Treatment versus Punishment: Understanding Racial Inequalities in Drug Policy

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Abstract

Context: Many observers believe that the policy response to the opioid crisis is less punitive than the crack scare and that the reason is that victims are (stereotypically) white.

Methods: To assess this conjecture, we compile new longitudinal data on district-level drug-related deaths and (co)sponsorship of legislation on drug abuse in the House of Representatives over the past four decades. Using legislator fixed effects models, we then test how changes in drug-related death rates in legislators' districts predict changes in (co)sponsorship of treatment-oriented or punitive legislation in the subsequent year and assess whether these relationships vary by race of victim or drug type.

Findings: Policy makers were more likely to introduce punitive drug-related bills during the crack scare and are more likely to introduce treatment-oriented bills during the current opioid crisis. The relationship between district-level drug deaths and subsequent sponsorship of treatment-oriented legislation is greater for opioid deaths than for cocaine-related deaths and for white victims than for black victims. By contrast, district-level drug deaths are not significantly related to sponsorship of punishment-oriented bills.

Conclusions: These results suggest that the racial inequalities and double standards of drug policy still persist but in different forms.

Keywords crack, opioids, policy

The opioid crisis continues to reach new levels of severity but seemingly receives disproportionately less public attention, media coverage, and legislative action than crack cocaine did in the 1980s and 1990s. The discrepancy in responses between these two cases is not easily explained by the objective severity of the crises. More than 70,000 Americans died of drug

Journal of Health Politics, Policy and Law, Vol. 45, No. 2, April 2020
DOI 10.1215/03616878-8004850

overdoses in 2017, a record total that far exceeds the number who died from car accidents or gun violence (Katz and Sanger-Katz 2018). Most of these deaths—47,600 (68%)—involved opioids (Scholl et al. 2018). Our analysis of the Center for Disease Control and Prevention's mortality data indicates the number of opioid-related deaths in 2016 alone (41,518) exceeds the number of cocaine-related deaths in the 1980s and 1990s combined (38,371). The scale of the opioid crisis thus outstrips any prior US drug epidemic. In addition, though the use of crack cocaine was associated with negative social and public health consequences such as increased homicides (Fryer et al. 2013; Golub and Johnson 1997), the opioid crisis has had massive social costs and has generated substantial negative externalities as well (see, e.g., Kolhatkar 2017).

As observers frequently note (e.g., Cohen 2015; Peterson and Armour 2018), the federal policy response to the opioid crisis seemingly emphasizes treatment and rehabilitation to a greater extent than the punitive approach that dominated drug policy in recent decades.[1] At the height of the crack scare, for instance, the 1992 Republican platform stated, "Drug users must face punishment, including fines and imprisonment, for contributing to the demand that makes the drug trade profitable" (Delegates to the RNC 1992). As a result of policy and administrative changes resulting from this punitive consensus, which was largely endorsed by both parties, the number of drug-related arrests and the number of people entering prison for drug crimes increased dramatically after the early 1990s (BJS 2019; Rothwell 2015). By contrast, the 2016 Republican platform highlighted how "the opioid crisis is ravaging communities all over the country, often hitting rural areas harder than urban," and called for "expeditious agreement" on a bill later signed by President Obama that sought to "expand prevention and education efforts while also promoting treatment and recovery" (CADCA 2019; Delegates to the RNC 2016). This discrepancy has frequently been noted by lawmakers and journalists, who conjecture that the shift is the result of greater empathy for stereotypically white opioid users compared to stereotypically black crack users (e.g., Glanton 2017; King 2017; Newkirk 2017).[2]

However, these conjectures about the differences between the policy response to the opioid crisis and the crack scare have not been systematically tested. In addition, little convincing evidence exists that isolates race

1. This policy difference appears to be replicated at the state level, though a comparison of drug policy across all 50 states is beyond the scope of this article. Mauer and Huling 1995 discuss changes in state approaches to drug policy during the crack scare. For recent reviews of the state policy response to the opioid crisis, see NCSL 2017 and Parker, Strunk, and Fiellin 2018.

