Inequality in Agency Responsiveness: Evidence from Salient Wildfire Events

October 7, 2020

Abstract

Government agencies may be an additional source of unequal representation beyond that stemming from the differential responsiveness of elected officials. In this paper, we use plausibly exogenous focusing events, which raise public demands for government provision of local public goods, to examine whether there is evidence for unequal responsiveness in agency decision-making. Using the empirical case of wildfire risk management in the western U.S., we find that when some communities experience nearby wildfire events, it raises the salience of wildfire risk and leads agencies to place a greater number of wildfire risk reduction projects nearby, even when wildfire risk has already been reduced. Importantly, this effect predominates among high socioeconomic status communities. We find that nearby fires increase rates of fuels treatment particularly among higher income, more educated, and whiter communities. The formal model and empirical evidence show that public agencies perpetuate inequality, via the costs of lobbying and the costs of agency lack of responsiveness that vary with demographics.

Keywords: responsiveness, representation, inequality, bureaucracy, public goods, salience, public land management

Scholars have long recognized a potential tension between equitable outcomes and government responsiveness: a responsive government has potential to be more responsive to some citizens than to others, which may result in inequitable policy outcomes in domains from the environment (Ringquist, 2005) to education (Meier et al., 1999). A substantial body of literature has examined how policymaker responsiveness (in terms of roll-call votes, communication, or enacted policies) to constituent preferences varies across demographic groups (e.g. Gilens, 2005; Bartels, 2008; Butler and Broockman, 2011). Yet policy outcomes only begin with legislators and legislation; downstream, disparities in responsiveness in the bureaucratic administration of policy can have implications for inequality as well. Agencies can be an alternative mechanism for unequal representation due to differential enforcement (Konisky and Reenock, 2018) or via their distribution of community-level resources. In this paper, we study responsiveness of agencies to demands for community-level resources from demographically-varying communities.

A number of possible channels can produce inequality in agency implementation decisions. Agency implementation, governed by administrative procedures (McCubbins et al., 1987), is a function of legislative constraints (Huber and Shipan, 2002), the political context (Aberbach, 1990), public pressure, and the agency's own preferences (Bawn, 1995; Anderson and Potoski, 2016; Epstein and O'Halloran, 1994). In this paper, we focus on identifying the role of public pressure. A formal model of agency decisionmaking in the face of lobbying shows how differences in costs of political mobilization or penalties communities can impose on non-responsive agencies, which we argue are correlated with demographic attributes of communities, can shape agency responsiveness and produce inequality.

Distinguishing the empirical role of a particular channel in contributing to inequality requires an exogenous shock to identify the effect of that channel. Our quasi-experimental difference-in-differences design makes use of an exogenous shock to demand for wildfire risk reduction projects that occurs after communities in the western US experience nearby wildfires. These exogenous focusing events increase attention to agencies' actions to reduce fire risk. We compare differences in fuel treatment rates, a primary tool federal and state land management agencies use to reduce wildfire risk by removing hazardous dead and live vegetation from the landscape, across demographically-varying communities before and after the occurrence of nearby fires. After including Census block fixed effects and county-year fixed effects, our empirical specification identifies differences in agency responsiveness to demographically-varying communities, ruling out the effects of legislative constraints, preferences of elected officials, and the preferences or biases of the agencies (Einstein and Glick, 2017; White et al., 2015). This unique and important empirical setting isolates the role of differences in community demographics in producing inequality of agency implementation.

We find that recent wildfires increase federal fuels project rates near high socioeconomic status (SES) communities, and that increases in project rates are strongest for higher income, more educated, and whiter communities. For example, we find a one standard deviation (10 percentage point) increase in the percentage of residents above the poverty line leads to up to a 40 percent increase in the likelihood that a community receives a fuels treatment. This is evidence that communities better able to demand resources from the federal government, whether because they are better able to mobilize or because they can better impose penalties on agencies for not meeting their demands, receive disproportionate benefits. In the case of wildfire, these benefits reduce fire intensity (Kalies and Kent, 2016) and, potentially, damages from wildfire (Kennedy and Johnson, 2014).

The paper proceeds as follows. In section 1, we discuss wildfire management in the western U.S., the role of fuel reduction projects, and the planning process used by federal agencies for determining how and where to situate these projects. This process includes significant opportunities for public comment and influence. Section 2 identifies the potential mechanisms for unequal distribution of agency resources, including differential responsiveness to communities. Section 3 develops a formal model that describes how inequality may result from differences in communities' abilities to mobilize or differences in their abilities to penalize agencies when they fail to meet community demands. This motivates the paper's empirical strategy, which—along with the data—is discussed in section 4. In section 5, we present our results. We conclude by discussing implications and limitations of this research, and potential paths forward for future research.

1 Wildfire Fuels Projects and Federal Land Management

Over the past several decades, wildfire activity has sharply increased within the western U.S. (Dennison et al., 2014; Westerling, 2016). Researchers have generally attributed this increase in wildfire activity to the combined effects of climate change (e.g., Westerling et al., 2006; Abatzoglou and Williams, 2016) and high fuel loads within western forests (Arno et al., 1995; Keane et al., 2002; Naficy et al., 2010). In many dry forest types (e.g., ponderosa pine forests within the U.S. southwest and Sierra Nevada mountains), high fuel loads, which are partially attributed to the legacy of fire exclusion over the course of the twentieth century, have led to larger and more severe wildfires (Schoennagel et al., 2004).

Fuels projects are local public goods intended to reduce wildfire risk by restoring the forest to conditions under which high intensity fires are less likely. Fuel reduction is achieved by burning ("prescribed fire") or physically removing ("mechanical thinning") accumulated vegetation. Empirical evidence (reviewed in Kalies and Kent, 2016) indicates that fuel reduction projects within dry forests in the western U.S. are effective at reducing fire intensity, especially when prescribed fire is used in conjunction with thinning. There is also some evidence that strategically-placed fuels projects can help prevent damage to homes and structures (Bostwick et al., 2011; Kennedy and Johnson, 2014). Between 2009 and 2018, Congress appropriated approximately $500 million per year to the U.S. Forest Service (USFS) and Department of Interior land management agencies (GAO, 2019). Nevertheless, only two to four million of acres of National Forest land are restored each year, while the USFS estimates that 85 million acres of National Forest land are in need of forest restoration.[1]

Federal agencies are therefore budget-constrained, and they may face competition among residential areas for their limited resources. Although scientific management is a foundational doctrine

[1] Forest restoration encompasses a broad range of management activities intended to restore ecological function, including prescribed burns and mechanical thinning.

of the USFS, previous research indicates that USFS and other federal land management agencies' decision-making is also influenced by public pressure (Sabatier et al., 1995; Johnson and Watts, 1989; Anderson et al., 2013). This may in part be due to the participatory decision-making structures that have defined federal land management planning since the passage of the National Environmental Policy Act (NEPA) in 1970. NEPA mandates that all federal agencies document actions that will significantly impact the environment with an Environmental Impact Statement (EIS). Further, it mandates a public comment period during which the public can voice opinions on a proposed action. Similarly, the National Forest Management Act of 1976 mandates that the public be allowed opportunities to comment on forest management plans. This openness to public input likely affects fuels project planning process. According to Hakanson (2010), forest managers often have an eye toward the NEPA process from a fuels project's conception.

