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A spatiotemporal analysis of the association of California city and county cannabis policies with cannabis outlet densities

Suggested running head: Local cannabis policies and outlet densities

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Data access and statistical computing code: The cannabis policy and dispensary data are not currently available, as the funded work utilizing these newly collected data is still ongoing. The datasets will be released upon completion of the aims of the grant funding the work. Statistical code is provided in the appendix.

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Abstract

Background: Cannabis outlets may affect health and health disparities. Local governments can regulate outlets, but little is known about the effectiveness of local policies in limiting outlet densities and discouraging disproportionate placement of outlets in vulnerable neighborhoods.

Methods: For 241 localities in California, we measured seven policies pertaining to density or location of recreational cannabis outlets. We geocoded outlets using web-scraped data from the online finder Weedmaps between 2018 and 2020. We applied Bayesian spatiotemporal models to evaluate associations of local cannabis policies with Census block group-level outlet counts, accounting for confounders and spatial autocorrelation. We assessed whether associations differed by block group median income or racial–ethnic composition.

Results: Seventy-six percent of localities banned recreational cannabis outlets. Bans were associated with fewer outlets, particularly in block groups with higher median income, fewer Hispanic residents, and more White and Asian residents. Outlets were disproportionately located in block groups with lower median income (posterior RR [95% credible interval]: 0.76 [0.70, 0.82] per \$10,000), more Hispanic residents (1.05 [1.02, 1.09] per 5%), and fewer Black residents (0.91 [0.83, 0.98] per 5%). For the six policies in jurisdictions permitting outlets, two policies were associated with fewer outlets and two with more; two policy associations were uninformative. For these policies, we observed no consistent heterogeneity in associations by median income or racial–ethnic composition.

Conclusions: Some local cannabis policies in California are associated with lower cannabis outlet densities, but are unlikely to deter disproportionate placement of outlets in racial–ethnic minority and low-income neighborhoods.

Keywords: Cannabis; marijuana; legalization; local policy; local control; local ordinances; outlet density; availability; outlets; retail

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Introduction

As of November 2021, recreational or “adult-use” cannabis is legal in 18 states and the District of Columbia.¹ Cannabis policies regulate the availability of cannabis by legally permitting outlets offering cannabis products for retail sale. Alcohol availability research indicates that higher residential outlet densities make it easier to find, purchase, and use legal intoxicants.^{2,3} Analogously, greater availability of medical cannabis dispensaries has been linked to cannabis use and frequency.^{4,5} Similar effects are expected for recreational cannabis outlets.² Increases in cannabis access and use may have both positive and negative health consequences. Cannabis consumption has been linked to motor vehicle crashes, psychotic disorders, respiratory disease, low birth weight, and cannabis use disorder, but substitution of opioids, tobacco, or alcohol for cannabis may prove beneficial.⁶⁻⁹ Outlets may also attract crime, although research on this topic is mixed.¹⁰⁻¹³

State cannabis legalization policies typically defer authority to regulate the density and locations of outlets to local governments. Local governments can limit the number of outlets permitted, establish minimum distances between outlets, and bar their location near sensitive locations such as schools. Local governments also share responsibility with state agencies for abating illegal outlets which are prevalent in California.^{14,15} The impacts of local cannabis policies on outlet densities may have implications for public health by limiting availability. Recreational cannabis outlets are disproportionately located in neighborhoods with high proportions of low-income and racial–ethnic minority residents.¹⁵⁻¹⁹ Policies that encourage greater reductions in outlets in vulnerable neighborhoods therefore have the potential to promote health equity.

Little is known about the impacts of local cannabis policies. Three studies assessed local policies in Colorado, Washington, and California following recreational cannabis legalization.^{20–}

²² All identified broad variation in local regulatory approaches, ranging from all-out bans to unlimited outlets, with a few jurisdictions allowing outlets while limiting their densities. To our knowledge, no prior study has evaluated how local policies influence outlet densities or socioeconomic and racial–ethnic equity in the distribution of outlet densities within jurisdictions.

We addressed these gaps with a spatiotemporal analysis of city and county cannabis policies and cannabis outlets in California. We evaluated whether specific local policies such as density limits cannabis outlets led to lower outlet densities. We also assessed whether the associations of local policies with outlet densities varied across neighborhoods depending on median income or racial–ethnic composition. We hypothesized that stricter local policies would be associated with lower outlet densities and less disproportionate placement of outlets in less-advantaged communities. Cannabis legalization research suggests that provisions enabling outlets are influential for cannabis consumption and related health outcomes.^{23–25} We focus on the local-level policies that determine how many outlets can open and in which communities. Understanding which local policies effectively limit and equalize outlet densities is critical for state and local policymakers seeking to make more informed decisions about which cannabis policies to pursue to protect public health and health equity from potential harms related to legal cannabis.

