estimation, we run Equation 2 for two samples, namely for observations from RTW versus non-RTW states.[9]

Table 5 contains the corresponding estimates. The results in panel A show that unionization does not have a significant effect in RTW law states, neither on changes in (on-site) releases nor in (on-site) cure. On the contrary, our main effects are significant in non-RTW states (panel B). Especially, we find that the decrease in (on-site) waste cure is stronger than in RTW states. Some scholars argue that the general attitude toward unions rather than effective RTW legislation determines union power (e.g. Farber, 1984). To proxy for this general attitude, we split the sample into observations from states *eventually* versus *not* adopting RTW legislation. Appendix Table 5 shows that our results of significant effects in non-RTW states hold. In sum, these findings suggest that unions require sufficient

[9] We assign states according to the year they pass RTW legislation. For instance, chemical-level waste from facilities in Michigan is in the non-RTW split of the sample before 2012 and in the RTW split afterwards.

Table 5: Union Power and Waste Handling

	Changes in waste releases		Changes in waste cure	
	Total (1)	On-site (2)	Total (3)	On-site (4)
Panel A: RTW state				
Unionization	0.131	0.160	−0.009	−0.058
	(0.158)	(0.215)	(0.266)	(0.369)
Panel B: Non-RTW state				
Unionization	0.072	0.134*	−0.422***	−1.005***
	(0.107)	(0.072)	(0.116)	(0.191)

Notes: This table presents the unionization effect on changes in waste releases and cure up to three years post election. We split our sample in observations from US states with Right-to-Work laws in panel A and US states without in panel B. We report bias-corrected local regression discontinuity estimates using the MSE-optimal bandwidth and triangular kernel function. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

bargaining power to cause real changes in waste management practices.

5.2 Chemical Toxicity

So far, our analysis deals with quantity-based measures of toxic waste that implicitly assume a general comparability across chemicals (Kassinis and Vafeas, 2006; Dutt and King, 2014; Wang et al., 2021). However, our chemical-level data also allows for explicit accounting of differing toxicity as done by other scholars (King and Lenox, 2002; Berrone and Gomez-Mejia, 2009; Kim et al., 2019). As such, we can investigate whether chemical toxicity moderates the sustainability-safety tradeoff.

On one hand, union's observed distaste for waste cure could intensify for extremely toxic materials due to higher safety concerns. Given the proximity of blue-collar workers to production (Bradley et al., 2017) and detailed data provided by EPA and other institutions it is also reasonable to assume that unions could gather this toxicity information. On the other hand, more toxic chemicals are arguably more regulated and have greater negative effects when released. Consequently, waste management practices for these chemicals might be less flexible in general.

For estimation, we use reportable quantities for emergency action to split the sample. All chemical-facility-year observations with a reportable quantity of 1, 10, and 100 pounds constitute the more toxic subsample. Higher, i.e. 1000 and 5000 pounds, or missing reportable quantities hence constitute the less toxic subsample.

Table 6: Chemical Toxicity

	Changes in waste releases		Changes in waste cure	
	Total (1)	On-site (2)	Total (3)	On-site (4)
Panel A: RQ $\leq$ 100 pounds				
Unionization	0.084	0.131	-0.159	-0.448^{***}
	(0.144)	(0.118)	(0.184)	(0.155)
Panel B: RQ $>$ 100 pounds				
Unionization	0.132	0.166^{**}	-0.256	-0.586^{**}
	(0.086)	(0.080)	(0.190)	(0.272)

Notes: This table presents the unionization effect on changes in waste releases and cure up to three years post election. We split our sample in chemicals that have low reportable quantities (RQ below or equal to 100 pounds), i.e. higher toxicity, and those with high reportable quantities (RQ above 100 pounds or not specified), i.e. lower toxicity. Panel A and B report the respective subsamples which are approximately evenly distributed over the total number of observations. We report bias-corrected local regression discontinuity estimates using the MSE-optimal bandwidth and triangular kernel function. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

As shown in Table 6, accounting for diverging toxicity does not affect the direction of our main result. Panel A and panel B both report increasing release ratios and decreasing cure ratios. We show that changes in on-site releases are not statistically significant despite their similar magnitude to our main effect, whereas changes in on-site waste cure drops significantly also for more toxic chemicals (panel A). On the contrary, the coefficients for on-site release and cure are statistically significant and larger in absolute terms for the subsample of less toxic chemicals (panel B). In Appendix Table 6, we report further heterogeneity analyses for toxicity subsamples using the human toxicity potential to split the sample (Hertwich et al., 2001). Again, real effects tend to be slightly smaller for more toxic chemicals.