2 Mechanisms and Evidence for Inequity in Agency Implementation

Legislatures face a basic tradeoff between political control and technical competence when they delegate authority to agencies (Bawn, 1995; Ringquist et al., 2003). Broadly speaking, agency implementation decisions are a function of agency preferences, including preferences derived from agency expertise, and political factors, the balance of which is determined in part by the legislature's decisions about adminstrative procedures (Epstein and O'Halloran, 1994). An agency makes decisions on the basis of the authority delegated to them by the legislature in organic statutes or particular legislation, the current political landscape (Wood and Bohte, 2004; Anderson and Potoski, 2016), public input (Anderson et al., 2013), and the agency's own preferences. Any one of these components could contribute to inequality (Dion et al., 1998; Scholz and Wang, 2006;

Konisky, 2009). In practice, it is impossible to separately observe each contribution to agency decisionmaking. As a result, it is difficult to identify the source of inequality.

Agency responsiveness to the public may lead to inequitable outcomes when (1) policy preferences differ across groups and (2) government responsiveness varies across groups (Wlezien and Soroka, 2011). Since the first condition is generally taken as given, research has focused on examining the second condition. Political participation varies across groups (e.g., Verba et al., 1995), with higher SES individuals participating at higher rates. Wealthier individuals (Rosenstone and Hansen, 1993; Ojeda, 2018; Erikson, 2015), white residents (Griffin et al., 2019; Bowler and Segura, 2011), and higher educated people (Rosenstone and Hansen, 1993; Nie et al., 1996) participate in politics at higher rates. These groups may apply greater pressure to politicians and government officials. Another reason to expect that government officials might respond differentially to high SES groups is that government officials tend to be relatively high income, highly educated individuals, and they may be more sympathetic to the views of similar individuals (Page et al., 2013). As the next section develops with a formal model, different communities have different benefits from lobbying, costs of lobbying, and abilities to make not meeting their demands costly to the government, all of which may lead to differential responsiveness on the part of the government.

Motivated by these ideas, a variety of studies over the past fifteen years have tested for inequality in responsiveness among policymakers, usually by following Gilens (2005) in relating political outcomes (e.g., roll-call votes, legislation) to constituent opinions across the income distribution. Most of these studies use either cross-sectional data (e.g., McCarty et al., 2009; Grimes and Esaiasson, 2014) or time-series data (e.g., Ura and Ellis, 2008; Wlezien and Soroka, 2011), and thus cannot identify differences in responsiveness across income groups from differences in policy preferences that are correlated across income groups over space or time. One exception uses an exogenous shift in responsiveness caused by a change in how local assessors are chosen to show that towns with elected assessors are less likely to perform property value reassessments, which tend to increase the effective tax rate paid by owners of high-value homes (Sances, 2016). Throughout the literature on government responsiveness and inequality, data limitations have led to difficulties in identifying differences in responsiveness across communities. Like Sances (2016), we add to this literature by making use of panel data to credibly identify these differences in responsiveness.

3 Demographics, Lobbying, and Agency Responsiveness

This section presents a model of local public good provision with lobbying. The decision-maker is a government agency that provides a local public good to a community. Communities, which have differential costs of lobbying that can depend on their SES status, can lobby the government for more of that public good based on their perceived value for it. The government agency makes allocation decisions on the basis of the costs of the good and of not meeting the community's demand. We show here that communities with a lower cost of lobbying, perhaps because of higher SES status, and communities that can place higher costs on government for not meeting their demands, again perhaps because of higher SES status, are allocated more of the public good. This differential responsiveness may produce inequality (Griffin and Newman, 2008; Franko, 2013) between higher SES communities with more political efficacy (Verba et al., 1995) and lower SES communities.

Formally, the cost of allocating Q units of the good is $C(Q) = \frac{1}{2}\eta Q^2$ where η is a positive parameter. The agency allocates Q_0 units of the public good to the community to satisfy its own programmatic and budgetary objectives. The public good provides benefits to the community given by $B(Q) = bQ$, where b is the marginal benefit from the good. Because members of the community do not bear the costs of public goods provision, they have an incentive to lobby for more than Q_0. We allow for the amount of community lobbying to depend on the community's perceived value $\tilde{b}$

of the public good, which may depart from its true value b. We assume the community and agency play a leader-follower game in which the community lobbies for the good, taking into account the best response function of the agency.[2] The cost of lobbying is $C_L(Q_L) = \frac{1}{2}\alpha Q_L^2$, where α is a positive parameter and Q_L is the additional amount of the good sought by the community. One can think of Q_L as being proportional to lobbying effort.[3] The community receives Q_0 when it does no lobbying, and so it finds the optimal Q_L by solving:

$$\max_{Q_L} \tilde{b}(Q_0 + Q_A(Q_L)) - \frac{1}{2}\alpha Q_L^2 \tag{1}$$

where $Q_A(Q_L)$ is the community's conjecture about the additional amount of the public good it will obtain from seeking Q_L. The solution to (1), Q_L^*, is given implicitly by $\tilde{b}\frac{dQ_A}{dQ_L} - \alpha Q_L^* = 0$.

When a community lobbies, the agency incurs a cost of not meeting the community's demand, given by $C_A(Q_A) = \frac{1}{2}\gamma(Q_L^* - Q_A)^2$ where γ is a positive parameter. The agency is the follower in the model and so it assumes that its choice of Q_A does not affect Q_L^*. It chooses the additional allocation to the community, Q_L, that minimizes the costs of providing more of the good and not meeting the community's demand, solving[4]:

$$\min_{Q_A} \frac{1}{2}\eta(Q_0 + Q_A)^2 - \frac{1}{2}\eta Q_0^2 + \frac{1}{2}\gamma(Q_L^* - Q_A)^2. \tag{2}$$

The solution is $Q_A^*(Q_L^*) = \frac{\gamma Q_L^* - \eta Q_0}{\eta+\gamma}$, which implies $\frac{dQ_A}{dQ_L} = \frac{\gamma}{\eta+\gamma}$. When the community lobbies for more of the public good, the agency increases the provision of the good but does not fully meet the

[2]Our model has the same structure as a Stackelberg industry. We make the community the leader in the game so that it conjectures that lobbying affects the provision of the public good.

[3]We express Q_L in terms of units of the public good because it allows us to define the agency's loss function, below, in the same units.

[4]The term $-\frac{1}{2}\eta Q_0^2$ appears in (2) so that costs reflect the *additional* costs to the government of providing the public good. The restriction $\tilde{b}\gamma^2 > \alpha(\eta + \gamma)\eta Q_0$ ensures that a strictly positive value of Q_A is chosen.

community's demand ($\frac{\gamma}{\eta+\gamma} < 1$). By substitution, we obtain:

$$
\begin{aligned}
Q_L^* &= \frac{\tilde{b}}{\alpha} \frac{\gamma}{\eta+\gamma} \\
Q_A^* &= \frac{\tilde{b}}{\alpha} \left[\frac{\gamma}{\eta+\gamma} \right]^2 - \frac{\eta}{\eta+\gamma} Q_0
\end{aligned}
\tag{3}
$$

In sum, the community lobbies for more of the public good ($Q_L^* > 0$) than the agency would have allocated for its own programmatic and budgetary reasons. The agency responds by providing an additional amount of the public good ($\frac{dQ_A}{dQ_L} > 0$). A community with higher perceived marginal benefits ($\tilde{b}$) will lobby more and be allocated more of the public good:

$$
\begin{aligned}
\frac{dQ_L^*}{d\tilde{b}} &= \frac{1}{\alpha} \frac{\gamma}{\eta+\gamma} > 0 \\
\frac{dQ_A^*}{d\tilde{b}} &= \frac{1}{\alpha} \left[\frac{\gamma}{\eta+\gamma} \right]^2 > 0
\end{aligned}
\tag{4}
$$

This yields *Hypothesis 1*: that communities whose perceived marginal benefits of government provision of a local public good increase will receive more of it.