Methods

Overall approach

We assessed local cannabis policies in 241 city and county jurisdictions across California. We merged these data with annual Census block group-level measures of cannabis

outlet densities and potential confounders. Using the resulting dataset of block groups nested within city and county jurisdictions, we examined within-block group changes in outlet densities from 2018 to 2020, beginning with the implementation of recreational cannabis retail sales on January 1, 2018. No human subjects were involved in this study.

Cannabis policy data and measures

We classified local cannabis policies for 12 of California's 58 counties representing 59% of the state population. The 12 counties were selected to capture a range of sizes, sociodemographic compositions, political orientations, and approaches to cannabis regulation,²⁰ and included 230 cities and 11 unincorporated county areas (San Francisco city and county constitute a single government).

Using a legal epidemiological approach,^{26,27} between November 2020 and January 2021 we systematically identified and coded the characteristics of currently applicable cannabis policies in all 241 jurisdictions. We used a structured data collection instrument to capture the presence or absence and content of pre-specified provisions. Two analysts coded all jurisdictions separately until they achieved >95% agreement. Complete protocols, data collection instruments, and further detail are provided in eAppendices 1-3.

California state law specifies a minimum set of regulations that apply to cannabis statewide. However, localities retain considerable discretion. The policy measures we collected were guided by an established taxonomy of all possible cannabis policies.²⁹ We coded all policies that: (a) were regulated at the local level, (b) varied across jurisdictions, (c) were more restrictive than state law, and (d) were plausibly related to public health given prior evidence, public health best practices, and expert opinion.^{20,21,29}

The exposures were cross-sectional measures of the seven binary policy measures that directly restricted the number, density, or locations of storefront recreational cannabis outlets (Table 1).

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Cannabis outlet data and measures

The outcome was the count of storefront recreational cannabis outlets (hereafter, “outlets”) in each Census block group and year. We web-scraped data on outlets annually between 2018 and 2020 from Weedmaps, a high-traffic online promotional cannabis business finder widely used in cannabis research.^{4,16,30,31} A prior validation study found that, compared to official license listings or other finders, Weedmaps was the most up-to-date and comprehensive source for capturing cannabis outlets.¹⁴ We focused on recreational rather than medical outlets because: following recreational legalization, few medical-only outlets remained; the applicable state laws for medical outlets are distinct; and Weedmaps measures of medical outlets were less valid over the study period. Recreational outlets included both newly opened outlets and outlets that converted from medical to recreational. We focused on storefront (brick-and-mortar) outlets, as opposed to home delivery retailers, because this study builds on conceptual models based on physical proximity to outlets offering in-person purchases.³ See eAppendix 3 for detail (“Cannabis outlet measurement”).

Web-scraped data included the outlet address and offerings (medical or recreational cannabis; storefront sales or home delivery). We geocoded the addresses to 2019 block groups using the ArcGIS World Geocoding Service (>99% success rate). are the most granular geographic level at which the Census Bureau reports demographic and socioeconomic characteristics, and are considered reasonable proxies for neighborhoods.³²

Covariates

Covariates included in the adjustment set were factors hypothesized to confound the policy–outlets relationship.³³ Potential confounders were measured at the block group-year level and included measures of demographic composition, socioeconomic factors, the commercial

environment, a local alcohol outlet policy stringency score, and the percent of voters favoring recreational cannabis legalization. eTable 1; <http://links.lww.com/EDE/B940> provides detail on the data sources and procedures for each covariate.

We conceptualized race–ethnicity as socially-defined categories that reflect the distribution of risk, opportunities, and discrimination.³⁴ Racial–ethnic groups were not mutually exclusive: Asian, Black, and White racial groups were defined irrespective of Hispanic identity, and the Hispanic group included people of any. Primary analyses adjusted for the proportions of Asian, Black, and Hispanic residents. Analyses considering effect measure modification by the racial–ethnic composition also utilized percent White.

Database development

City policies apply within city borders, and county policies apply to the unincorporated areas of counties outside cities. To assign the block group-level outlet data to jurisdiction-level policy data, we overlaid shapefiles of block group, city, and county boundaries in ArcGIS Pro (see eAppendix 3 “Database development”). We excluded three jurisdictions with no residential populations. We excluded an additional 30 block groups due to covariate missingness. The final analytic dataset included 13,979 block groups nested within 238 city and unincorporated county jurisdictions.

