Drinking Water Violations and Environmental Justice in the United States, 2011–2015

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Objectives. To assess the extent to which drinking water violations in the United States differed on the basis of county race/ethnicity and socioeconomic status using the primary county served by the community water system (CWS) as the unit of analysis and to determine whether counties with higher proportions of underrepresented groups were disproportionately burdened with repeat violations.

Methods. We used multivariable logistic regression to calculate odds ratios of contextual environmental justice covariates associated with initial and repeat drinking water violations. We obtained violations from the federal Safe Drinking Water Information System. Results were nonstratified and stratified on the basis of population size served by the CWS.

Results. Stratified multivariable logistic regression results revealed previously unobservable patterns in nonstratified findings. Minorities face significant challenges, including exposure to poor water quality. The most notable differences in both initial and repeat violations that we observed were among CWSs that serve large populations. Our most consistent finding was the positive association of initial and repeat violations with the proportion of those who were uninsured, irrespective of stratification.

Conclusions. Greater efforts are needed to ensure that counties with higher proportions of minorities, uninsured households, and low-income households have access to safe drinking water, irrespective of the size of population served by the CWS. (*Am J Public Health.* 2018; 108:1401–1407. doi:10.2105/AJPH.2018.304621)

See also Galea and Vaughan, p. 1288.

Although the United States has among the safest drinking water supplies in the world, harmful contaminants that are detrimental to public health may still be found in public water supplies, as most recently demonstrated by the Flint water crisis.[1] Through the Safe Drinking Water Act of 1974 (SDWA) and its subsequent amendments of 1986 and 1996, the Environmental Protection Agency (EPA) was charged with establishing national health standards to protect drinking water and was granted federal power to regulate public water systems to ensure adherence with set standards.[2] A public water system provides drinking water primarily through piped infrastructure to at least 15 connections or serves 25 or more people for a minimum of 60 days annually; it may be public or privately owned.[3] We examined community water systems (CWS), a type of public water system that supplies drinking water to the same population year round. Most residential settings, including homes, apartments, and mobile park homes, are served by a CWS.[3] Despite the SDWA, many drinking water suppliers receive initial (i.e., first-time) violations as well as repeat violations.[4–6]

Several area-specific studies have found that minorities and lower socioeconomic status (SES) populations are more affected by drinking water violations than are other populations.[7–9] A study of the CWSs of local government–owned utilities that serve populations of 10 000 or more found that communities at or above 40% of the federal poverty level (as determined from the American Community Survey, 5-year estimates, 2010–2014) and with a higher percentage of Hispanics or Blacks had a significant increase in the number of predicted drinking water violations.[10] These studies have increased our understanding of the extent of environmental injustices associated with CWSs. The EPA defines environmental justice as

> the fair treatment and meaningful involvement of all people regardless of race, color, national origin, or income with respect to the development, implementation and enforcement of environmental laws, regulations, and policies. Fair treatment means no group of people should bear a disproportionate share of the negative environmental consequences resulting from industrial, governmental, and commercial operations or policies.[11]

Environmental injustices are a function of environmental discrimination.[9] Environmental discrimination can arise from industry and nonindustry policies and regulations, resulting in underrepresented groups (i.e., racial/ethnic minorities and those of low SES) being exposed to environmental hazards that can directly affect quality of life.[9] For example, in Flint, Michigan, an economically depressed city that has a predominantly Black population, a state-appointed emergency management body decided to introduce a more corrosive water source into an aging water system without also providing adequate corrosion control. The subsequent failure of government officials to respond to health concerns associated with the change in

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drinking water source resulted in a higher elevated blood level percentage among children in Flint as compared to those living outside Flint (i.e., whose water source was unchanged). The areas in Flint that had a higher proportion of Black children had the highest water lead levels and the highest elevated blood levels among children as compared to other areas.[12]