2. Contrary to these stereotypes, the opioid crisis has claimed numerous nonwhite victims (see, e.g., Shihipar 2019).

or drug type as the key factors that explain any such differences, which could instead reflect a broader shift toward viewing drug addiction as a type of disease rather than a crime (e.g., Pew Research Center 2014).

In this article, we therefore measure the policy response to the opioid crisis in Congress and compare its content with the response to the crack scare.[3] Drawing from theory and prior research on policy responsiveness, we consider the following four research questions. First, we test whether the legislative response to the crises has differed in the aggregate, comparing the bills introduced during these epidemics and the extent to which they focus on treatment versus punishment. Second, we assess whether legislators respond to district-related drug deaths with drug policy legislation and, further, whether they respond with a treatment- or punishment-oriented approach. Third, we consider whether these patterns of responsiveness to drug deaths differ among opioids, cocaine, and methamphetamine and between white and black victims. Fourth, we test if these relationships vary over time, comparing the crack scare, the opioid crisis, and the period between them, which allows us to examine whether the recent shifts toward more empathetic approaches (if any) hold across different drug types and victims' race. Finally, we evaluate the robustness of our findings to controlling for measures of homicide deaths at the district level and test for heterogeneity in responsiveness to drug-related deaths by legislator party or factors that affect media coverage.

We evaluate these theories with newly coded data on legislative sponsorship and cosponsorship of drug-related bills in the US House of Representatives and data on drug-related deaths at the congressional district level over the past four decades. Using legislator fixed effects models, we test how changes in drug-related death rates in legislators' districts predict changes in (co)sponsorship of treatment-oriented or punitive legislation in the subsequent year and assess whether these relationships vary by race of victim or drug type. Our findings indicate that policy makers were more likely to introduce punitive drug-related bills during the crack scare and are more likely to introduce treatment-oriented bills during the current opioid crisis. We also find that legislators respond to drug deaths in their district by sponsoring more treatment-oriented legislation, but this relationship is only observed for opioid deaths and white victims. Legislators are

3. We focus on the federal legislative response to these drug epidemics given the nationwide attention paid to these crises and the lack of available data tracking state-level bills on drug policy across all 50 states. However, previous evidence suggests that state and federal drug policies tend to move in tandem (Mauer and Huling 1995; NCSL 2017; Parker et al. 2018). Evaluating the extent to which our findings hold at the state level is thus an important topic for future research.

specifically more responsive to opioid-related deaths than cocaine-related deaths (especially during the opioid crisis) and to white drug deaths than to black drug deaths. By contrast, we observe no evidence of a relationship between district-level drug deaths and punishment-oriented bills regardless of drug, race of victim, or era.

Theoretical Approach

What factors cause legislators to propose changes to drug policy? If political elites responded directly to objective conditions, legislative attention to the opioid crisis would be expected to be far greater than that of the crack scare. However, scholars have long emphasized that objective conditions are relevant but not decisive in setting the national agenda. Changes in issue salience often result instead from political entrepreneurs exploiting exogenous events or institutional processes to advance their policy goals (Adler and Wilkerson 2013; Kingdon and Thurber 1984). Compelling "focusing events" can also help to put issues on the policy agenda (Birkland 1997), which is shaped in part by episodic and often nonlinear changes in media coverage (Boydstun 2013; Weaver, McCombs, and Shaw 2004). By contrast, a lack of media coverage can reduce public and legislative attention to a problem and thereby reduce the likelihood of a policy response (Eisensee and Strömberg 2007).

Prior research shows that attention to the issue of illegal drugs is often divorced from objective measures of severity. In the case of crack cocaine, media coverage was extensive and frequently inaccurate (e.g., the panic over so-called crack babies; see Newkirk 2017). News reports hyped myths about crack cocaine that reinforced negative racial stereotypes (Golub and Johnson 1997)—part of a pattern of racialized news reporting that increased support for punitive approaches to crime, especially among people with negative racial attitudes (Dixon 2006; Gilliam Jr. and Iyengar 2000; Hurwitz and Peffley 1997). Politicians leveraged the increased salience of drug use to make a punishment-oriented approach to the issue an important public priority (Baumgartner and Jones 1993: 153–61). This tactic resonated with public opinion, which was heavily propunishment at the time (Enns 2014, 2016). By contrast, the opioid crisis did not center in urban areas among nonwhite Americans, lacked identifiable perpetrators like crack cocaine dealers, and came at a time when public demand for a punitive approach to crime had declined (Enns 2014, 2016). Politicians have therefore not exploited the issue as extensively as they exploited crack; similarly, media depictions have tended to be more sympathetic and less racialized (Dasgupta, Mandl, and Brownstein 2009; Harbin n.d.; Netherland and Hansen 2016).