Statistical analysis

To quantify the association of local policies with outlet densities, we used a hierarchical Bayesian spatiotemporal Poisson regression. This approach uses conditional autoregressive random effects to account for spatial autocorrelation in outlet densities across neighboring block groups that otherwise gives incorrect statistical inferences (i.e., block groups adjacent to one another are likely to have similar outlet counts and covariates, violating the independence-of-

units assumption of standard statistical approaches). The model specification is presented in Box 1. We modeled outlet counts relative to the expected count assuming a distribution directly proportional to land area to reflect physical access.^{35,31,36} The primary associations of interest were the areal relative risks (RRs) of outlets associated with each policy. We included block group-level spatially structured random intercepts to account for dependence of neighboring units, block group-level random intercepts and slopes assuming independence-of-units to allow the level and linear trend in outlets to vary independently for each block group, and jurisdiction-level random intercepts to account for time-constant characteristics of jurisdictions.

Box 1. Bayesian spatiotemporal model specification

$$Y_{jit} | \mu_{jit} \sim \text{Poisson}(E_{jit} \exp(\mu_{jit}))$$

$$\mu_{jit} = \beta_0 + \beta_1 P_{1jt} + \beta_2 P_{2jt} + \dots + \beta_k P_{kjt} + \beta_{k+1} X_{jit} + \alpha_t + \omega_{ji} + \theta(t)_{ji} + \psi_{ji} + \phi_j$$

i : block group ($i=1, \dots, 13,979$)

j : city/unincorporated county ($j=1, \dots, 238$)

t : year ($t=0,1,2$)

Y_{jit} : number of outlets

E_{jit} : expected outlet count assuming a distribution directly proportional to land area ($E_{jit} =$

$$L_{jit} \frac{\sum Y_{jit}}{\sum L_{jit}} \text{ where } L_{jit} \text{ is the corresponding land area)$$

$\exp(\mu_{jit})$: areal relative risk of outlets (RR)

β_0 : intercept

$P_{1jt}-P_{kjt}$: binary policy variables ($k=1, \dots, 6$)

$\exp(\beta_1), \dots, \exp(\beta_k)$: estimated areal relative risk (RR) indicating the associations of each policy P_k with outlets

X_{jit} : matrix of confounders

β_{k+1} : vector of the associations of the confounders with outlets

α_t : year fixed effects to account for secular trends

ω_{ji} : block group random effects intercepts assuming independence of units (spatially unstructured)

$\theta(t)_{ji}$: block group linear random slopes assuming independence of units (spatially unstructured)

ψ_{ji} : spatially structured random effects intercepts allowing outlet counts for neighboring block groups to be dependent, with neighboring block groups defined as those that shared a border (i.e., “queen adjacencies”: in the neighborhood structure matrix, cells take the value of 1 if block groups i and i' share a border and 0 otherwise)

ϕ_j : jurisdiction random effects intercepts assuming independence of units

First, to characterize places with outlets, we fit spatiotemporal models with each covariate in turn as the only fixed predictor. Then we estimated the associations of the policies with outlets, adjusting for all covariates. We considered two sets of policy effects: First, associations for outlet bans among all study areas, and second, for the jurisdictions permitting outlets (6,291 block groups in 56 jurisdictions), associations for the 6 policies regulating outlet density/location. For both sets, we estimated the overall association of the policies with outlets and used interaction terms to test whether the associations varied by block group median income or racial–ethnic composition. To report interaction results, we computed associations for block

groups at the 25th and 75th percentile of each moderating variable. For all estimates, we report the marginal posterior means and 95% credible intervals.

Following recommended practice and prior empirical work, we implemented estimation using Integrated Nested Laplace Approximation with the INLA package within R version 4.0.4.³⁷⁻⁴⁰ We used the “BYM2” spatiotemporal model instead of the typical BYM or Leroux specification because this method better handles non-contiguous county “islands” and generates clearly interpretable parameters.^{41,42} Based on reference guides and prior empirical work, we used the INLA default priors.³⁷⁻⁴⁰ We considered a five-unit change in the Watanabe-Akaike information criterion (WAIC) to indicate improved model fit.^{43,44} Statistical code is provided in eAppendix 3.

Secondary and sensitivity analyses

Because policies regulating the density or location of outlets are particularly relevant to urban areas, we considered models restricted to cities, excluding unincorporated county areas. Second, we tested models with expected counts of outlets proportional to population instead of land area ($E_{jit} = M_{jit} \frac{\sum Y_{jit}}{\sum M_{jit}}$ where M_{jit} is the corresponding population). Third, because we did not observe all possible combinations six density- or location-related policies relevant to jurisdictions that permitted outlets (eTable 2; <http://links.lww.com/EDE/B940>), we summarized the combined effects of the six policies by estimating models replacing the individual policy variables with a summed policy count score (range 0-6). Fourth, we tested whether removing random effects led to better model fit.