To date, no studies have assessed both the extent to which county-level racial/ethnic and different socioeconomic factors are associated with initial and repeat drinking water violations across the United States. We aimed to (1) assess the extent to which initial drinking water violations differed on the basis of county race/ethnicity and several SES indicators using the primary county served by the CWS as the unit of analysis, and (2) determine whether counties with higher proportions of underrepresented groups are disproportionately burdened with repeat drinking water violations. Additionally, we addressed a gap in the literature by examining differences in initial and repeat drinking water violations on the basis of size of population served by the CWS.[13]

METHODS

The EPA has established National Primary Drinking Water Regulations for the US drinking water supply that pertain to 88 contaminants and drinking water treatment techniques. These regulations are legally enforceable standards that must be monitored monthly or more frequently on the basis of population served by the CWS or previous violations to ensure that maximum contaminant–level thresholds are not exceeded.[1,14,15] The level of safety (i.e., levels that do not cause known or expected health risk) is to be monitored, and safe levels are to be enforced through the collaboration of the EPA, the states, and the CWSs, with the states granted primary control (i.e., "primacy"). A state is granted primacy if it is able to demonstrate that it will comply with at least the minimum SDWA standards within its jurisdiction and that it will report initial violations and repeat violations to the EPA quarterly. With the exception of Wyoming and the District of Columbia (which chose not to apply for primacy), states have primacy status. The EPA oversees drinking water compliance using state-level violation data.[16]

The EPA regularly collects data on drinking water violations and reports this information in the EPA Safe Drinking Water Information System (SDWIS).[17] We downloaded publicly available data on quarterly initial drinking water violations ("initial violations") and repeat water violations ("repeat violations"). We linked these data to the EPA Envirofacts by the public water system ID.[18] By doing this, we constructed a US drinking water database on the basis of previously studied contaminants at the county level.[7,8,12,19–21] The database included CWSs that were active 2011 to 2015 (n = 59 595). We excluded Indian reservations and US territories (n = 1564) because these communities are confronted with documented historical injustices that contribute to water quality, access, and administrative challenges that require separate analysis outside the scope of this study.[22]

Reported initial and repeat violations during the years 2011 to 2015 were the dependent variables and were operationalized as a binary variable (1 if a violation ever occurred for CWS and 0 otherwise). We aligned initial violations and repeat violations data to 2011 through 2015 to match the years we selected from the American Community Survey. We focused on the contaminants previously examined in quantitative studies: arsenic; atrazine; chlorine; total coliforms (pre-TCR and TCR); copper; di(2-ethylhexyl) adipate (DEHA); di(2-ethylhexyl) phthalate (DEHP); haloacetic acids (HAA5); nitrates; nitrite; lead; radium 226 and radium 228 (combined radium); total trihalomethanes (TTHMs); 1,1,1-trichloroethane; 1,1,2- trichloroethane; trichloroethylene (TCE); and uranium (n = 58 031).[4,7,8,12,19–21] These contaminants have potential negative health effects if there is long-term exposure above the maximum contaminant–level threshold, such as damage to the circulatory, cardiovascular, central nervous, respiratory, and reproductive systems; increased risk of cancer; kidney and liver disease; physical and developmental problems; and gastrointestinal illnesses.[15]

We obtained county-level covariates from the 2011 to 2015 American Community Survey because CWSs and regulatory agencies do not collect demographic data.[13] We excluded 13 counties because of data suppression (i.e., the county did not have the minimum number of cases).[23] Contextual environmental justice covariates (units of analysis were county-level proportions) included female-headed households, renter-occupied households, adults with less than high school education, non-Hispanic Blacks, non-Hispanic Asians, Hispanics, non-Hispanic Whites, and uninsured households (i.e., not having medical insurance).[24] We also obtained median income (in $10 000 increments). With the exception of female-headed household and median income, the covariates were not normally distributed; we used log-transformations to account for skewness.