As a result of these differences, attention to and interest in the crack scare greatly exceeded that of the opioid crisis despite the latter's far larger death toll. In 1989, a time when overdose deaths were a small fraction of the current total, 64% of Americans said drugs were the most important problem facing the country (CBS News/*New York Times* 1989). Only 2% said the same in December 2018 (Gallup 2019). Similarly, during the 1989–90 period, for example, 417 *New York Times* front-page stories mentioned crack compared with only 68 for opioids in 2017–18.[4] Public support for tough-on-crime policies has ebbed since its high-water mark in the early 1990s (Enns 2016). We therefore expect to observe a less intense and less punitive legislative response to the opioid crisis than to the crack scare. We test this expectation empirically by describing changes in treatment- and punishment-oriented legislation over the past four decades, drawing on comprehensive data of bills introduced in the US House of Representatives.

To better understand the factors that promote different responses to the two drugs, we specifically consider whether and how legislators respond to the severity of these drug epidemics in their districts. Previous research provides theoretical reasons to expect district-level responsiveness. In some cases, district conditions or characteristics may serve as a proxy for constituent preferences (Peltzman 1984). In other cases, legislators may anticipate future constituent preferences over outcomes (Canes-Wrone, Herron, and Shotts 2001) and assume they will be held accountable retrospectively (e.g., as occurred with local casualties in a war they supported—see Grose and Oppenheimer 2007). Finally, some legislators may simply seek to act on behalf of perceived constituent interests as a trustee model of representation would predict.

The available evidence, though limited, does suggest that legislators respond to district conditions and would thus be expected to respond to the severity of drug-related deaths in their districts. For instance, studies find a correspondence between district conditions and voting records on agriculture (Bellemare and Carnes 2015), poverty (Miler 2018), and free trade (Conconi, Facchini, and Zanardi 2012; Xie 2006). Further evidence

4. Results based on Nexis Uni searches for publication (*New York Times*) AND crack AND ("Section 1; Page 1" OR "Section A; Page 1" OR A1) for 1/1/1989–12/31/1990 and publication (*New York Times*) AND opioid AND ("Section 1; Page 1" OR "Section A; Page 1" OR A1) for 1/1/2017–12/31/2018.

suggests that legislators respond to changes in the status quo within their district. For instance, Winburn and Sullivan (2011) find that legislators from districts affected by Hurricane Katrina introduced more disaster relief bills after the storm, while Cayton (2017) finds that legislators from districts hardest hit by the Great Recession were more likely to vote to extend unemployment benefits.

These relationships are documented most systematically in legislative voting by Adler, Cayton, and Griffin (2018), who find that district conditions are related to voting in Congress even after accounting for constituent preferences. Similarly, Lazarus (2013) and Waggoner (2019) find that sponsorship of issue-specific legislation is strongly associated with employment levels in related industries. These relationships appear to be strongest in the House for electorally vulnerable members (Lazarus 2013), though it is important to note that such effects are typically strongly conditioned by party (see, e.g., Adler, Cayton, and Griffin 2018; Kriner and Shen 2014) and are not always observed (see in particular Fowler and Hall 2016).

There are reasons to doubt, however, that the likelihood or content of legislators' policy response to changing conditions in their districts will necessarily be proportional to the severity of the problem. First, the volume of coverage that various risks receive in the media, which has an important influence on legislative behavior (see, e.g., Arnold 2004), rarely correspond to objective measures of severity (see, e.g., Bomlitz and Brezis 2008; Frost, Frank, and Maibach 1997). Similarly, public concern tends to be driven more by cues from elites than by objective conditions—Beckett (1994) found, for instance, that the perceived importance of drugs and crime tracked with statements by government officials, not incidence rates. Finally, legislative attention tends to be driven by the strategic choices of political actors (e.g., the president and party leaders) as well as unexpected events and institutional rules and processes (Adler and Wilkerson 2013; Baumgartner and Jones 1993; Kingdon and Thurber 1984).