Likewise, a community with a lower cost of lobbying (α) will lobby more and receive more of the good:

$$
\begin{aligned}
\frac{dQ_L^*}{-d\alpha} &= \frac{\tilde{b}}{\alpha^2} \frac{\gamma}{\eta+\gamma} > 0 \\
\frac{dQ_A^*}{-d\alpha} &= \frac{\tilde{b}}{\alpha^2} \left[\frac{\gamma}{\eta+\gamma} \right]^2 > 0
\end{aligned}
\tag{5}
$$

Finally, a community that imposes a higher penalty on the government (γ) will lobby more and receive more of the good:

$$
\begin{aligned}
\frac{dQ_L^*}{d\gamma} &= \frac{\tilde{b}}{\alpha} \frac{\eta}{(\eta+\gamma)^2} > 0 \\
\frac{dQ_A^*}{d\gamma} &= \frac{2\tilde{b}\gamma\eta}{\alpha(\eta+\gamma)^3} + \frac{\eta Q_0}{(\eta+\gamma)^2} > 0
\end{aligned}
\tag{6}
$$

Together, these last two results imply *Hypothesis 2*: that communities with higher socioeconomic status, which has been shown to be correlated with higher political efficacy (e.g. Verba et al., 1995), receive higher allocations of local public goods. This higher political efficacy likely manifests in both a lower cost of lobbying and a better ability to impose higher costs on the government if it does not meet their demands.

4 Methods

4.1 Empirical Strategy

Our empirical strategy identifies differences in responsiveness across communities with different demographic characteristics using exogenous shocks to the salience of wildfire risk—and thus the perceived benefits of fuels projects—caused by the occurrence of recent nearby wildfires. Salience is a behavioral phenomenon in which individuals disproportionately weight concerns that have drawn their attention (Taylor and Thompson, 1982). For example, prices of homes in high flood risk areas are lower than homes outside these areas, but only in years after a flood has occurred nearby (Bin and Landry, 2013). Salient disaster events are often referred to as focusing events and have been shown to influence political agenda-setting (Birkland, 1997).

Recent nearby wildfires increase the salience of wildfire risk within communities (McCoy and Walsh, 2014); their emotional interest, concreteness, or temporal, spatial, and sensory proximity and visibility shape the ease with which information is brought to mind, which can influence or bias judgments (Kahneman, 2003; Taylor and Thompson, 1982; Tversky and Kahneman, 1973). In prior work (Wibbenmeyer et al., 2019), we found that federal wildfire fuels management projects are more likely to be placed near communities that have experienced recent wildfires. We attribute this pattern to salience of wildfire risk in these areas and the ensuing public pressure community members place on agencies. Where wildfire risk is more salient and therefore the perceived benefits for fuels projects are higher, homeowners and community members may seek more agency actions to reduce wildfire hazard. Here, we use the occurrence of wildfires as a shock to public pressure and use this to identify differential rates of bureaucratic responsiveness across demographic groups.

Formally, we model dependent variable y_{it}, which measures the placement of fuels projects in the area surrounding Census block i within the Wildland Urban Interface (WUI) in year t, using a standard difference-in-differences framework. We measure y_{it} in two ways, as described in section 4.2. We take WUI blocks as treated if they have experienced a nearby wildfire in the past three years. We choose three years as the relevant cut-off because the salience of wildfire events is short-lived and does not drive fuels project decision-making after about 3 years (Wibbenmeyer et al., 2019).[5] We define nearby fires as those occurring within 2 km (and 5 km for robustness) of a WUI block, since very nearby fires will be most salient to homeowners and most likely to drive increases in public pressure (see also Stokes (2016)). Therefore, we define the variable $recentfire_{it}$ as equal to 1 if WUI block i experiences a wildfire within 2 km in the past three years and zero otherwise. We write the difference-in-difference specification as:

$$y_{it} = \alpha_i + \beta recentfire_{it} + recentfire_{it} \times \mathbf{x}_i' \boldsymbol{\gamma} + \delta_{rt} + \varepsilon_{it}. \tag{7}$$

The coefficient β describes the main effect of a recent fire on the placement of fuels projects. Because we are interested in how responsiveness to salient wildfire events varies with demographics across communities, we allow the effect to vary with demographic characteristics $\mathbf{x}_i$. The degree to which the effect of wildfire occurrence varies with demographic variables is captured by the $K \times 1$ vector of parameters $\boldsymbol{\gamma}$. Given that demographic variables are standardized to a distribution with mean

[5] This finding is also consistent with other empirical work on the effects of salient disaster events on home prices, e.g., McCoy and Walsh (2014).

zero and a standard deviation of 1, every element γ_k of $\boldsymbol{\gamma}$ can be interpreted as the increase in responsiveness due to a one standard deviation increase in demographic variable k.

In order for β and $\boldsymbol{\gamma}$ to be identified, it is required that there exist no unobserved factors that affect the likelihood of fuels projects being located near WUI block i and are correlated with the occurrence of a recent fire. Due to amenity-driven sorting, higher socioeconomic status individuals may be more likely to live in areas with higher wildfire risk and higher fuels project rates (Stetler et al., 2010). Furthermore, areas with higher wildfire risk are more likely to have experienced recent wildfires and are more likely to be chosen as the location for fuels reduction projects. To account for differences in the fuels project rates across blocks due to fixed natural or social charactieristics, we include block-level fixed effects α_i.

Still, a threat to identification would exist if wildfire risk facing individual WUI blocks were to vary over time in a way that was correlated with block demographic characteristics. To guard against this possibility, we include a set of county-by-year fixed effects—denoted δ_{rt}, where r indexes counties—which account for differences across counties and within years in fuels project rates. After including block and county-by-year fixed effects, we identify β and $\boldsymbol{\gamma}$ with variation in differences between departures from the within county-by-year average fuels project rate for block i in year t, and the average departure from the within county-by-year fuels project rate across all years for block i.

We argue that our empirical specification allows us to attribute observed differences in post-fire treatment rates to inequality in agency responsiveness to communities, as opposed to statutory limitations, political oversight, or the agency's own preferences. Differences in the degree to which statutory requirements stemming from the delegation of authority to the agency by Congress or embodied in administrative procedures enable fuels treatments in particular locations should be accounted for by Census block-level fixed effects. County-year effects account for changes in the statutory context over time, including varying changes across counties. In order for the statutory context to explain our findings, statutory limitations would need to shift the within-county distribution of fuel treatments toward locations correlated with higher SES and at locations and times correlated with the occurrence of a nearby fire. In a similar way, we expect that political oversight of agencies (e.g., oversight pressure from members of Congress, see Aberbach (1990)) may affect within-county fuel treatment location decisions, and that political oversight may change over time and across locations, but these effects will be accounted for by the combination of block-level fixed effects and county-year effects included in our model. Finally, differences in agency preferences explain our results only if relative preferences across potential within-county fuel treatment locations change at times and locations that experience nearby fires in ways that are positively correlated with SES. In sum, the most likely factor to be driving within-county shifts in the distribution of fuel treatments toward locations near high SES blocks after nearby fires is post-fire changes in demands for fuel treatments and differences in agency responsiveness to those demands across communities.

Fuels project rates are spatially correlated, both due to underlying spatial correlation in wildfire risk and mechanically due to the way in which our dependent variables are constructed. Our dependent variables are defined as a function of the placement of fuels projects within some distance from a given block. However, the same fuels project may increase the project rates for multiple adjacent WUI blocks. Moreover, treatment is not randomly assigned to blocks. It is spatially correlated, since a fire that occurs near one block also occurs in the proximity of adjacent blocks. To account for non-independence among observations within our sample of blocks, we cluster standard errors by Census tract. Census tracts are generally quite large within the western U.S. Our sample consists of of more than 6 million blocks, but contains only 6,096 tracts across 487 counties.