Tighter environmental permits for releasing these chemicals and more detrimental effects of pollution exposure are possible underlying explanations. Nevertheless, this heterogeneity analysis suggests that increasing on-site releases and especially decreasing on-site cure are present in both toxicity subsamples.

5.3 Industry Affiliation

Subsequently, we explore whether a facility being affiliated to a heavy versus a non-heavy industry matters for waste releases and waste cure. Workers expectations on job risks, such as exposure to dangerous procedures or chemical emissions might depend on the industry affiliation of their employers. Viscusi (1979) finds that workers have no perfect information on job risk and their perception is positively correlated with industry risk.

Specifically, we conjecture that workers in "dirty", i.e. heavy, industries likely expect a certain extent of exposure to hazardous substances and also regard releases and waste management practices as an integral part of performing business. Additionally, industries likely face different possibilities and (regulatory) constraints in adjusting waste management as a response to unions' pressures. To investigate the role of industry affiliation and to approximate corresponding potential preferences of workers, we split our sample into chemical observations from heavy industry (e.g. chemicals, oil and gas) versus non-heavy industry facilities (e.g. food processing, electronics).[10]

Panel A of Table 7 shows significant positive effects of unionization in non-heavy industries on changes in total (column (1)) and on-site releases (column (2)) as well as significant negative effects on changes in total (column (3)) and on-site waste cure (column (4)). Panel B mostly corroborates the direction of these effects also for observations from heavy industries, although only the on-site cure coefficient is statistically different from zero.

[10]Specifically, we assign facilities with a primary NAICS starting with 21, 221–223, 324–327, 331–332, and 562 to the heavy-industry subsample and all others to the non-heavy industry subsample.

Table 7: Industry Affiliation

	Changes in waste releases		Changes in waste cure	
	Total	On-site	Total	On-site
	(1)	(2)	(3)	(4)
Panel A: Non-heavy industry				
Unionization	0.198**	0.237***	−0.408***	−0.832***
	(0.098)	(0.058)	(0.122)	(0.174)
Panel B: Heavy industry				
Unionization	0.125	0.124	0.018	−0.477*
	(0.123)	(0.130)	(0.297)	(0.267)

Notes: This table presents the unionization effect on changes in waste releases and cure up to three years post election. We split our sample in observations from non-heavy industries in panel A and heavy industries in panel B. We report bias-corrected local regression discontinuity estimates using the MSE-optimal bandwidth and triangular kernel function. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

Overall, our results of decreasing waste ratios and simultaneously increasing release ratios are stronger for facilities in non-heavy industries. This finding suggests that workers from heavy industries have different risk expectations and interests regarding toxic waste management. Moreover, regulations and higher marginal costs of changing waste management practices might limit the wiggle room for unions to influence the sustainability-safety tradeoff in heavy industries.

6 Further Analyses

In our last results section, we aim to examine alternative explications and further disentangle the drivers of our main effects, increasing on-site release and decreasing on-site cure. To this end, we first investigate whether changes in production output mechanically affect our results. Second, we explore whether our main effects might be driven by motivations to save costs. Finally, we test our hypothesis of more pollution prevention activities due to their multi-win character.

6.1 Production Output

Production output is highly positively correlated to waste releases and waste cure. Hence, quantitative changes in these waste management procedures could be purely mechanical and result from changing production output after unionization, e.g. because of lower productivity and shirking due to misaligned incentives (Bradley et al., 2017). To examine the role of production output, we use the production ratio as presented in Section 3.2, i.e. output related to chemical use in one year divided by output in the previous year.