We used multivariable logistic regressions to evaluate reported initial and repeat violations; we considered covariates statistically significant if P was less than .05 (2-tailed) using a 95% confidence interval (CI). We stratified results by size of the population the CWS served on the basis of the stratum federal agencies use for funding purposes (very small: ≤ 3300; small: 3301–10 000; medium: 10 001–50 000; and large: ≥ 50 001).[25,26] We performed data management and analyses using Stata version 14 (StataCorp LP, College Station, TX).[27]

RESULTS

Descriptive results are presented in Table 1. The median percentage of female-headed households was 42.3% (SD = 9.5) and of renter-occupied households was 29.6% (SD = 7.9). The median household income was $49 220 (SD = $13 310). Across counties, the median percentages were 78.6% (SD = 19.9) non-Hispanic White, 3.0% (SD = 12.1) non-Hispanic Black, 1.1% (SD = 3.3) non-Hispanic Asian, 5.9% (SD = 14.5) Hispanic, 12.6% (SD = 5.9) with less than high school education, and 12.9% (SD = 5.3) uninsured.

Results of the nonstratified logistic regression analysis are shown in Table 1. The proportion of uninsured had the highest odds ratio (OR) associated with both reported initial violations and repeat violations (OR = 1.77; 95% CI = 1.64, 1.91 and OR = 1.67; 95% CI = 1.52, 1.82, respectively). The OR of 1.77 means that

TABLE 1—Descriptive Results for County-Level Initial and Repeat Drinking Water Violations: United States, 2011–2015

County-Level Variable	Median (SD)	Initial Violations, OR (95% CI)	Repeat Violations, OR (95% CI)
Proportion female single headed households	0.423 (0.095)	0.813 (0.636, 1.040)	0.468 (0.350, 0.625)
Median household income is $10 000	4.922 (1.331)	0.915 (0.892, 0.938)	0.931 (0.904, 0.959)
Log proportion renter occupied	0.296 (0.079)	0.899 (0.810, 0.998)	1.065 (0.942, 1.204)
Log proportion less than high school education	0.126 (0.059)	0.842 (0.777, 0.913)	0.805 (0.733, 0.885)
Log proportion non-Hispanic Black	0.030 (0.121)	0.899 (0.886, 0.912)	0.907 (0.892, 0.922)
Log proportion non-Hispanic Asian	0.011 (0.033)	1.015 (0.998, 1.033)	0.990 (0.969, 1.010)
Log proportion Hispanic	0.059 (0.145)	0.878 (0.858, 0.898)	0.939 (0.914, 0.964)
Log proportion non-Hispanic White	0.786 (0.199)	0.597 (0.556, 0.642)	0.697 (0.643, 0.756)
Log proportion uninsured	0.129 (0.053)	1.773 (1.641, 1.915)	1.668 (1.524, 1.825)

Note. CI = confidence interval; OR = odds ratio. The sample size was n = 58 018.

a 1-unit increase in the proportion of those uninsured within a county, other things being equal, increased the odds of having an initial violation by 77%. For most other covariates, we found a negative association with both initial violations and repeat violations. For example, a 1-unit increase in the proportion of renters within a county, all else being equal, decreased the odds of only an initial violation by 10.1%. The proportion of renters was negatively associated with initial violations only, the proportion of female-headed households was negatively associated with repeat violations only, and the proportion of non-Hispanic Asians was not associated with either outcome.

Table 2 presents initial violations stratified by population a CWS serves. Median household income remained negatively associated with initial violations for the very small and small strata, but we did not observe a negative association in the very small stratum for proportions of those with less than high school education and for proportions of renters. For these 3 covariates, the negative association with initial violations was most pronounced for large populations served by a CWS. The proportion of female-headed households was significant and changed direction across populations served, with the exception of the proportion of medium populations served (i.e., this was nonsignificant).