4.2 Data

The units of observation are U.S. Census blocks from 15 western states.[6] We focus specifically on blocks classified as WUI in 2000, since these are communities that are likely to face wildfire risk. Because we are interested in determinants of public fuels management project locations, we further limit our sample of Census blocks to those within 10 km of public lands managed by the USFS, Bureau of Land Management (BLM), or National Park Service (NPS). The USFS, BLM, and NPS together manage approximately 1.5 million square kilometers of land in the western US and are responsible for 93% of federal fuels management projects within the timespan of our data. After these restrictions, our data comprises more than 350,000 WUI census blocks.

Data on fuels treatment locations come from the National Fire Plan Operations and Reporting System (NFPORS). The NFPORS dataset records the point location (latitude and longitude), dates, and area of all fuels reduction projects conducted by the USFS, BLM, and NPS during years 2003-2011. Since NFPORS does not provide fuels project boundaries, we used reported point locations and project areas to impute project boundaries, under the assumption that project boundaries are circular. We compare variation across WUI blocks in the degree to which fuels projects are placed nearby, measured in two ways. First, we define the dependent variable y_{it} as an indicator for whether any fuels projects were placed within a given distance of a WUI block in a given year. As an alternative, we measure the percentage of public lands that were treated within some distance of a WUI block in each year.

Data on the occurrence of fires are drawn from the USGS Monitoring Trends in Burn Severity (MTBS) project, which uses satellite remote sensing data to map all fires larger than 1,000 acres occurring within the U.S. While the MTBS data do not include all fires within the period, they include the largest and therefore likely the most salient wildfires. We measure the distance from

[6]The states comprise US Forest Service regions 1-6. They are Arizona, California, Colorado, Idaho, Kansas, Montana, Nebraska, Nevada, New Mexico, North Dakota, Oregon, South Dakota, Utah, Washington, and Wyoming.

each Census block to the nearest wildfire in each year for the period 2000-2011, and we define the indicator $recentfire_{it}$ as equal to one for blocks that have experienced a wildfire within some threshold distance in the past 3 years.

Columns 2-4 of Table 1 report fuels project rates after the sample has been limited to those blocks for which a fire has occurred within 2, 5, and 10 kilometers, respectively. Comparing fuels projects rates in these columns with column 1 of Table 1, which reports the percent of the sample overall receiving treatments within a given radius, blocks are more likely to receive a fuels project when they have experienced a recent nearby fire. However, while the pattern of fuels project rates observed in Table 1 is consistent with the hypothesis that wildfire risk salience increases the demand for fuels reduction projects, it is possible that this pattern is driven by unobserved block characteristics that are correlated with both fire occurrence and fuel reduction project decisions. In section 5, we present results from estimates of equation 7 which use block and county-by-year fixed effects to control for such unobserved factors.

Finally, we collected a series of variables describing each block's demographic characteristics: income, education, age, rental rates, and race and ethnicity variables measured at the Census tract level and population density measured at the Census block level (US Census Bureau, 2000). Since our fuels treatment data span the years 2003-2011, we use demographic variables from the 2000 Census, and therefore our demographic variables are time-invariant.[7] Columns 1 and 2 of Table 2 reports the means and standard deviations of demographic variables within our sample of WUI blocks. As noted above, we standardize each demographic variable prior to estimating the regression models.

Columns 3-5 of Table 2 report means of demographic variables within block-years receiving fuel reduction projects within 2, 5, and 10 kilometers, respectively. Since demographic variables are not

[7]Note that we interact these demographic variables with dummy variables for whether the area has experienced a recent fire that acts as a focusing event. Thus, we can estimate coefficients on the interaction even though the covariates are not time-varying.

observed as time-varying, means for demographic variables are means of demographic characteristics for blocks that ever received fuels projects within a certain distance (e.g., two kilometers), weighted by how frequently they received fuels projects. Blocks for which fuels projects occur more frequently nearby tend to be wealthier, more educated, and have a higher percentage of white residents. Most significantly, when fuels projects occur within 2 km of WUI blocks, these blocks are 87 percent white, while blocks within the sample overall are 78 percent white.

Patterns in demographic variables are largely consistent with hypothesis 2. Wealthier, whiter, and more educated Census blocks are more likely to receive fuels projects. However, these patterns in and of themselves should not be interpreted as evidence that managers are more responsive to such communities. For example, these patterns could also emerge due to amenity-driven sorting. White, wealthier, and more educated individuals may be more likely to live in high amenity, high fire risk areas, and these areas are likely to be chosen as the location for fuels reduction projects. To identify differences in responsiveness to demographics, we make use of the occurrence of fires, which after accounting for fixed differences in wildfire risk provide a plausibly exogenous shock to public demand for fuels projects.

5 Results

We are interested in how bureaucratic responsiveness varies across different types of communities in ways that result in unequal representation. In particular, we are interested in how responsiveness varies with community income level, racial composition, and educational attainment. Within our sample of WUI communities these variables are highly correlated.[8] The strong correlations among

[8]Figure A.1, included the Supplementary Information, illustrates joint distributions for demographic variables within the sample of WUI blocks. The lower left panel, for example, indicates that blocks are most likely to have a very high percentage of white residents and per capita income of approximately $20,000. Further, it shows that very few blocks are observed to have a low percentage of white residents but a high per capita income. Similarly we do not observe blocks with high per capita income but low levels of educational attainment, or blocks with a high educational attainment but a low percentage of white residents.

demographic variables within our sample make it difficult to separately identify which variables bear primary responsibility for any differences in responsiveness. Therefore, in Tables 3 and 4, we test how responses vary with individual demographic characteristics.

Table 3 provides estimates of equation 7, where the dependent variable is an indicator for whether any public lands within 2 km of WUI block i received fuel treatments in year t.[9] The coefficient reported in column 1 indicates that the probability a fuels project is placed within 2 km is approximately 1.6 percentage points higher for blocks that have experienced a wildfire within 2 km in the past three years, though the difference is not statistically significant. Table 1 indicates that approximately 10 percent of blocks receive projects within 2 kilometers in a given year; therefore, recent fires cause an approximately 16 percent increase in the probability a fuels project will be placed nearby.[10]

Characteristics associated with high socioeconomic status each significantly increase the probability of fuels projects following recent fires. Table 3 indicates, for example, that when percentage above the poverty line and percentage of the population with a college education or more increase by one standard deviation, the likelihood of receiving a fuels project increases by 4.4 and 3.1 percentage points, respectively, following a nearby fire. Since the overall average rate at which communities receive fuels projects is 0.10, these coefficients represent 40% and 30% increases from baseline, respectively.

In addition to demographic characteristics, we control for differences in responsiveness by pop-

[9]We specify a linear probability model rather than a logit or probit model because it facilitates the inclusion of a large number of fixed effects and the use of cluster-robust standard errors.

[10]This result is similar, though somewhat smaller in magnitude, to the result reported in Wibbenmeyer et al. (2019). Several factors explain the discrepancy between the finding here and the finding reported in Wibbenmeyer et al. (2019). First, that paper's analysis was performed using public land grid cells as the unit of analysis, whereas here, consistent with our interest in the role of community characteristics, we treat communities as the unit of analysis. Second, to simplify the analysis we regress outcomes on an indicator for whether communities received a fuel project in any of the previous three years; Wibbenmeyer et al. (2019) examined lagged effects of recent fires separately. Because the effects observed in this analysis are smaller in magnitude, we focus on effects within 2 km thresholds, where observed effects are relatively larger and we are better able to tease out heterogeneity across characteristics associated with socioeconomic status.

ulation and housing ownership rates. Population and ownership rates are not direct measures of socioeconomic status, yet they may nonetheless affect responsiveness. Areas with higher population may face higher costs of lobbying, or, alternatively, they may be able to impose higher costs on agencies that do not meet their demands. Areas with a high proportion of renters tend to be relatively lower income, but property owners in these areas may be wealthier, and property ownership may be more concentrated. This may result in higher political efficacy due to decreased costs of lobbying from easier coordination or increased ability to impose costs on agencies. In general, we find that population density and ownership percentage are correlated negatively with post-fire responsiveness, though statistical significance varies throughout our results.