Results

Table 2 presents characteristics of the study block groups. The study covered 24 million people with varied demographics, socioeconomic positions, commercial environments, and

political orientations. Of the 238 jurisdictions with residential populations, 182 (76%) banned outlets. Outlet bans were more common in jurisdictions with more White residents, higher median income, and less poverty. Among the six policies applicable to jurisdictions allowing outlets, the most common were buffers around sensitive locations (86%), location restrictions (77%), and density limits (55%) (Table 2). Limits on the overconcentration of outlets in vulnerable neighborhoods (10%) and buffers around alcohol outlets (2%) were rare. Nearly half of jurisdictions allowing outlets (41%) required buffers between one cannabis outlet and another. Across the study jurisdictions, the total number of outlets increased from 170 in 2018 to 390 in 2020. Five percent of outlets were in jurisdictions that banned them, reflecting gaps in implementation, enforcement, and grandfathering.

Table 3 presents the associations of each block group characteristic with observed outlet counts relative to expected. Throughout the study period, most block groups had 0 outlets—fewer than the number expected assuming a distribution directly proportional to land area. There were more outlets than expected in places with fewer Black residents (RR: 0.91 [95% CI: 0.83, 0.98] per 5%), more Hispanic residents (RR: 1.05 [95% CI: 1.02, 1.09] per 5%), and lower median income (RR: 0.76 [95% CI: 0.70-0.82] per \$10,000), as well as more poverty, less education, fewer family households, more renters, more unemployment, more alcohol outlets, stricter alcohol policies, and more pro-cannabis voters.

eTable 4; <http://links.lww.com/EDE/B940> presents the estimated hyperparameters for the fully adjusted spatiotemporal models. Across models, the proportion of the marginal variance in the block group random intercepts explained by the BYM2 spatially structured block group random intercepts (as opposed to the block group random intercepts for which we assume independence and identical distribution) ranged from 0.02 to 0.53.

The Figure presents the adjusted associations of outlet bans with cannabis outlet counts, overall and by neighborhood median income and racial–ethnic composition. As hypothesized, bans were associated with substantially lower outlet counts (RR: 0.04 [95% CI: 0.01-0.11]). These associations were more pronounced for block groups at the 75th percentiles of median income (RR: 0.02 [95% CI: 0.00-0.06]), percent White residents (RR: 0.01 [95% CI: 0.00-0.03]), and percent Asian residents (RR: 0.02 [95% CI: 0.00-0.06]), and at the 25th percentile of percent Hispanic residents (RR: 0.01 [95% CI: 0.00-0.04]). We improved model fit by incorporating interaction terms between outlet bans and median income, percent Asian, percent Hispanic, or percent White ($\Delta WAIC > 9.6e23$), but not percent Black. Results from sensitivity analyses restricting to cities and towns, with expected outlet counts proportional to population, and using alternative combinations of random effects were consistent with the main results (eFigures 1-2; <http://links.lww.com/EDE/B940>; eTable 5; <http://links.lww.com/EDE/B940>).

eFigure 3; <http://links.lww.com/EDE/B940> presents the adjusted associations for the six policies available to local jurisdictions that did not ban outlets. Associations between each policy and outlet counts varied, and were generally imprecise. Outlet counts were lower in jurisdictions adopting location restrictions (RR: 0.66 [95% CI: 0.16-2.72]) and buffers between outlets (RR: 0.57 [95% CI: 0.16-2.01]). In contrast, outlet counts were higher in jurisdictions that placed buffer zones around sensitive location such as schools (RR: 2.78 [95% CI: 0.38-24.39]) and limits on overconcentration in vulnerable areas (RR: 2.46 [95% CI: 0.42-12.87]). For density limits (RR: 1.01 [95% CI: 0.30-3.26]) and buffers around alcohol outlets (RR: 1.17 [95% CI: 0.02-62.24]), estimates were uninformative. There was some heterogeneity in policy associations by block group median income and racial–ethnic composition, but not in a consistent direction. For these models, incorporating interaction terms between the policies and percent Hispanic

improved model fit ($\Delta WAIC > 8.5e21$), but not incorporating interaction terms between policies and median income or other racial–ethnic composition variables did not.