Among racial and ethnic groups, the direction of the association with initial violations varied on the basis of the population size served by CWS; this was unobservable in our nonstratified analysis. The exception was the proportion of non-Hispanic Whites, which

TABLE 2—Descriptive Results for County-Level Initial Drinking Water Violations Stratified by Population Served: United States, 2011–2015

County-Level Variable	Very Small (Population ≤ 3300; n = 47 499), OR (95% CI)	Small (Population = 3301–10 000; n = 5657), OR (95% CI)	Medium (Population = 10 001–50 000; n = 3769), OR (95% CI)	Large (Population ≥ 50 001; n = 1093), OR (95% CI)
Proportion female single headed households	0.703 (0.537, 0.919)	2.511 (1.045, 6.031)	1.960 (0.643, 5.975)	0.007 (0.000, 0.155)
Median household income is $10 000	0.933 (0.908, 0.959)	0.795 (0.715, 0.883)	0.987 (0.877, 1.111)	0.317 (0.216, 0.463)
Log proportion renter occupied	1.093 (0.975, 1.226)	0.709 (0.480, 1.048)	0.416 (0.249, 0.693)	0.002 (0.000, 0.012)
Log proportion less than high school education	1.028 (0.943, 1.122)	0.248 (0.181, 0.340)	0.360 (0.246, 0.527)	0.012 (0.004, 0.033)
Log proportion non-Hispanic Black	0.879 (0.866, 0.893)	1.065 (1.009, 1.124)	1.087 (1.005, 1.175)	1.336 (1.049, 1.701)
Log proportion non-Hispanic Asian	1.039 (1.020, 1.057)	0.748 (0.694, 0.806)	0.807 (0.705, 0.922)	2.903 (1.679, 5.017)
Log proportion Hispanic	0.859 (0.838, 0.880)	1.052 (0.954, 1.160)	1.194 (1.036, 1.376)	3.787 (2.200, 6.520)
Log proportion non-Hispanic White	0.593 (0.548, 0.641)	0.559 (0.428, 0.731)	0.894 (0.617, 1.296)	0.437 (0.181, 1.053)
Log proportion uninsured	1.618 (1.489, 1.759)	2.413 (1.791, 3.253)	1.914 (1.337, 2.740)	4.509 (1.360, 14.95)

Note. CI = confidence interval; OR = odds ratio.

was negatively associated with initial violations for nonstratified analysis and for very small and small population groups served by a CWS; this was nonsignificant for medium and large groups. Yet as the population size served by CWS increased, the proportion of non-Hispanic Blacks and Hispanics in a community changed to being positively associated with initial violations. The proportion of non-Hispanic Asians in a community was positively associated with initial violations with very small and large populations served but negatively associated with initial violations for small and medium strata. The proportion of those uninsured was consistent with our nonstratified analysis, and it had a positive OR, with initial violations across all stratification levels.

We found several differences when we compared initial violations to repeat violations across the population served by a CWS (i.e., stratum; Table 3). In particular, covariates shifted significance and reversed direction (i.e., positive to negative OR and vice versa). Of note, the proportion of uninsured residents was the most consistent covariate, as it had a positive OR across every stratum for initial violations and repeat violations, with the exception of the nonsignificant finding for repeat violations within the large stratum. In the very small stratum, the proportion of renters was nonsignificant for initial violations, but it emerged as positively associated with repeat violations (OR = 1.23; 95% CI = 1.08, 1.40). Likewise, in the small stratum, the proportion of Hispanics in a community was nonsignificant for initial violations but had a significant, positive OR with repeat violations (OR = 1.17; 95% CI = 1.04, 1.32).