Table 3 captures differences in the extent to which blocks are treated, but may underestimate differences in responsiveness if managers are not only more likely to implement projects but are also more likely to implement larger projects around certain types of blocks. In Table 4 we use as the dependent variable the percentage of public lands within 2 kilometers on which fuels projects are implemented, and we report results from the same set of regressions as in Table 3. Overall, results are comparable. Occurrence of fire near a WUI block with average SES characteristics increases the percentage area receiving projects by about 0.6 percentage points, though the increase is not statistically significant. Higher SES communities are more likely receive fuels projects following nearby wildfires; one standard deviation increases in per capita income, percent above poverty line, percent with a college education or greater, and percent white non-Hispanic each result in an approximately 1 percent increase in the percentage of public lands near (within 2 km) a community receiving a fuels project. On average, approximately 3 percent of public land within 2 km of a community receives a fuel project in a given year; therefore, among higher SES communities, nearby fires result in an additional 33 percent increase in the percent of public land receiving fuels projects.

A disadvantage of the results presented in Tables 3-4 is that because demographic characteristics are highly correlated, it is not possible to know whether differences in responsiveness are due, for example, to differences in education or differences in racial composition. In Table 5, we include each of the demographic interactions together in the same regression. Table 5 also varies the threshold distances for fires and for fuels projects. These models provide insights into which of these demographic variables is most responsible for driving differences in responsiveness across locations. In regressions with 2 km thresholds for nearby fires and fuel projects, percentage above poverty line and percentage with a college education or more appear to be driving differences in responsiveness. In regressions with 5 km thresholds, percentage white non-Hispanic appears to have the strongest effect. However, because demographic variables are highly correlated with one another, interaction coefficients within this model are not estimated with great precision.

As a final check of combined effects of SES variables on responsiveness, we generate two indices of SES and test the differential responsiveness across communities with differing SES index levels. First, we create an SES summary index by adding the standardized SES variables and standardizing the result. Table 6 indicates that fires within 2 km increase the probability of fuels projects within 2 km for high SES communities. Second, we use principal components factor analysis to reduce our SES variables to a single factor that explains 60 percent of the total variance in the individual SES variables. Table 7 shows results very similar to those in Table 6. Where the SES factor variable is one standard deviation above the average, the post-fire (within 2 km) increase in probability of receiving a fuels project within 2 km is 3.9 percentage points greater than in average communities.

Table 8 reports results from a placebo test included to eliminate the possibility that fires and fuels reduction projects are jointly determined (perhaps as a function of fire risk that is not captured in our block and county-year fixed effects). We estimate a version of equation 7 with an indicator for whether a fire occurred over the *next* three years, as we would not expect the likelihood of observing a fuels reduction project today to be influenced by the occurrence of future fires. Significant lead effects could be the result of omitted time-varying block-level factors that are correlated with wildfires and fuels reduction projects. The estimated coefficients on the main effects and on the interaction terms are small relative to the estimates using fires from the past three years and not significantly different from zero with only a few exceptions.

6 Discussion

In this paper, we find evidence for unequal responsiveness on the part of the bureaucracy to higher SES communities. In particular, we find consistent evidence that high SES communities receive more local public goods when they experience a shock in demand. Forest managers are more likely to implement fuels projects near communities that have recently experienced a fire when those communities have a higher percentage of high income, high education, and white residents. This finding is consistent with previous work on inequality in government responsiveness (Gilens, 2005; McCarty et al., 2009; Gilens, 2011). Unlike these earlier studies, which focus on responsiveness among elected officials, we focus on inequality in responsiveness of public agencies. Further, our study is one of few to provide quasi-experimental evidence regarding inequality in government responsiveness across communities. This paper shows that similar events can yield very different policy outcomes for different types of communities, a finding that should spawn further research on the mechanisms by which inequity is perpetuated.

The formal model indicates that such unequal responsivness can be due to differences in the cost of lobbying or the costs that communities can impose on agencies by mobilizing. One limitation of this analysis is that because we have no direct measure of citizen political engagement, we cannot discern from the main analysis whether differences across communities in fuels project rates after fires are due to differences in salience-motivated political action or to bias on the part of bureaucrats (Einstein and Glick, 2017). To test whether demographic factors are associated with more mobilization on fire-related programs, we conduct an additional analysis. We regress the number of Firewise USA® communities per 1000 residents in California counties on demographics (see Supplementary Information). Becoming a Firewise USA® community requires forming a committee of residents and other stakeholders, assessing wildfire risk, and developing and undertaking a plan to reduce wildfire risk. Results across the 58 California counties show that counties with more white, non-Hispanic residents have more Firewise USA® communities per 1000 residents, as do counties with higher fire risk. The other demographic indicators are not associated with more Firewise USA® communities and a summary variable that collapses all socioeconomic characteristics is also not associated with more community participation. These supplementary analyses provide suggestive evidence that communities with more white, non-Hispanic residents are better able to mobilize, indicating that the additional resources provided by government agencies may be at least partially a function of mobilization, not just biased bureaucrats.

Future research should continue to explore this differential responsiveness of government agencies, including addressing sources of mobilization. That we find evidence of inequality of responsiveness in fire management, which is far removed from the social policies, especially education, where much of the research has been focused (e.g. Meier et al., 1999; Dee, 2004), indicates both the possible ubiquitousness of differential responsiveness and the need for further research in diverse policy areas. More research on patterns of responsiveness should be paired with a deeper qualitative understanding of community mobilization and how organizations play a role in that mobilization (Han, 2009). An improved understanding of community mobilization may provide insights regarding strategies lower SES communities can use to achieve the political efficacy demonstrated by more privileged communities (Verba et al., 1995).

References

Abatzoglou, J. T. and A. P. Williams (2016). Impact of anthropogenic climate change on wildfire across western us forests. *Proceedings of the National Academy of Sciences 113*(42), 11770–11775.

Aberbach, J. D. (1990, Jun). *Keeping a Watchful Eye: The Politics of Congressional Oversight*. Brookings Institution Press.

Anderson, S. E., H. E. Hodges, and T. L. Anderson (2013). Technical management in an age of openness: The political, public, and environmental forest ranger. *Journal of Policy Analysis and Management 32*(3), 554–573.

Anderson, S. E. and M. Potoski (2016). Agency structure and the distribution of federal spending. *Journal of Public Administration Research and Theory 26*(3), 461–474.

Arno, S., J. Scott, and G. Hartwell (1995). Age-class structure of old growth ponderosa pine/douglas-fir stands and its relationship to fire history, int-rp-481.

Bartels, L. M. (2008). *Unequal democracy: The political economy of the new gilded age*. Princeton, New Jersey: Princeton University Press.

Bawn, K. (1995). Political control versus expertise: Congressional choices about administrative procedures. *American Political Science Review 89*(1), 62–73.

Bin, O. and C. E. Landry (2013). Changes in implicit flood risk premiums: Empirical evidence from the housing market. *Journal of Environmental Economics and Management 65*(3), 361–376.