Table 8: Production Output

	Production	Changes in waste releases		Changes in waste cure	
	ratio	Total	On-site	Total	On-site
	(1)	(2)	(3)	(4)	(5)
Unionization	0.100	0.108	0.131^{*}	-0.281^{*}	-0.599^{**}
	(0.098)	(0.093)	(0.077)	(0.166)	(0.234)
Production ratio		0.265^{***}	0.272^{***}	0.206^{**}	0.274^{*}
		(0.059)	(0.069)	(0.084)	(0.144)

Notes: This table presents the unionization effect on the production ratio in column (1). We report the unionization effect on changes in waste releases and cure up to three years post election when controlling for the production ratio in columns (2)-(5). We report bias-corrected local regression discontinuity estimates using the MSE-optimal bandwidth of our main analysis and triangular kernel function. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

In Table 8, we first test whether unionization has a significant effect on this production ratio. Column (1) shows that unionization does not significantly affect the production ratio up to three years after the election takes place. This finding is consistent with DiNardo and Lee (2004) and does not point to the presence of shirking on job duties by unionized workers.

Next, we report unionization and production effects on waste management strategies in columns (2) to (5). We find no relevant changes in coefficients and significance levels when including the production ratio as a control variable, compared to our main specification in Table 2. Hence, we rule out that the increase in on-site waste releases and strong decrease in cured waste after unionization is a purely mechanic consequence of changing production levels.[11] Furthermore, these findings address concerns with respect to the observed pre-election unbalance of the production ratio between treatment and control group.

6.2 Financial Constraints

Next, we consider financial constraints, approximated by the minimum Paydex credit score, as a possible mechanism of our results. The unionization effect on financing decisions, such as debt-to-equity ratios, is well documented in the literature (Bronars and Deere, 1991; Klasa et al., 2009; Matsa, 2010). Moreover, other studies present evidence that more financially constrained firms release more toxic chemicals (Xu and Kim, 2022), potentially because the costs associated with end-of-process waste cure procedures require sufficient funding (Frondel et al., 2007; Dutt and King, 2014). Our sample reflects these findings as the Paydex score and changes in waste cure are positively correlated.

Table 9: Financial Constraints

	Paydex score	Changes in waste releases		Changes in waste cure	
		Total	On-site	Total	On-site
	(1)	(2)	(3)	(4)	(5)
Unionization	6.099	0.183^{**}	0.163^{**}	-0.265	-0.700^{***}
	(3.817)	(0.071)	(0.074)	(0.226)	(0.260)
Paydex score		-0.000	-0.001	0.005	0.009^{**}
		(0.002)	(0.002)	(0.003)	(0.004)

Notes: This table presents the unionization effect on the minimum Paydex credit score in column (1). We report the unionization effect on changes in waste releases and cure up to three years post election when controlling for the Paydex score in columns (2)-(5). We report bias-corrected local regression discontinuity estimates using the MSE-optimal bandwidth of our main analysis and triangular kernel function. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

We examine whether unionization directly impacts financial constraints and thereby in-

[11] In untabulated robustness checks, we use the log of sales and the log of employees from NETS instead of TRI's production ratio to proxy for production output. Our results remain virtually unchanged.

directly impacts waste cure and releases. As shown in column (1) of Table 9, we find a positive but insignificant local unionization effect on the Paydex score. We show that financial constraints do not significantly mediate the relationship between unionization and changes in waste releases and total waste cure (columns (2) to (4)). In column (5), we show that, ceteris paribus, an increase in the Paydex score increases on-site cure. However, the mediating effect is relatively small. These results suggest that financial constraints, and possibly cost savings, are not the main motivation of our waste cure effect. Thus, workplace safety concerns appear to be more relevant for the observed decrease in waste cure.

6.3 Prevention Activities

Our empirical results indicate that unions care about workplace safety, leading to reduced on-site waste cure but increased on-site releases. However, the magnitude and significance of the effect on waste releases is generally weaker. Subsequently, we investigate potential channels mitigating the increase in waste releases.

First, employee training should be a primary concern for labor unions to achieve higher workplace safety. We proxy training by the binary variable "Cata" which is equal to one if a facility reports catastrophic releases and zero otherwise. These releases occur rarely and unplanned due to human or technical failure and might decrease with better training possibilities. As shown in column (1) of Table 10, we find a significant negative local effect of unionization on Cata, i.e. catastrophic releases. Hence, unions might bargain for higher safety and training standards facilitating this decrease in catastrophic releases.