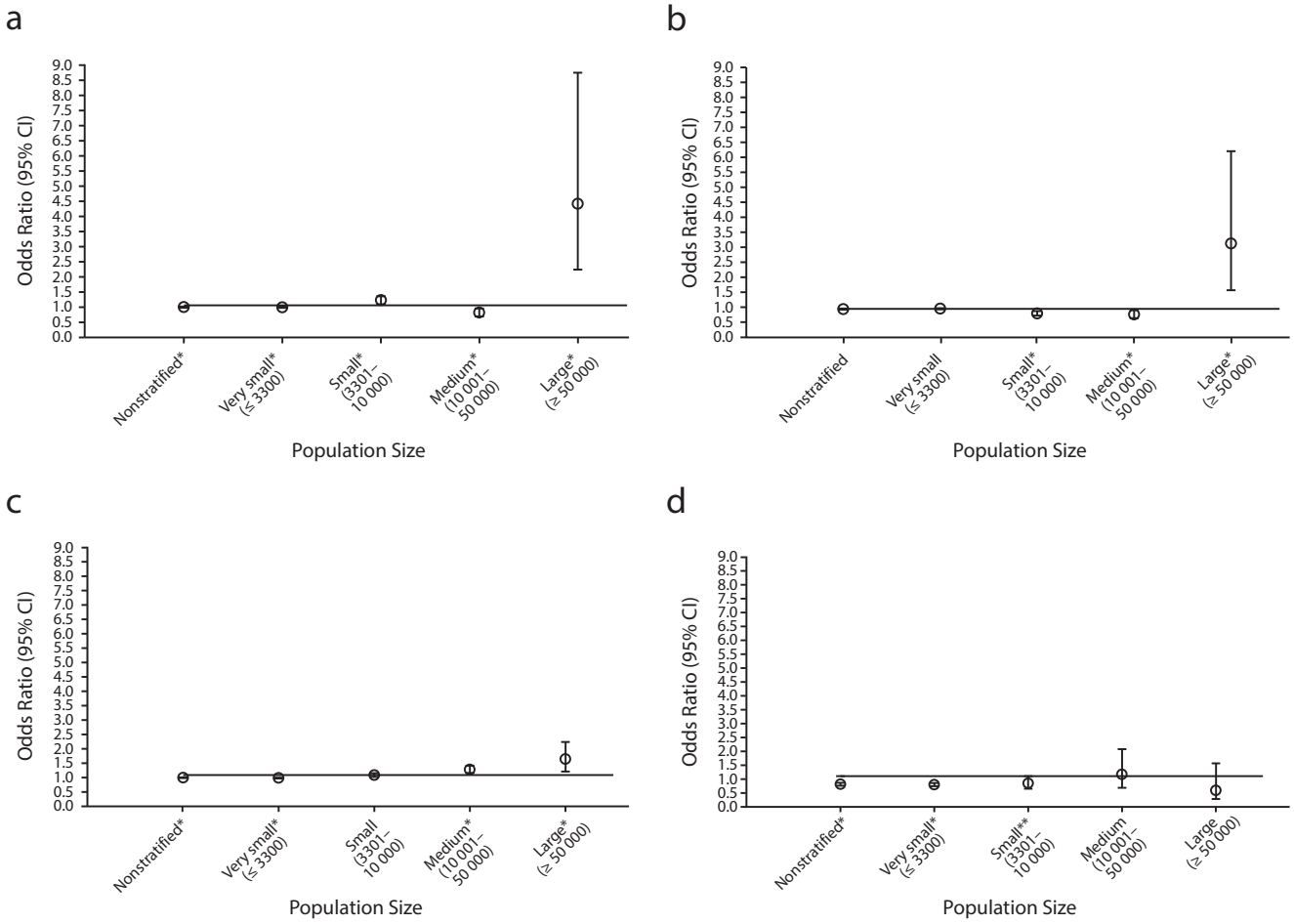
An unexpected finding was that the median income had a positive OR with repeat violations (OR = 1.28; 95% CI = 1.10, 1.50) in the medium stratum, as this association was not observed in other analyses. This means that a $10 000 increase in median household income within a medium stratum, other covariates held constant, increased the odds of having a repeat violation by 28%. Another noteworthy result was that the proportion of Hispanics in a community changed direction from being positively associated (OR = 1.19; 95% CI = 1.04, 1.38) with initial violations to negatively associated with repeat violations (OR = 0.75; 95% CI = 0.63, 0.91) for medium populations served by a CWS. The following covariates were significant and either positively or negatively associated as denoted with initial violations but were nonsignificant for repeat violations: the proportion of female-headed households for small (positive) and large (negative) strata, the proportion of renters and the proportion of those with less than high school education for medium stratum (negative), and the proportion non-Hispanic Blacks for small stratum (positive).