Birkland, T. A. (1997, September). *After Disaster: Agenda Setting, Public Policy, and Focusing Events*. Washington, DC: Georgetown University Press.

Bostwick, P., J. Menakis, and T. Sexton (2011). How fuel treatments saved homes from the 2011 wallow fire. US Forest Service Fuel Treatment Effectiveness Assessement. Available online: `https://www.fs.usda.gov/Internet/FSE_DOCUMENTS/stelprdb5318765.pdf` [last accessed May 15, 2015].

Bowler, S. and G. Segura (2011). *The future is ours: Minority politics, political behavior, and the multiracial era of American politics*. CQ Press.

Butler, D. M. and D. E. Broockman (2011). Do politicians racially discriminate against constituents? a field experiment on state legislators. *American Journal of Political Science 55*(3), 463–477.

Dee, T. S. (2004). Teachers, race, and student achievement in a randomized experiment. *The Review of Economics and Statistics 86*(1), 195–210.

Dennison, P. E., S. C. Brewer, J. D. Arnold, and M. A. Moritz (2014). Large wildfire trends in the western united states, 1984–2011. *Geophysical Research Letters 41*(8), 2928–2933.

Dion, C., P. Lanoie, and B. Laplante (1998). Monitoring of pollution regulation: Do local conditions matter? *Journal of Regulatory Economics 13*(1), 5–18.

Einstein, K. L. and D. M. Glick (2017). Does race affect access to government services? an experiment exploring street-level bureaucrats and access to public housing. *American Journal of Political Science 61*(1), 100–116.

Epstein, D. and S. O'Halloran (1994, Aug). Administrative procedures, information, and agency discretion. *38*, 697–722.

Epstein, D. and S. O'Halloran (1994). Administrative procedures, information, and agency discretion. *American Journal of Political Science 38*(3), 697–722.

Erikson, R. S. (2015). Income inequality and policy responsiveness. *Annual Review of Political Science 18*, 11–29.

Franko, W. W. (2013). Political inequality and state policy adoption: Predatory lending, children's health care, and minimum wage. *Poverty & Public Policy 5*(1), 88–114.

GAO (2019). Wildland fire: Federal agencies' efforts to reduce wildland fuels and lower risk to communities and ecosystems. Technical Report GAO-20-52, Government Accountability Office.

Gilens, M. (2005). Inequality and democratic responsiveness. *Public Opinion Quarterly 69*(5), 778–796.

Gilens, M. (2011). Policy consequences of representational inequality. In P. K. Enns and C. Wlezien (Eds.), *Who Gets Represented*, pp. 247–86. New York, NY: Russell Sage Foundation.

Griffin, J., Z. Hajnal, B. Newman, and D. Searle (2019). Political inequality in america: Who loses on spending policy? when is policy less biased? *Politics, Groups, and Identities 7*(2), 367–385.

Griffin, J. D. and B. Newman (2008). *Minority report: Evaluating political equality in America.* University of Chicago Press.

Grimes, M. and P. Esaiasson (2014). Government responsiveness: A democratic value with negative externalities? *Political Research Quarterly 67*(4), 758–768.

Hakanson, E. K. (2010). Fuels management policy and practice in the US Forest Service. Ph.D. Thesis, The University of Montana.

Han, H. (2009). *Moved to action: Motivation, participation, and inequality in American politics.* Stanford University Press.

Huber, J. D. and C. R. Shipan (2002, Sep). *Deliberate Discretion? The Institutional Foundations of Bureaucratic Autonomy.* Cambridge University Press.

Johnson, R. N. and M. J. Watts (1989). Contractual stipulations, resource use, and interest groups: implications from federal grazing contracts. *Journal of Environmental Economics and Management 16*(1), 87–96.

Kahneman, D. (2003). A perspective on judgment and choice: Mapping bounded rationality. *American Psychologist 58*(9), 697–720.

Kalies, E. L. and L. L. Y. Kent (2016). Tamm review: Are fuel treatments effective at achieving ecological and social objectives? A systematic review. *Forest Ecology and Management 375*, 84–95.

Keane, R. E., K. C. Ryan, T. T. Veblen, C. D. Allen, J. A. Logan, B. Hawkes, and J. Barron (2002). The cascading effects of fire exclusion in rocky mountain ecosystems. In J. S. Baron (Ed.), *Rocky Mountain futures: An ecological perspective*, pp. 133–152. Washington, DC: Island Press.

Kennedy, M. C. and M. C. Johnson (2014). Fuel treatment prescriptions alter spatial patterns of fire severity around the wildland–urban interface during the Wallow Fire, Arizona, USA. *Forest Ecology and Management 318*, 122–132.

Konisky, D. M. (2009). Inequities in enforcement? environmental justice and government performance. *Journal of Policy Analysis and Management 28*(1), 102–121.

Konisky, D. M. and C. Reenock (2018). Regulatory enforcement, riskscapes, and environmental justice. *Policy Studies Journal 46*(1), 7–36.

McCarty, N., K. T. Poole, and H. Rosenthal (2009). *Polarized America: The dance of ideology and unequal riches*. Cambridge, Massachusetts: MIT Press.

McCoy, S. J. and R. P. Walsh (2014). WUI on fire: Risk, salience & housing demand. NBER Working Paper No. 20644.

McCubbins, M. D., R. G. Noll, and B. R. Weingast (1987). Administrative procedures as instruments of political control. *Journal of Law, Economics, & Organization 3*(2), 243–277.

Meier, K. J., R. D. Wrinkle, and J. L. Polinard (1999). Representative bureaucracy and distributional equity: Addressing the hard question. *The Journal of Politics 61*(4), 1025–1039.

Naficy, C., A. Sala, E. G. Keeling, J. Graham, and T. H. DeLuca (2010). Interactive effects of historical logging and fire exclusion on ponderosa pine forest structure in the northern Rockies. *Ecological Applications 20*(7).

Nie, N. H., J. Junn, K. Stehlik-Barry, et al. (1996). *Education and democratic citizenship in America*. University of Chicago Press.

Ojeda, C. (2018). The two income-participation gaps. *American Journal of Political Science 62*(4), 813–829.

Page, B. I., L. M. Bartels, and J. Seawright (2013). Democracy and the policy preferences of wealthy Americans. *Perspectives on Politics 11*(1), 51–73.

Ringquist, E., J. Worsham, and M. Eisner (2003). Salience, complexity, and the legislative direction of regulatory bureaucracies. *Journal of Public Administration Research and Theory 13*(2), 141–164.

Ringquist, E. J. (2005). Assessing evidence of environmental inequities: A meta-analysis. *Journal of Policy Analysis and Management 24*(2), 223–247.

Rosenstone, S. J. and J. M. Hansen (1993). *Mobilization, participation, and democracy in America*. Longman Publishing Group.

Sabatier, P. A., J. Loomis, and C. McCarthy (1995). Hierarchical controls, professional norms, local constituencies, and budget maximization: An analysis of US Forest Service planning decisions. *American Journal of Political Science*, 204–242.

Sances, M. W. (2016). The distributional impact of greater responsiveness: Evidence from New York towns. *The Journal of Politics 78*(1), 105–119.

Schoennagel, T., T. Veblen, and W. Romme (2004). The interaction of fire, fuels, and climate across Rocky Mountain forests. *BioScience 54*(7), 661–676.

Scholz, J. T. and C.-L. Wang (2006). Cooptation or transformation? local policy networks and federal regulatory enforcement. *American Journal of Political Science 50*(1), 81–97.

Stetler, K. M., T. J. Venn, and D. E. Calkin (2010). The effects of wildfire and environmental amenities on property values in northwest Montana, USA. *Ecological Economics 69*, 2233–2243.