Next, we analyse a potential channel through which facilities and unions can generate a multi-win outcome: pollution prevention through innovative product and process modifications. Specifically, by eco-designing products or by inventing new production processes and technologies, facilities can manufacture the same level of output with a lower chemical waste intensity. Consequently, such innovation also reduces pressure on the sustainability-safety

Table 10: Prevention Activities

	Cata	Inno	Changes in waste releases		Changes in waste cure	
			Total	On-site	Total	On-site
	(1)	(2)	(3)	(4)	(5)	(6)
Unionization	-0.067^{***}	0.045^{**}	0.122	0.166^{**}	-0.222	-0.589^{**}
	(0.026)	(0.019)	(0.086)	(0.072)	(0.181)	(0.244)
Cata			0.195	0.164	0.059	0.001
			(0.163)	(0.174)	(0.203)	(0.141)
Inno			-0.154^{*}	-0.073	-0.170^{*}	-0.175^{**}
			(0.084)	(0.124)	(0.097)	(0.087)

Notes: This table presents the unionization effect on catastrophic releases (Cata) and innovative pollution prevention (Inno) in columns (1) and (2). In columns, 3 to 6, we present unionization effects on changes in waste releases and cure up to three years post election when controlling for the alternative waste prevention mechanisms, catastrophic releases and innovative prevention. We report bias-corrected local regression discontinuity estimates using the MSE-optimal bandwidth of our main analysis and triangular kernel function. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

tradeoff, as there is less necessity for cure for the same level of releases. We operationalize innovative prevention by means of a binary variable ("Inno") equal to one if facilities report a pollution prevention activity related to innovation and zero otherwise. We report the positive unionization effect on Inno, i.e. the adoption of innovative pollution prevention activities, in column (2) of Table 10.[12]

Finally, we examine whether catastrophic releases and innovative prevention activities relate to changes in waste management practices. To this end, we add corresponding indicator variables as controls and repeat the discontinuity estimations for the release and cure effects in columns (3) to (6).[13] As predicted, facilities that report innovative prevention activities are associated with less waste releases (columns (3) and (4)) and less cure (columns (5) and

[12]Our results are robust to using a more narrow definition of innovation, EPA's category of product innovation. We also investigate other source reduction activities such as the adoption of good operating practices, but do not find a significant effect of unionization on these indicators. Since Cata and Inno are both binary variables, we rerun the RDD in columns (1) and (2) of Table 10 using logit and poisson estimation. Again, results remain robust (untabulated).

[13]The similarity of the unionization discontinuity estimates when controlling and when not controlling for our mechanisms ease concerns about measurement error due to unionization-induced changes in mechanisms that confound with changes in waste management practices (Angrist and Pischke, 2009; Heckman et al., 2013).

(6)), ceteris paribus. Nevertheless, our unionization effects stay large even after controlling for both potential channels, catastrophic releases and innovative prevention.

7 Discussion and Conclusion

This paper investigates changes in waste management practices by facilities following a union election. Theoretical reasoning suggests that unions essentially face a tradeoff between protecting their members' safety at the workplace and supporting sustainability by protecting members, neighboring communities, and the environment from pollution. In the context of waste management, the tradeoff exists because curing waste is costly and relatively dangerous for its workers whereas releases pollute the environment and negatively affect the health of workers and community members.

Exploiting a quasi-experimental setting with close-call union elections, we document that facilities decrease on-site waste cure and increase on-site releases after unionization. The discontinuity estimates of our local RDD are economically large and hold across several contextual and econometric robustness checks. Moreover, we test and show the generalizability of our results to all union elections using a global RDD. In cross-sectional analyses, we find that our results are not statistically significant in states with lower union bargaining power as indicated by effective RTW laws. Chemical toxicity and industry affiliation also play a role, with more toxic chemicals and heavy industries attenuating our main effects.

Taken together, these results add to previous evidence of a negative effect of unionization on globally relevant emissions (Ertugrul and Marciukaityte, 2021; Heitz et al., 2021). The locality of toxic waste tightens the tradeoff workers face as they are also exposed to negative pollution externalities on the local chemical-facility level. Although workers derive disutility from both releases and cure, the distaste for waste cure dominates. Moreover, we rule out changes in production output as a mechanical driver of our results. Similarly, we show that unionization does not lead to higher financial constraints. Our results call upon managers and policymakers to reconsider current waste cure practices with a focus on increasing safety and workers' trust.