In sensitivity analyses restricted to cities and towns, density limits, location restrictions, sensitive location buffers, alcohol outlet buffers, and buffers between outlets showed no discernible association with outlet counts, but overconcentration limits remained associated with more outlets (eFigure 4; <http://links.lww.com/EDE/B940>). Results for sensitivity analyses with expected outlet counts proportional to population were similar to the main results (eFigure 5; <http://links.lww.com/EDE/B940>). In models utilizing the sum of the six policies as the primary exposure, greater policy stringency was associated with a moderate but imprecise reduction in overall outlets (RR: 0.60 [95% CI: 0.29-1.24]), with more pronounced but imprecise associations for block groups with high (75th percentile) proportions of Hispanic residents (RR: 0.49 [95% CI: 0.23-1.02]) and Asian (RR: 0.45 [95% CI: 0.21-0.95]) residents (eFigure 6; <http://links.lww.com/EDE/B940>). Models removing block group random slopes and spatially structured and unstructured block group random intercepts fit the data better than models including these components, but there were no substantive differences in the estimated associations (eTable 5; <http://links.lww.com/EDE/B940>). The one exception was for density limits, for which removing the block group random effects changed the RR (95% CI) from 0.99 (0.31, 3.33) to 1.54 (0.46, 5.41).

Discussion

In this spatiotemporal analysis of city and county cannabis control policies, we found that local policies banning outlets were strongly associated with lower geographic densities of recreational cannabis outlets. In jurisdictions that did not ban outlets, we evaluated the potential for specific local policies to limit densities and promote equitable distribution of outlets. Here,

our findings were mixed: some policies were associated with fewer outlets and others with more, but estimates were imprecise. Outlets disproportionately opened in block groups with more Hispanic residents and less socioeconomic advantage, yet local policies restricting outlets did not appear to counteract this pattern. Instead, in jurisdictions adopting outlet bans, the lower outlet counts were most pronounced for block groups with higher incomes, and more White and Asian residents. For jurisdictions permitting outlets, the six policy associations followed no consistent pattern in terms of the most-affected block groups. These findings are important for public health and health equity because if city and county policies can effectively limit outlet densities, they may encourage safer population levels of consumption.^{3,20,45} To promote health equity, such policies would need to encourage greater reductions in outlets in vulnerable neighborhoods.

Our finding that outlets disproportionately opened in block groups with more Hispanic residents and less socioeconomic advantage is likely driven by the disproportionate absence of outlet bans in these places. These findings are consistent with prior research reporting similar patterns for California, Colorado, Washington, Oregon, and Canada.¹⁵⁻¹⁹ Economic theory suggests that outlets are likely to open in low-income areas (where retail rents are lower) but adjacent to high-income areas (where demand is highest) because this placement maximizes sales opportunities while minimizing operating costs.⁴⁶ Although we are not aware of any evidence that the economic benefits of outlets accrue to the neighborhoods where outlets are located, outlets may offer economic opportunities for community members. This idea has motivated explicit efforts by some localities to prioritize retail licenses for individuals and communities negatively impacted by the past criminalization of cannabis.^{20,45,47} Yet, to the extent that outlets are harmful to health—this is still an open question—regulators should be concerned about the potential implications of the uneven distribution of outlets for health equity.

Although most localities in our study banned cannabis outlets, some outlets persisted in banned areas. Policies are rarely universally effective, or perfectly and equally enforced. Outlets may be present in places with local bans for several reasons, including enforcement gaps and overriding laws that grandfathered licenses to outlets in banned areas. Still, outlet bans appear to be a highly effective tool for communities seeking to control the proliferation of outlets.

Although outlet bans apply to all block groups within the jurisdiction, outlet bans appeared more effective in areas with more social advantage (higher median income, more White residents). The frequency and consequences of differential enforcement across neighborhoods should be investigated.

For the six local policies limiting outlet densities and locations, the magnitudes of most associations were meaningful, but there was insufficient statistical support to make firm conclusions. Imprecision arose because most jurisdictions banned outlets, outlets were rare, and spatial autocorrelation was high. If results are truly null, this would be unsurprising, as many well-meaning policies are ineffective. If the estimates are real differences, any interpretations are conditional on meeting the assumptions necessary for causal inference (no unmeasured confounding, positivity, and no interference). The negative associations we observed for location limits and buffers between outlets may reflect effective policies. The positive associations we observed for sensitive location buffers and overconcentration limits may reflect reverse causation whereby policies are adopted in response to high concentrations of outlets and are either ineffective or have not yet had time to work. A central challenge here is disentangling the causal effects of policies from confounding—whereby advantaged communities adopt restrictive policies.

If results for the six policies are real causal effects, they did not appear to systematically benefit socially advantaged block groups. This might be expected because high socioeconomic status, White, and other advantaged groups may use their disproportionate political power to exclude cannabis outlets from opening in their neighborhoods.⁴⁸ Residential segregation along racial–ethnic and socioeconomic lines set the stage for “not-in-my-backyard” (NIMBY) activism.^{49–51} NIMBY initiatives have thwarted public health equity on issues ranging from homelessness to AIDS, alcohol control, substance use treatment, and air pollution.^{49,52–54} Cannabis legalization has raised concern that NIMBYism and other mechanisms of structural racism⁵⁵ would lead to regulations that protected White, advantaged communities from outlets while increasing density in non-White or disadvantaged communities. If estimated associations for the six policies reflect causation, the findings suggest that these policies are unlikely to counteract inequitable distributions of outlets (though some were explicitly designed to do this), but also unlikely to exacerbate inequalities. Local policymakers seeking to address the inequitable distribution of outlets may need to test alternative strategies.