Stokes, L. C. (2016). Electoral backlash against climate policy: A natural experiment on retrospective voting and local resistance to public policy. *American Journal of Political Science 60*(4), 958–974.

Taylor, S. E. and S. C. Thompson (1982). Stalking the elusive 'vividness' effect. *Psychological Review 89*(2), 155–181.

Tversky, A. and D. Kahneman (1973). Availability: A heuristic for judging frequency and probability. *Cognitive Psychology 5*(2), 207–232.

Ura, J. D. and C. R. Ellis (2008). Income, preferences, and the dynamics of policy responsiveness. *PS: Political Science & Politics 41*(4), 785–794.

US Census Bureau (2000). Census 2000 summary file 3 sample data. Available online: `https://factfinder.census.gov/` [accessed March 26, 2018].

Verba, S., K. L. Schlozman, and H. E. Brady (1995). *Voice and equality: Civic voluntarism in American politics*, Volume 4. Cambridge Univ Press.

Westerling, A., H. Hidalgo, D. Cayan, and T. Swetnam (2006). Warming and earlier spring increases western U.S. forest wildfire activity. *Science 313*(5789), 940–943.

Westerling, A. L. (2016). Increasing western us forest wildfire activity: Sensitivity to changes in the timing of spring. *Phil. Trans. R. Soc. B 371*(1696), 20150178.

White, A. R., N. L. Nathan, and J. K. Faller (2015). What do i need to vote? bureaucratic discretion and discrimination by local election officials. *American Political Science Review 109*(1), 129–142.

Wibbenmeyer, M., S. E. Anderson, and A. J. Plantinga (2019). Salience and the government provision of public goods. *Economic Inquiry 57*(3), 1547–1567.

Wlezien, C. and S. N. Soroka (2011). Inequality in policy responsiveness? In P. K. Enns and C. Wlezien (Eds.), *Who Gets Represented*, pp. 285–310. New York, New York: Russell Sage Foundation.

Wood, B. D. and J. Bohte (2004). Political transaction costs and the politics of administrative design. *The Journal of Politics 66*(1), 176–202.

Table 1: Rate at which WUI blocks received fuel treatments within a given radius, for full sample and for WUI block-years in which a recent nearby fire has occurred

Radius	Full sample	Block-years with recent fire within distance 2 km	5 km	10 km
2 kilometers	0.10	0.16	0.14	0.13
	117,160	9,761	18,905	32,555
5 kilometers	0.16	0.23	0.21	0.19
	350,475	22,755	49,548	91,855
10 kilometers	0.26	0.39	0.35	0.32
	786,863	47,798	103,991	200,181
Number of WUI blocks	364,689	41,145	90,459	157,463
Number of block-year obs.	3026800	122,219	301,332	629,513

Table 2: Demographic and political characteristics for the entire sample of WUI blocks, and for WUI blocks receiving nearby fuel reduction projects

	Full sample		Block-years with fuels projects within distance		
			2 km	5 km	10 km
Population	47.3	[98.0]	34.3	38.8	43.4
Per cap. income	20805.0	[10198.8]	22460.3	21933.2	21600.8
Pct. above pov. line	0.87	[0.097]	0.88	0.88	0.88
Pct. high school grad.	0.84	[0.11]	0.87	0.87	0.86
Pct. college or greater	0.23	[0.15]	0.26	0.26	0.25
Pct. white non-Hispanic	0.78	[0.22]	0.87	0.85	0.83
Pct. Hispanic	0.14	[0.17]	0.080	0.089	0.10
Pct. own place of residence	0.74	[0.16]	0.75	0.73	0.73
Number of WUI blocks	302,680		9,791	21,266	28,218
Number of block-year obs.	3,026,800		117,160	350,475	786,863

Table 3: Linear probability models of the effect of fire within 2 km of a WUI community on the probability the community will receive a fuels project within 2 km over the next 3 years, for communities varying by demographic characteristics.

	(1)	(2)	(3)	(4)	(5)
Fire within 2 km	.016	.015	.013	.016	.016
	[.0098]	[.01]	[.0099]	[.0099]	[.01]
SES and recent fire ints.					
Per cap. income		.019**			
		[.0068]			
Pct. above pov. line			.044**		
			[.0088]		
Pct. college or greater				.031**	
				[.0087]	
Pct. white non-Hispanic					.028**
					[.011]
Control and recent fire ints.					
Population		-.0039	-.0054*	-.005	-.0014
		[.0025]	[.0025]	[.0027]	[.0019]
Pct. own place of residence		-.021*	-.032**	-.023*	-.019*
		[.0097]	[.0096]	[.0096]	[.0089]
Distance to fuel project	2	2	2	2	2
Distance to fire	2	2	2	2	2
No. WUI blocks	108,224	108,216	108,216	108,212	108,221
No. county-years	3,970	3,970	3,970	3,970	3,970
No. obs.	1,082,240	1,082,160	1,082,160	1,082,120	1,082,210

Note: All columns include WUI community fixed effects and county-by-year fixed effects. Robust standard errors are clustered by Census tract, ** p<0.01, * p<0.05, +p< 0.10.

Table 4: Effect of fire within 2 km of a WUI community on percent of public land within 2 km receiving fuels projects over the next 3 years, for communities varying by demographic characteristics.

	(1)	(2)	(3)	(4)	(5)
Fire within 2 km	.006	.0056	.005	.0061	.0061
	[.0047]	[.005]	[.0047]	[.0048]	[.005]
SES and recent fire ints.					
Per cap. income		.0089**			
		[.0033]			
Pct. above pov. line			.017**		
			[.0044]		
Pct. college or greater				.013**	
				[.0042]	
Pct. white non-Hispanic					.0093
					[.0051]
Control and recent fire ints.					
Population		-.0012	-.0017	-.0016	-.00025
		[.00098]	[.001]	[.0011]	[.00081]
Pct. own place of residence		-.0086	-.012*	-.0092	-.0069
		[.005]	[.0052]	[.005]	[.0045]
Distance to fuel project	2	2	2	2	2
Distance to fire	2	2	2	2	2
No. WUI blocks	108,224	108,216	108,216	108,212	108,221
No. county-years	3,970	3,970	3,970	3,970	3,970
No. obs.	1,082,240	1,082,160	1,082,160	1,082,120	1,082,210

Note: All columns include WUI community fixed effects and county-by-year fixed effects. Robust standard errors are clustered by Census tract, ** p<0.01, * p<0.05, +p< 0.10.

Table 5: Effect of fire on fuels projects conditional on demographics, varying dependent variables (an indicator for nearby projects and the percent of nearby public land receiving projects) and varying threshold distances (distance to fuel project= $\{2, 5\}$ and distance to fire (d)= $\{2, 5\}$).

	(1) Any projects nearby	(2) Any projects nearby	(3) Pct. pub. land treated	(4) Pct. pub. land treated
Fire nearby	.013	.0069	.0048	.0044**
	[.0098]	[.0079]	[.0048]	[.0016]
SES and recent fire ints.				
Per cap. income	-.011	-.012	-.0021	-.0029
	[.0085]	[.0077]	[.0039]	[.0016]
Pct. above pov. line	.03**	.016*	.011*	.0027
	[.0096]	[.0069]	[.0044]	[.0015]
Pct. college or greater	.027*	-.0017	.01*	.00092
	[.011]	[.0093]	[.0051]	[.002]
Pct. white non-Hispanic	.019	.03**	.0053	.0055**
	[.01]	[.0078]	[.0049]	[.0016]
Control and recent fire ints.				
Population	-.005*	-.0034	-.0018	-.0002
	[.0024]	[.0017]	[.0011]	[.00037]
Pct. own place of residence	-.034**	-.02**	-.013*	-.0047**
	[.0099]	[.0061]	[.0053]	[.0016]
Distance to fuel project	2	5	2	5
Distance to fire	2	5	2	5
No. WUI blocks	108,209	213,372	108,209	213,372
No. county-years	3,970	4,250	3,970	4,250
No. obs.	1,082,090	2,133,720	1,082,090	2,133,720

Note: All columns include WUI community fixed effects and county-by-year fixed effects. Robust standard errors are clustered by Census tract, ** $p<0.01$, * $p<0.05$, +$p< 0.10$.