Furthermore, we find that unions decrease catastrophic releases and support the adoption of innovative prevention activities like modifying the design or composition of a product in order to prevent pollution. These activities relax the sustainability-safety tradeoff. Given their multi-win character, managers, governmental institutions, and policymakers should focus on supporting pollution prevention activities with financial resources and expert knowledge. Efforts to increase pollution prevention induce positive external effects across facilities, e.g. because knowledgeable workers change employers and environmental externalities decline. The public good character of such efforts justifies policy support like EPA's Pollution Prevention program.

Finally, our paper highlights the role and importance of blue-collar workers on facilities' environmental performance. We investigate unions' perspective on the sustainability-safety (i.e., release-cure) tradeoff in manufacturing and demonstrate novel channels, through which unions provide benefits for their members. As such, we shed light on potential environmental consequences from increasing union power relative to other stakeholders. i.e. through new legislation currently discussed by the US Senate, the Protecting-the-Right-to-Organize Act. What is good for (unionized) workers is not necessarily good for other stakeholders, such as neighboring communities and the planet. Our findings suggest that successful union elections after 1990 have not been a major driver behind the observed remarkable decline in US manufacturing emissions (Shapiro and Walker, 2018). Future research could investigate whether loosening procedural standards through Advanced Recycling legislations will accelerate or depress this trend, in light of workers' distaste for waste cure.

Appendix

Appendix Table 1: List of Variables

Variable Name	Description	Source
Adjusted vote share	Number of valid votes for joining the union divided by total votes. For even total votes, $^{0.5}/_{\text{number of votes cast}}$ is subtracted from the vote share (DiNardo and Lee, 2004).	NLRB
Catastrophic releases	Dummy variable equal to 1 if a facility reported releases not associated with routine production processes of a specific chemical and zero otherwise.	TRI file 2a
Innovative prevention	Dummy variable equal to 1 if a facility, for a given chemical, reports prevention activity related to innovative product and process modifications (W-codes W50, W51, W59, W80, W82, W83, W84, W89).	TRI file 2a
Paydex score	Facility-level minimum business credit score based on trade credit performance.	NETS
Production ratio	Output changes in the manufactured product or supporting operation related to the use of chemical i at facility j from last $t-1$ to current year t ($PR_{i,j,t}/PR_{i,j,t-1}$).	TRI file 2a
Waste cure ratio	Changes in waste that is recycled, used for energy recovery, or otherwise treated for chemical i at facility j (total, on-site, and off-site) from last $t-1$ to current year t ($WC_{i,j,t}/WC_{i,j,t-1}$).	TRI file 1a
Waste release ratio	Changes in waste that is released to air, land, and water for chemical i at facility j (total, on-site, and off-site) from last $t-1$ to current year t ($WR_{i,j,t}/WR_{i,j,t-1}$).	TRI file 1a
Unionization	Dummy variable equal to 1 if the adjusted vote share is greater or equal to 50.	NLRB
Voters	Number of eligible voters at facility union elections.	NLRB

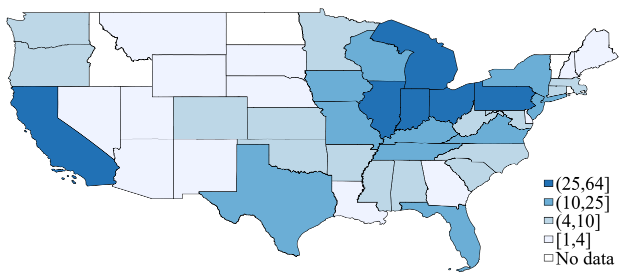
Appendix Table 2: Descriptive Statistics

	Mean	SD	Min	p25	Median	p75	Max
Vote share	44.44	20.34	7.89	30.32	40.71	54.87	100.00
Voters	265.51	481.85	50	80	135.50	290	7000
Total waste release ratio	0.93	0.53	0.00	0.65	0.99	1.11	3.00
On-site release ratio	0.93	0.49	0.00	0.70	1.00	1.07	3.00
Off-site release ratio	0.91	0.66	0.00	0.33	0.96	1.21	3.00
Total waste cure ratio	0.97	0.51	0.00	0.72	0.99	1.19	3.00
On-site cure ratio	0.97	0.48	0.00	0.76	1.00	1.18	2.97
Off-site cure ratio	0.94	0.58	0.00	0.59	0.93	1.19	3.00
Total waste ratio	0.99	0.47	0.00	0.76	0.99	1.17	3.00
Production ratio	1.00	0.37	0.00	0.88	1.00	1.12	3.00
Paydex score	67.95	8.43	21.00	64.00	70.00	74.00	83.00
Innovative prevention	0.02	0.12	0.00	0.00	0.00	0.00	1.00
Catastrophic releases	0.02	0.15	0.00	0.00	0.00	0.00	1.00