Figure 1 shows the different OR patterns of nonstratified and stratified racial/ethnic subgroups for repeat drinking water violations. For example, except for non-Hispanic Asians, nonstratified ORs were significant and negatively associated with repeat drinking water violations (i.e., as depicted by the 1.0 OR reference line). Because CWSs varied in size by population served, we found that, overall, the proportion of non-Hispanic Whites remained significant and negatively associated with repeat drinking water violations, whereas the proportions of other racial/ethnic groups did not. The large CWSs showed the most pronounced ORs for proportion of minority groups. This is of particular interest because nonstratified analysis has previously emphasized drinking water quality in very small and small systems. Our findings reveal that minority groups served by large systems have increased significant odds for repeat violations.

TABLE 3—Descriptive Results for County-Level Repeat Drinking Water Violations Stratified by Population Served: United States, 2011–2015

County-Level Variable	Very Small (Population ≤ 3300; n = 47 499) OR (95% CI)	Small (Population = 3301–10 000; n = 5657) OR (95% CI)	Medium (Population = 10 001–50 000; n = 3769) OR (95% CI)	Large (Population ≥ 50 001; n = 1093) OR (95% CI)
Proportion female single headed householder	0.458 (0.335, 0.626)	1.491 (0.497, 4.470)	0.400 (0.090, 1.772)	0.052 (0.001, 2.165)
Median household income is $10 000	0.942 (0.913, 0.971)	0.814 (0.717, 0.924)	1.284 (1.097, 1.503)	0.326 (0.203, 0.523)
Log proportion renter occupied	1.229 (1.076, 1.403)	0.930 (0.574, 1.509)	0.753 (0.376, 1.506)	0.003 (0.000, 0.017)
Log proportion less than high school education	0.920 (0.833, 1.017)	0.324 (0.222, 0.474)	0.682 (0.405, 1.146)	0.015 (0.004, 0.051)
Log proportion non-Hispanic Black	0.897 (0.881, 0.913)	0.999 (0.935, 1.067)	1.197 (1.074, 1.334)	1.563 (1.127, 2.168)
Log proportion non-Hispanic Asian	1.002 (0.981, 1.024)	0.837 (0.764, 0.918)	0.800 (0.674, 0.949)	3.277 (1.650, 6.506)
Log proportion Hispanic	0.932 (0.907, 0.959)	1.171 (1.036, 1.323)	0.753 (0.626, 0.906)	4.399 (2.198, 8.804)
Log proportion non-Hispanic White	0.683 (0.626, 0.746)	0.731 (0.536, 0.999)	1.065 (0.571, 1.988)	0.477 (0.155, 1.469)
Log proportion uninsured	1.491 (1.355, 1.641)	2.243 (1.560, 3.224)	2.974 (1.794, 4.931)	4.147 (0.886, 19.42)

Note. CI = confidence interval; OR = odds ratio.



Note. CI = Confidence interval; OR = odds ratio. The test for interaction confirmed heterogeneity across the 4 populations served by community water system categories OR. We performed the test by allowing populations served by community water system categories to vary with log proportion of race/ethnicity subgroups.
*$P < .05$; **$P < .01$.