Table 6: Differential responsiveness by SES, with SES measured by a summary index

	(1) Any projects nearby	(2) Any projects nearby	(3) Pct. pub. land treated	(4) Pct. pub. land treated
Fire nearby	.013	.0069	.0046	.0044**
	[.0097]	[.0078]	[.0048]	[.0016]
SES summation index	.041**	.014	.017**	.0029
	[.0099]	[.0079]	[.0048]	[.0018]
Control and recent fire ints.				
Population	-.0047	-.0061**	-.0015	-.00067
	[.0025]	[.0021]	[.001]	[.0004]
Pct. own place of residence	-.031**	-.015**	-.012*	-.004**
	[.01]	[.0058]	[.0054]	[.0015]
Distance to fuel project	2	5	2	5
Distance to fire	2	5	2	5
No. WUI blocks	108,209	213,372	108,209	213,372
No. county-years	3,970	4,250	3,970	4,250
No. obs.	1,082,090	2,133,720	1,082,090	2,133,720

Note: The SES summary index was constructed by summing the standardized SES variables, and standardizing the result. All columns include WUI community fixed effects and county-by-year fixed effects. Robust standard errors are clustered by Census tract, ** p<0.01, * p<0.05, +p< 0.10.

Table 7: Differential responsiveness by socioeconomic status (SES)

	(1) Any projects nearby	(2) Any projects nearby	(3) Pct. pub. land treated	(4) Pct. pub. land treated
Fire nearby	.013	.007	.0047	.0044**
	[.0097]	[.0078]	[.0048]	[.0016]
SES factor variable	.039**	.011	.016**	.0024
	[.0095]	[.0078]	[.0047]	[.0018]
Control and recent fire ints.				
Population	-.0049	-.0061**	-.0016	-.00067
	[.0026]	[.0021]	[.001]	[.00041]
Pct. own place of residence	-.03**	-.014*	-.012*	-.0038*
	[.01]	[.0058]	[.0054]	[.0015]
Distance to fuel project	2	5	2	5
Distance to fire	2	5	2	5
No. WUI blocks	108,212	213,378	108,212	213,378
No. county-years	3,970	4,250	3,970	4,250
No. obs.	1,082,120	2,133,780	1,082,120	2,133,780

Note: The SES factor variable is constructed using principal-components factoring and explains 60.6% of total variance within the SES variables. All columns include WUI community fixed effects and county-by-year fixed effects. Robust standard errors are clustered by Census tract, ** p<0.01, * p<0.05, +p< 0.10.

Table 8: Placebo test results from regressions of nearby fuels projects on occurrence of nearby wildfires over the following three years.

	(1) Any projects nearby	(2) Any projects nearby	(3) Pct. pub. land treated	(4) Pct. pub. land treated
Fire nearby	-.004	.014	-.0019	.0013
	[.013]	[.0092]	[.0065]	[.0023]
SES and recent fire ints.				
Per cap. income	-.02*	-.021*	-.0038	-.0012
	[.01]	[.0096]	[.0053]	[.0022]
Pct. above pov. line	.0065	.015	-.0067	.00057
	[.014]	[.0092]	[.0068]	[.002]
Pct. college or greater	.027	.018	.0095	-.0004
	[.016]	[.01]	[.0082]	[.0025]
Pct. white non-Hispanic	.013	-.0065	.003	.00011
	[.0092]	[.0093]	[.0038]	[.0017]
Control and recent fire ints.				
Population	-.0023	-.0052**	-.00067	-.00048
	[.002]	[.0019]	[.00088]	[.00045]
Pct. own place of residence	-.011	-.011	.0013	.00035
	[.0098]	[.0081]	[.0051]	[.002]
Distance to fuel project	2	5	2	5
Distance to fire	2	5	2	5
No. WUI blocks	108,209	213,372	108,209	213,372
No. county-years	3,970	4,250	3,970	4,250
No. obs.	1,082,090	2,133,720	1,082,090	2,133,720

Note: All columns include WUI community fixed effects and county-by-year fixed effects. Robust standard errors are clustered by Census tract, ** $p<0.01$, * $p<0.05$, +$p< 0.10$.

A Supplementary Information

The National Firewise Communities Program is a national interagency program that encourages communities to develop and implement local solutions for wildfire preparedness. Participating agencies include the U.S. Forest Service, the Bureau of Land Management, and other federal land management agencies. To be recognized as a Firewise USA® community, a community must create and implement a local plan with cooperative assistance from state forestry agencies and local fire staff.

We collected data on the number of Firewise USA® communities in each county in California, along with data from the U.S. Census Bureau on socioeconomic status: per capita income, percent of residents above the poverty line, percent of residents with a college education or higher, and the percent of residents who are white non-Hispanic. Additional county-level data was assembled on population, percent of residents who own their residence, and the percent of the county with vegetation that has a high degree of departure from historical conditions, classified as Vegetation Class 3 by the Landfire project (see `https://www.landfire.gov/about.php-#planning` for details). In California, Class 3 vegetation is associated with high fire risk.

As reported in Table A.1, we regressed the number of Firewise USA® communities per 1000 residents on the socioeconomic status variables (model 1) and on the socioeconomic status variable plus controls (model 2). Percent white non-Hispanic and Vegetation Class 3 are found to have a positive and significant ($p < 0.05$) effects on the number of Firewise USA® communities. In the average county, a one percentage point increase in white non-Hispanic residents increases the number of Firewise USA® communities by about 1.67, a 44% increase. We also created indices for the four socioeconomic status variables (models 3 and 4), but found they had insignificant effects on the number of Firewise USA® communities.

Table A.1: The effects of socioeconomnic status and vegetation conditions on the number of Firewise USA® communities in California counties

Socioeconomic status	(1)	(2)	(3)	(4)
Per cap. income	.008	.012		
	[.009]	[.009]		
Pct. above pov. line	-.765	-1.141		
	[.772]	[.79872]		
Pct. college or greater	-.375	-.493		
	[.601]	[.617]		
Pct. white non-Hispanic	.365**	.403*		
	[.118]	[.161]		
SES factor variable			.007	
			[.023]	
SES summation				.018
				[.023]
Controls				
Pct. vegetation class 3		1.307**	1.010*	1.038*
		[.472]	[.488]	[.487]
Population		.007	-.007	-.007
		[.017]	[.017]	[.017]
Pct. own place of residence		.142	.596+	.552+
		[.414]	[.298]	[.303]
Constant	.369	.442	-.347+	-.321
	[.532]	[.510]	[.191]	[.193]
No. obs.	58	58	58	58

Note: The dependent variable is the number of Firewise USA® communities per 1000 residents in a county. Per capita income is measured in thousands of dollars and Population is measured in millions of people. Pct. vegetation class 3 measures the percentage of a county covered by vegetation that has a high degree of departure from historical vegetation conditions. **$p<0.001$, *$p<0.05$, +$p<0.10$.

Figure A.1: Joint distribution of demographic variables within the sample of WUI blocks. Values above or below or the 97.5 or 2.5 percentiles, respectively, for any demographic variable have replaced with the 97.5 or 2.5 percentile value.