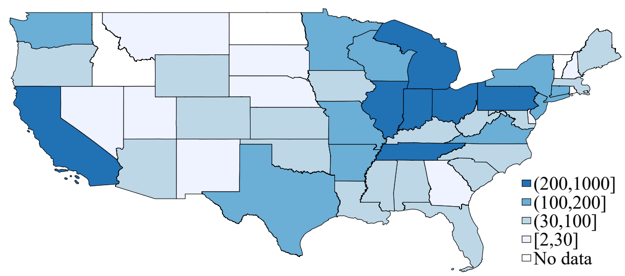
Notes: This table presents summary statistics for our main variables. All variables are defined in Appendix Table 1.

Appendix Figure 1: Cross-State Characterization of Sample

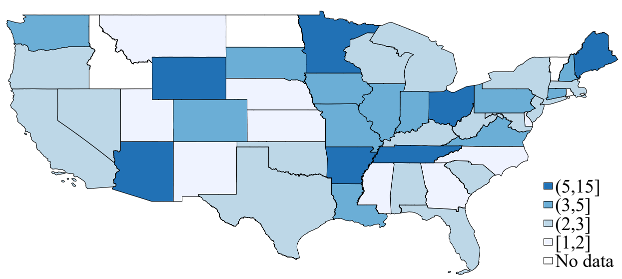
A Elections



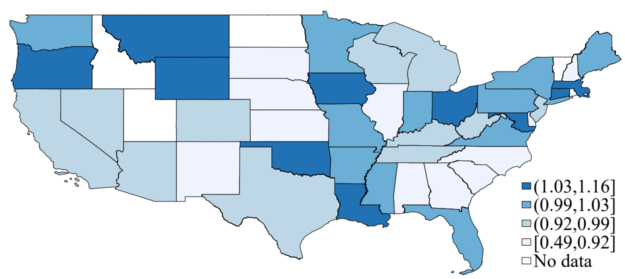
B Observations



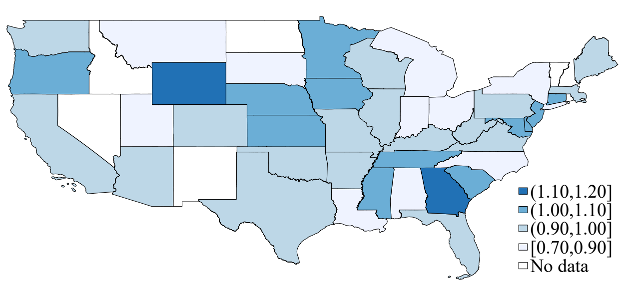
C Average chemicals reported by year



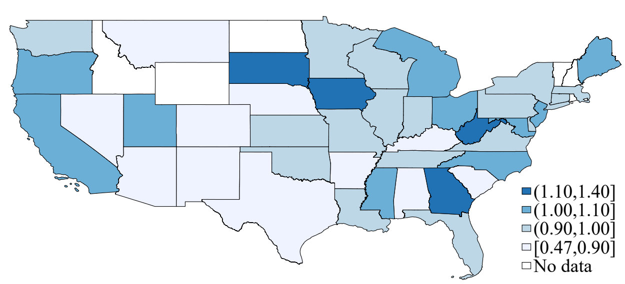
D Changes in production activity



E Changes in waste releases

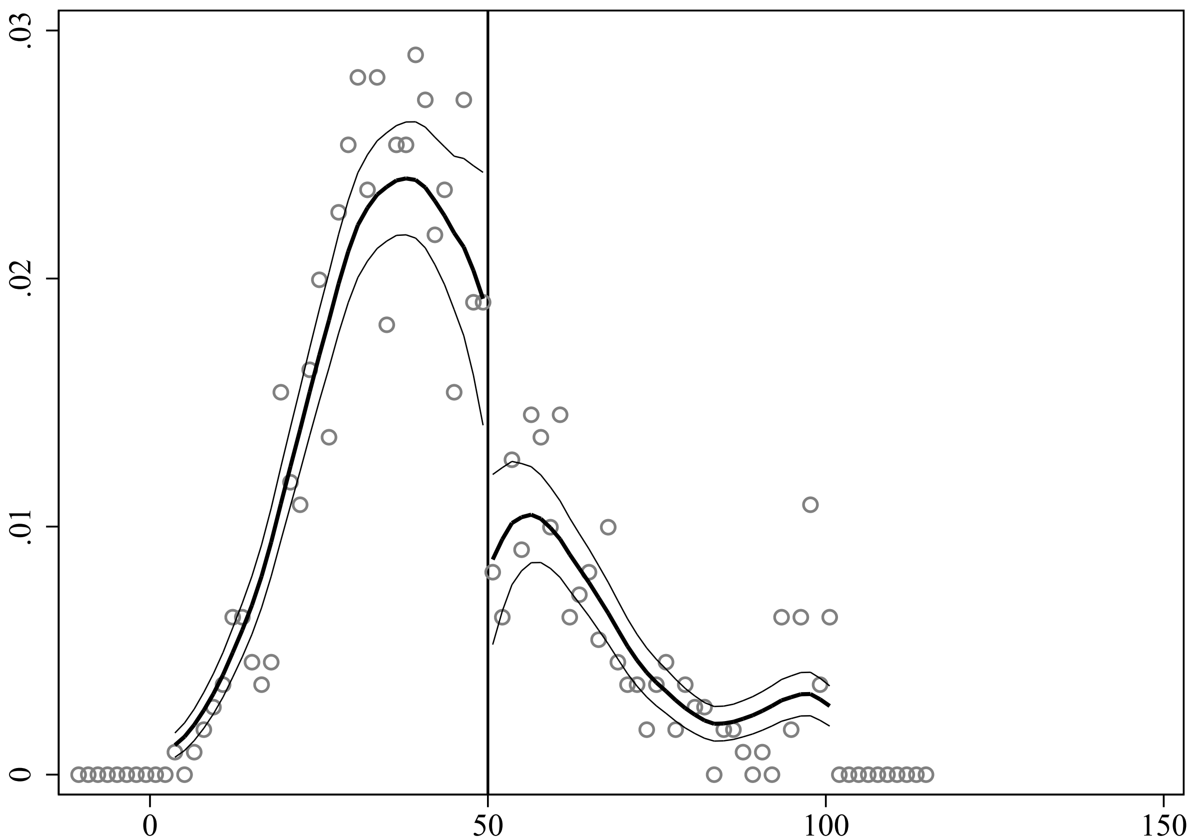


F Changes in waste cure



Notes: This figure presents the sample distribution of our observations and main variables of interest over the United States.

Appendix Figure 2: McCrary (2008) Discontinuity Test



Notes: This figure plots the McCrary (2008) discontinuity test for the density of the adjusted vote share variable. Dots represent bins of vote share data. For estimation of the density function and corresponding confidence intervals, the raw data rather than the bins are used.