Our findings are also important in light of research showing that recreational outlets are co-located with alcohol outlets.^{16–18} High densities of alcohol outlets are associated with binge drinking, crime, and injuries; are disproportionately located in marginalized communities; and can be regulated by local policies.^{56,57,28} New cannabis outlets generate potential for dual-burden harms associated with the spatial co-location of cannabis and alcohol outlets, particularly in communities with less power to deter this activity. Siloed policy approaches—rather than integrated approaches that consider co-location—may further exacerbate problems, yet we found only one locality that regulated the locations of cannabis outlets in relation to alcohol outlets. To

be cautious, localities should consider policies regulating co-location of alcohol and cannabis outlets, and the health implications of alcohol–cannabis outlet co-location should be assessed.

Limitations

Our inability to incorporate cannabis home delivery is an important limitation of this study. Methods for operationalizing access to cannabis delivery remain undeveloped, but cannabis delivery constitutes a growing portion of the retail market, a pattern accelerated by the COVID-19 pandemic.⁵⁸ Given that most jurisdictions banning outlets also ban delivery businesses, the associations we have observed may be relevant to delivery businesses as well, but this should be evaluated empirically in future research.

Other limitations include the potential for uncontrolled confounding. We may have also underestimated effect measure modification by controlling for confounders of the policy–outlet relationship that are also on the pathway from median income or racial–ethnic composition to outlet densities. Additionally, illegal outlets may be undercounted in our data in 2020, because legal action in the previous year encouraged Weedmaps to purge listings of illegal outlets. We assessed local policies cross-sectionally in 2020 and assumed them to be time-invariant over the study period. Policies may have been adopted several months or years prior to 2020. We could not assess how within-place temporal changes in policies affected outlet densities, either immediately or lagged. We modeled the temporal relationships between outlet densities and time-varying covariates such as sociodemographics, but we could not model other temporal dynamics, including whether a recreational outlet was previously medical-only versus newly opened. Reverse causation, in which local policies are adopted in response to outlet densities, is also possible. However, the cannabis norms and political orientations that determine local policies are unlikely to change substantially and systematically over the 3-year study period. We

focused on a subset of California, which limits generalizability. Nonetheless, our study areas captured the majority of the California population and diverse approaches to cannabis regulation. Although block groups are very small spatial units, it is possible that analyses at other levels of spatial aggregation could produce different results (the “modifiable area unit problem”). Some mismeasurement of spatial effects is possible because block groups at the edge of the study regions lacked measurements for all neighbors, but any bias is likely to be small because this concern applies to only a small minority of study areas. Finally, we define “equity” as the absence of differential associations between policies and outlets by block group median income and racial–ethnic composition, but other measures may also be appropriate.

Conclusions

As with all policies, cannabis legalization likely involves balancing harms and benefits. For jurisdictions that have chosen to legalize recreational cannabis, the optimal density of outlets is unknown. If lessons from alcohol and tobacco apply to cannabis, limiting outlet densities may protect public health.^{2,20,59–63} Alternatively, if cannabis outlets promote substitution of alcohol, tobacco, or opioids for cannabis, and these substances are less harmful than cannabis, then health may be improved.^{8,9,64–66}

Local control of legal cannabis has resulted in considerable variation in cannabis policies across California with important implications for health equity. This analysis suggests that bans on outlets were disproportionately adopted in jurisdictions with more White residents, higher median income, and less poverty, and this pattern has resulted in the disproportionate placement of cannabis outlets in less advantaged communities. Moreover, although local policies in jurisdictions permitting cannabis outlets have the potential to address inequitable distributions of cannabis outlets, those policies adopted to date do not appear to have achieved this. Findings

from this study should be incorporated into broader assessments of the costs and benefits of recreational cannabis legalization considering short-term and long-term public health and social welfare outcomes. Alternative policy and public health approaches that protect vulnerable communities from disproportionate harms related to cannabis should be explored.