FIGURE 1—Comparison of Stratified and Nonstratified ORs for County-Level Repeat Drinking Water Violations by Proportion of (a) Hispanics, (b) Non-Hispanic Asians, (c) Non-Hispanic Blacks, and (d) Non-Hispanic Whites: United States, 2011–2015

DISCUSSION

We have presented an original analysis of initial and repeat drinking water violations using an environmental justice lens and have 3 key findings. First, the majority of initial violations occurred among systems serving very small (≤ 3300 inhabitants) and small (3301–10 000 inhabitants) populations. This finding aligns with results from previous work.[4,25] Our second finding, as posited by VanDerslice,[13] revealed that analysis by size of population served by a CWS revealed results masked by nonstratified analysis. The third and novel key finding was that lower SES and minority groups are associated with an increased OR for initial and repeat drinking water violations.

To our knowledge, our study is the first to examine contextually relevant environmental justice covariates associated with both initial drinking water violations and repeat violations stratified by population served by CWS. The most similar drinking water research focused on nonstratified drinking water violations for inactive and active CWSs. Previous studies found CWSs that served a larger percentage of Hispanics and renters had higher nitrate levels in the drinking water[7] and increased odds for arsenic violations.[8] Conversely, selective implementation and enforcement of the revised 2002 SDWA arsenic standard in Arizona did not reveal disproportionate impacts on minorities or low-income groups.[19] We found that the association between initial and repeat drinking water violations and race/ethnicity and SES covariates varied in direction and strength on the basis of the stratum. The single consistent association irrespective of stratification was between initial and repeat drinking water violations and the proportion of uninsured residents. These results further support the strong relationship between uninsured status and environmental justice because those lacking insurance are potentially most likely to be disproportionately exposed to hazards.[9] Our finding aligns with previous environmental justice research that reported that cumulative cancer risks increased in direct proportion with a higher percentage of uninsured populations.[28]

Separately, median household income results for stratified and nonstratified analyses predominantly support previous environmental justice research, which found that high-income areas are more likely to have better water quality.[10,29] An unexplained finding was that median household income was nonsignificant for initial violations in the medium stratum but had a significantly increased OR for repeat violations. It is noteworthy that the proportion of renters was found to be negatively associated with initial violations among medium and large strata. A possible explanation for this is the complexity of the housing market, whereby the majority of renters are of modest income but occupants span the income spectrum and the geographic variations in renting.[30] Contrary to our expectations, the proportion of those with less than a high school education decreased the odds of initial violations across small, medium, and large strata. A potential explanation for this is that this population may reside within a multigenerational household[31] in which some members have attained higher educational levels, thereby altering the results.

Overall, our results provide new information about how population size served by a CWS may affect the odds of county-level race/ethnicity proportions associated with initial and repeat violations. We observed the decreased OR associated with initial and repeat violations for non-Hispanic Whites among only very small and small strata. Furthermore, the race/ethnicity proportion results revealed that with the exception of non-Hispanic Asians, ORs associated with initial and repeat violations were decreased for all other subgroups at the very small stratum. This is noteworthy, because this stratum had the highest percentage of violations, but we still observed the lowest OR among non-Hispanic Whites.

The proportion of Hispanics in a community emerged as having a significantly increased OR in the small stratum for repeat violations, which was not present for initial violations. Conversely, for this covariate in the medium stratum, initial violations had an increased OR, but repeat violations had a decreased OR. However, minorities face significant challenges with exposure to poor water quality. This finding was most notable among CWSs that serve large populations for initial and repeat violations.

Stratified analysis revealed that the proportion of female-headed households in a community had a statistically negative or positive association with initial and repeat violations depending on the stratum and had a nonsignificant association in the nonstratified analysis, which is an important finding. In the United States, the relationship between living in a female-headed household, environmental justice, and drinking water quality is underexamined, despite inclusion of female-headed household as a covariate of interest in studies measuring increased risk of exposure to harmful environmental factors. Our inconsistent findings across strata reflected contradictory US-based research results of a positive association of female-headed household and poor air quality[24,32] as well as no association.[28] To develop a full picture of the role of living in a female-headed household and drinking water quality, we recommend that additional studies include this covariate.