Appendix Table 3: Estimation Results for Contextual Robustness Tests

	Changes in waste releases		Changes in waste cure	
	Total	On-site	Total	On-site
	(1)	(2)	(3)	(4)
Panel A: Ratio threshold at 2				
Unionization	0.164*	0.216***	−0.219	−0.497**
	(0.085)	(0.078)	(0.149)	(0.223)
Panel B: Ratio threshold at 4				
Unionization	0.187*	0.226***	−0.300*	−0.625**
	(0.100)	(0.085)	(0.175)	(0.282)
Panel C: Ln(ratios)				
Unionization	0.192	0.259**	−0.172	−1.094**
	(0.275)	(0.132)	(0.203)	(0.525)
Panel D: Cumulative ratios				
Unionization	0.372	0.500**	−0.286	−1.189*
	(0.268)	(0.251)	(0.616)	(0.704)
Panel E: Year + chemical FE				
Unionization	0.052	0.139**	−0.477***	−0.988***
	(0.069)	(0.066)	(0.157)	(0.159)
Panel F: At least 25 voters				
Unionization	0.080	0.112**	−0.107	−0.068
	(0.067)	(0.056)	(0.150)	(0.250)
Panel G: At least 75 voters				
Unionization	0.047	0.116*	−0.407***	−0.999***
	(0.093)	(0.069)	(0.150)	(0.185)
Panel H: Always mandated				
Unionization	0.087	0.128*	−0.184	−0.598**
	(0.084)	(0.066)	(0.147)	(0.249)