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Figure titles and legends

Figure: Adjusted associations of bans on cannabis outlets with cannabis outlet densities, estimated from Bayesian spatiotemporal models, overall and by median income and racial–ethnic composition, California, 2018-2020

Legend: Reported values are the posterior mean and posterior 95% credible intervals for the model parameters estimated in INLA. Estimates by median income and racial–ethnic composition correspond to block groups at the 25th and 75th percentiles of median income and racial–ethnic composition. All models were adjusted for demographic composition (total population, population change, age, and race–ethnicity), socioeconomic factors (educational attainment, poverty, median income, unemployment, home ownership, family households), commercial environment (per capita densities of general retail businesses and payday loan, tobacco, and pawnshop businesses; off-premise, restaurant, and bar/pub alcohol outlet densities), a local alcohol outlet policy stringency score, and the percent of voters favoring recreational cannabis legalization as a proxy for pro-cannabis norms.

Table 1: California city and county policies regulating the number, density, and locations of recreational storefront cannabis outlets

Local policy	Description
<i>Policies applicable in all jurisdictions</i>	
Ban on outlets	Retail sales of cannabis through outlets are permitted statewide with a state-issued license. However, localities can ban outlets from operating within their borders.
<i>Policies applicable in jurisdictions without bans</i>	
Density limits	No statewide density limits exist, but localities can adopt such restrictions. Density limits include caps on the number of cannabis outlets that are permitted in the jurisdiction, by count, square mile, or per capita.
Geographic buffers around sensitive locations	Statewide, outlets must be at least 600 feet away from schools, daycares, and youth centers. Localities can mandate larger minimum distances or expand the list of sites considered to be sensitive locations.
Location restrictions	Beyond buffers around sensitive locations, the state places no additional restrictions on where outlets can be located. Localities can further restrict outlet placement, beyond what is allowed for retail businesses generally—for example, requiring that outlets be located only on one street or in one specific commercial zone.
Limits on overconcentration in vulnerable neighborhoods	Statewide, determinations of whether to grant, deny or renew a retail license involve considering whether there exists an “excessive concentration” of outlets in the area where the licensee will operate. Localities can prohibit the establishment or renewal of outlets in or adjacent to low-income neighborhoods, areas of high crime, areas with existing high densities of outlets, or other vulnerable neighborhoods.
Geographic buffers around alcohol outlets	Alcohol sales are banned inside cannabis outlets throughout the state. Localities can restrict where outlets are located in relation to alcohol outlets (e.g., not in the same strip mall) or require that outlets be placed a minimum distance away from alcohol outlets.
Geographic buffers between outlets	The state places no restrictions on how far apart outlets must be from one another. Localities may require that outlets be spaced a minimum distance apart.

Table 2: Characteristics of study cities and unincorporated county areas, overall and by bans on outlets, California, 2020

Characteristic	All jurisdictions	Bans dispensaries	Permits dispensaries
Jurisdictions (N)	238	182	56
Block groups (N)	13,979	7,688	6,291
Total population (persons)	24,315,643	13,839,708	10,475,935
Demographics (median [Q1, Q3][Q1, Q3])			
Median age	38 (32, 44)	37 (32, 43)	40 (33, 47)
% female	51 (49, 53)	51 (49, 53)	51 (49, 53)
% non-Hispanic Asian	6 (1, 16)	6 (1, 17)	6 (1, 16)
% non-Hispanic Black	1 (0, 6)	1 (0, 4)	2 (0, 9)
% Hispanic	38 (16, 71)	37 (17, 70)	39 (15, 72)
% non-Hispanic White	55 (39, 76)	57 (43, 77)	50 (34, 75)
Socioeconomic status (median [Q1, Q3])			
% with high school degree or GED	17 (11, 24)	18 (12, 25)	16 (10, 23)
% with some college or Associate's degree	28 (20, 37)	31 (22, 39)	26 (18, 34)
% with Bachelor's degree or higher	23 (11, 40)	23 (12, 39)	23 (11, 42)
Median income (in \$1000's)	62 (43, 89)	68 (47, 94)	55 (38, 81)
% below 150% of federal poverty level	23 (11, 41)	19 (9, 35)	28 (14, 48)
% unemployed	4 (3, 6)	4 (3, 6)	5 (3, 6)
% renters	36 (16, 61)	28 (12, 53)	47 (24, 67)
% family households	75 (62, 83)	79 (68, 85)	69 (53, 79)
% population change since 2000	8 (-1, 16)	7 (-1, 15)	9 (-1, 17)
Commercial environment (median [Q1, Q3])			
General retail outlet density (per sq mi)	2,552 (926, 5,487)	2090 (742, 4434)	3256 (1245, 6828)
Density of payday loan, tobacco, and pawnshop businesses (per 10 sq mi)	34 (0, 419)	0 (0, 308)	109 (0, 583)
Alcohol outlet density (per 10 sq mi)	54 (0, 199)	39 (0, 158)	82 (0, 263)
Off-premise alcohol outlet density (per 10 sq mi)	10 (0, 102)	3 (0, 80)	22 (0, 135)
Bar/pub outlet density (per 10 sq mi)	0 (0, 0)	0 (0, 0)	0 (0, 0)
Restaurant alcohol outlet density (per 10 sq mi)	0 (0, 78)	0 (0, 61)	0 (0, 106)
Policy environment (median [Q1, Q3])			
% voting for recreational cannabis legalization	56 (52, 65)	53 (51, 55)	65 (60, 65)
Alcohol outlet density policy stringency score	1 (0, 2)	1 (0, 2)	1 (1, 4)
Adoption of cannabis policies (N [%])			
Outlet ban	182 (76)	182 (100)	0
Outlets permitted	56 (24)	0	56 (100)
Density limit	31 (13)	0	31 (55)
Location restriction	43 (18)	0	43 (77)
Buffers around sensitive locations	48 (20)	0	48 (86)