These findings may help us to understand how environmental justice and drinking water quality are linked and that although the majority of violations occur in CWSs that serve very small and small populations, medium and large strata should not be overlooked. We found that the proportion of uninsured residents and of race/ethnicity were associated with an increased OR, whereas median household income was associated with a decreased OR for initial violations. We observed similar trends for repeat violations. Our results do not reveal that counties with a higher proportion of racial/ethnic subgroups or households with lower SES were disproportionately burdened with repeat violations. It is notable that the ORs for the proportion of non-Hispanic Blacks and Hispanics were higher for repeat violations than for initial violations in the large stratum.

Strengths and Limitations

Our study contributes to the body of knowledge on environmental justice and drinking water quality, as it is the first US-based study, to our knowledge, that examines county-level risk factors for both initial drinking water violations and repeat violations using contextual covariates typically associated with environmental justice issues. Another strength is that we used stratification categories that state- and federal-level agencies use for funding and reporting purposes,[25,26] unlike past studies.[8,10]

A limitation of this study is that states may upload unreliable data on both violations and enforcement actions to the SDWIS.[33] As of August 23, 2017, the EPA is still working on the implementation of 2011 US Government Accountability Office recommendations to improve state water quality reporting and programs through regional offices providing guidance and oversight.[34] However, the SDWIS is the most comprehensive publicly available national database to examine US drinking water quality. Moreover, this study combines data from SDWIS with information from other databases, creating a unique data set from which to study drinking water violations. Another limitation of this research is the ecological study design. Our study was limited by the lack of available georeferenced CWS data to conduct an analysis of disparities (i.e., race/ethnicity and SES) at a more granular level (i.e., below the county level).

Public Health Implications

Our results suggest that drinking water stakeholders (including consumers, administrators, and public health practitioners) need to make greater efforts to ensure that populations with a higher proportion of minorities, of people lacking medical insurance, and of low-income households have access to safe drinking water, irrespective of the population size served by the CWS. Considering the association between drinking water contaminants and potential negative health outcomes,[15] it is imperative that populations exposed to poor water quality be identified to prevent public health outbreaks. Although Congress has mandated that the nation's public water supply be regulated to protect public health, the public has little knowledge of how water regulation actually works.[1] Any effort to improve access to safe drinking water is moot without reliable and timely reporting of initial violations and repeat violations.

There are 3 key initiatives that can contribute to this effort. First, it is essential to create awareness among the general public about how the SDWA operates, including

information on the provision for citizens to take civil action against any federal agency or EPA administrator allegedly in violation of the SDWA.[16]

Second, citizen engagement is required to stem the current tide of calls by government leaders for deregulation.[35] Action would include pressing local representatives as well as Congress to invest in upgrading and maintaining infrastructure (e.g., water pipes and water pumps), technical assistance and training of CWS personnel, and additional reporting resources (e.g., investment in personnel and computing tools) to assist the EPA with SDWA oversight to improve compliance and enforcement of violations. Such investment would improve the reporting of initial violations and repeat violations so that the CWS, the state, and the EPA know where and what corrective action is required to ensure that the SDWA goal to protect public health for all populations is realized.

Third, public health practitioners need to be involved in the process to evaluate the effectiveness of the state primacy practice.[4] Public health practitioners can use county-level results to address drinking water quality concerns to respective EPA regional and state officials of populations affected by initial violations and repeat violations. As discussed by Weinmeyer et al.,[1] the current application of the SDWA has resulted in a burdensome, unreliable, and redundant system that at best is challenging to manage and diminishes the EPA's ability to monitor drinking water violations and identify those at greatest risk for poor drinking water quality. AJPH

CONTRIBUTORS

Y. J. McDonald drafted the introduction, strengths and limitations, and public health implications sections. N. E. Jones performed the data analysis. Both authors designed the research, collected and interpreted the data, drafted the methods, results, and discussion sections, critically revised the article, and approved the final version of the article.

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HUMAN PARTICIPANT PROTECTION

No protocol approval was necessary because no human participants were involved in this study.

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