Notes: This table presents contextual robustness checks to our main analysis. In panels A and B, we change the threshold value of our waste cure and release ratios to two and four, respectively. In panel C, we estimate our effect with the natural logarithm of these ratios. Panel D shows results for cumulative ratios, estimating compound unionization effects of three years following the election. In panel E, we include year and chemical fixed effects (FE) as covariates. In panels F and G, we restrict our sample to elections with at least 25 and 75 eligible voters, respectively. In panel H, we restrict our sample to chemicals that were mandatory to report in every year of our sample period. We report bias-corrected local regression discontinuity estimates using the MSE-optimal bandwidth and triangular kernel. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

Appendix Table 4: Estimation Results for Econometric Robustness Tests

	Changes in waste releases		Changes in waste cure	
	Total	On-site	Total	On-site
	(1)	(2)	(3)	(4)
Panel A: Second-order polynomial				
Unionization	0.057	0.147^{*}	-0.381^{**}	-0.883^{***}
	(0.112)	(0.089)	(0.185)	(0.322)
Panel B: Epanechnikov kernel				
Unionization	0.097	0.149^{**}	-0.223	-0.596^{**}
	(0.084)	(0.072)	(0.185)	(0.253)
Panel C: Donut RDD				
Unionization	0.143	0.229^{***}	-0.201	-0.588^{**}
	(0.106)	(0.087)	(0.175)	(0.251)

Notes: This table presents econometric, local RDD-specific robustness checks of our main results. Specifically, we estimate the discontinuity around the cutoff using a local second-order polynomial in panel A. In panel B, we use an epanechnikov kernel function that, compared to the default triangular kernel, gives relatively more weight to observations further from the cutoff. Finally, we exclude observations directly at the cutoff, i.e. those with an (adjusted) vote share of 50 percent, to estimate a "Donut" RDD in panel C (Barreca et al., 2011). Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

Appendix Table 5: Union Power – Robustness

	Changes in waste releases		Changes in waste cure	
	Total	On-site	Total	On-site
	(1)	(2)	(3)	(4)
Panel A: Eventually RTW				
Unionization	0.042	0.091	0.037	-0.421
	(0.107)	(0.115)	(0.180)	(0.303)
Panel B: Non-RTW				
Unionization	0.206	0.358^{***}	-0.380^{***}	-0.978^{***}
	(0.150)	(0.115)	(0.108)	(0.166)

This table presents the unionization effect on changes in waste releases and cure up to three years post election. We split our sample in observations from US states that eventually become RTW states in panel A and those that do not become RTW states in panel B. We report bias-corrected local regression discontinuity estimates using the MSE-optimal bandwidth and triangular kernel function. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

Appendix Table 6: Chemical Toxicity – Human Toxicity Potential

	Changes in waste releases		Changes in waste cure	
	Total	On-site	Total	On-site
	(1)	(2)	(3)	(4)
Panel A: High Air-HTP				
Unionization	-0.045	0.001	-0.323^{*}	-0.700^{**}
	(0.104)	(0.082)	(0.188)	(0.333)
Panel B: Low Air-HTP				
Unionization	0.372^{***}	0.350^{***}	-0.178	-0.724^{**}
	(0.112)	(0.124)	(0.286)	(0.352)
Panel C: High Water-HTP				
Unionization	0.092	0.151	-0.124	-0.631^{*}
	(0.114)	(0.093)	(0.196)	(0.351)
Panel D: Low Water-HTP				
Unionization	0.201^{**}	0.182^{*}	-0.695^{***}	-1.134^{***}
	(0.102)	(0.107)	(0.184)	(0.187)

Notes: This table presents the unionization effect on changes in waste releases and cure up to three years post election. We split our sample in chemicals that are above the human toxicity potential median of all chemicals in our sample and those below. We distinguish between air- and water human toxicity potential. We report bias-corrected local regression discontinuity estimates using the MSE-optimal bandwidth and triangular kernel function. Robust standard errors clustered at the facility are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$ indicate significance.

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