Limit on overconcentration in vulnerable neighborhoods	6 (3)	0	6 (11)
Buffers around alcohol outlets	1 (0)	0	1 (2)
Buffers between outlets	23 (10)	0	23 (41)
Storefront recreational cannabis outlets			
Number, 2018	170	9	161
Number, 2019	349	24	325
Number, 2020	390	21	369
Density per 10 square miles (mean [min, max])	2 (0, 380)	0 (0, 166)	4 (0, 380)

Legend: Abbreviations: Q1: 25th percentile. Q3: 75th percentile. Statistics reported in this table were calculated across the 13,979 study block groups nested within city and unincorporated county jurisdictions in 2020. eTable 1 provides detail on the data sources and data processing for each covariate.

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Table 3: Bivariate associations of census block group characteristics with cannabis outlet densities, estimated from Bayesian spatiotemporal models, California, 2018-2020

Block Group Characteristic	Outlet relative risk (posterior mean [95% credible interval])
Year	
2018	(ref)
2019	0.01 (0.00, 0.02)
2020	5.5e-5 (1.2e-5, 2.0e-4)
Population (per 10,000 persons)	0.43 (0.07, 2.4)
Median age (years)	1.02 (0.99, 1.04)
Racial and ethnic composition	
% non-Hispanic Asian ^a	0.96 (0.89, 1.03)
% non-Hispanic Black ^a	0.91 (0.83, 0.98)
% Hispanic ^a	1.05 (1.02, 1.09)
% non-Hispanic White	0.99 (0.95, 1.03)
Median income (per \$10,000)	0.76 (0.70, 0.82)
% below 150% of federal poverty level ^a	1.15 (1.10, 1.20)
Education	
% with high school degree or GED ^a	1.03 (0.95, 1.11)
% with some college or Associate's degree ^a	0.99 (0.92, 1.06)
% with Bachelor's degree or higher ^a	0.93 (0.89, 0.98)
% family households ^a	0.76 (0.72, 0.80)
% renters ^a	1.21 (1.17, 1.25)
% unemployed ^a	1.30 (0.90, 1.85)
% population change since 2000 ^a	1.03 (0.96, 1.10)
General retail outlet density (per 10,000 persons)	1.02 (0.97, 1.05)
Density of payday loan, tobacco, and pawnshop businesses (per 100 persons)	1.02 (0.97, 1.06)
Total alcohol outlet density (per 1000 sq mi)	1.07 (1.06, 1.09)
% bar/pub alcohol outlets ^a	1.15 (1.09, 1.20)
% off-premise alcohol outlets ^a	1.03 (1.00, 1.05)
Alcohol outlet density policy stringency score ^b	1.51 (0.97, 2.4)
% voting for recreational cannabis legalization ^a	3.7 (2.6, 5.6)

Legend: Reported values are the posterior mean and posterior 95% credible intervals for the model parameters estimated in INLA, using each covariate in turn as the only fixed predictor. eTable 1 provides detail on the data sources and procedures for each covariate. Associations for year are negative because: (1) the models include block group-level random slopes, which help us to account for unmeasured confounding resulting from temporal correlations between block group policy implementation and block group-specific secular trends in the outcome (i.e., the impacts of heterogeneous growth on the fixed parameter estimates of policy effects); and (2) the outlet counts are modeled relative to the expected count of outlets assuming a distribution directly proportional to land area. Most block groups have no cannabis outlets, but the

expected count for all outlets is a small number greater than 0, so most block groups have fewer outlets than expected for all time periods. Superscripts: ^a Percentage variables were formulated in units of 5 percentage points. ^b Local alcohol policy data were collected using procedures identical to those described for local cannabis policies. Using the subset of policy measures that directly dictate the number, density, or locations of alcohol outlets, the alcohol outlet density policy stringency score was calculated using the weighting scheme developed by Thomas and colleagues.²⁸

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Figure 1