Representation and policy responsiveness can also be affected by organized interest group influence rather than problem severity (see, e.g., Gilens and Page 2014). For example, research shows officials elected in off-cycle elections are more likely to pursue public policy that serves the interests of organized groups (Anzia 2013) and that congressional staff members often rely on interest groups to form policy positions and gauge constituent preferences (Hertel-Fernandez, Mildenberger, and Stokes 2019).

In addition, prior work has found evidence of racial inequality in legislative responsiveness. Such inequality can take the form of direct discrimination—for example, Butler and Broockman (2011) find that white legislators are more likely to respond to emails from putatively white constituents, while minority legislators respond to putatively black constituents more often. Legislators may also differ in responsiveness to the preferences of constituents in their districts. Following the 1992 redistricting, for instance, white incumbents who lost black constituents became less responsive to black policy preferences (Overby and Cosgrove 1996). Finally, in previous research, race has consistently been found to be a significant factor in welfare policy. For example, states with higher proportions of black welfare recipients have stricter eligibility rules and offer less generous benefits (Fellowes and Rowe 2004).

We consider whether such racial inequalities exist in drug policy, a domain in which the form of elite responsiveness may depend on the stereotypical race of a drug's users or the race of the victims themselves. As noted above, negative racial stereotypes invoked by the crack scare were associated with support for punitive responses to the issues of drugs and crime (Dixon 2006; Gilliam Jr. and Iyengar 2000; Golub and Johnson 1997; Hurwitz and Peffley 1997; Newkirk 2017). As such, deaths from cocaine, especially among nonwhite victims, may be especially likely to induce a fear-oriented policy response that emphasizes punishment (Dasgupta, Mandl, and Brownstein 2009; Harbin n.d.; Netherland and Hansen 2016). By contrast, victims of the opioid crisis are seen as stereotypically white and may be viewed more sympathetically (Keller 2017; Lopez 2017; McKenzie 2017; Peterson and Armour 2018). In fact, many have claimed that the opioid crisis inspired a more treatment-oriented policy response than did the crack scare because of racial inequality in US society (e.g., Glanton 2017; King 2017; Newkirk 2017).

To empirically test these claims, we measure legislative responsiveness to drug-related deaths, evaluating whether treatment- or punishment-oriented responses vary with the drug in question and the race of the victims. This approach allows us to address the concern that the difference in legislative responses between the two drug epidemics reflects a broader shift toward viewing drug addiction as a type of disease rather than a crime (see, e.g., Pew Research Center 2014).

To better understand these relationships, we also consider legislator responsiveness to deaths from methamphetamine, a drug predominately used by whites that has generated less public sympathy than opioids but has been portrayed less negatively than crack (Cobbina 2008; Murakawa 2011). The comparison to methamphetamine will help us better understand whether policy responses to the opioid crisis have been different because its victims are stereotypically white or because they might have addictions that began with prescription drugs.

Finally, we consider two possible moderators of the relationships of interest. First, given the evidence noted above that legislative responsiveness may vary by party (e.g., Adler, Cayton, and Griffin 2018; Kriner and Shen 2014), we test whether the relationship between drug-related deaths and subsequent (co)sponsorship of treatment- or punishment-oriented legislation differs between Democrats and Republicans. Second, research shows that media coverage can have important effects on legislative behavior (e.g., Arnold 2004; Snyder and Strömberg 2010). We therefore evaluate whether variation in media coverage influences legislative responsiveness to drug-related deaths using the Snyder and Strömberg (2010) approach of exploiting district congruence with media markets, which is a plausibly exogenous source of coverage variation. We specifically test whether the relationship between drug-related deaths and legislative responsiveness varies with district/media market congruence for deaths within the district and for deaths within the media market as a whole.[5]

Data

We measure the federal legislative response to the crack scare and the opioid crisis using data for the 96th–114th Congresses (1983–2016) from the Congressional Bills Project (Adler and Wilkerson n.d.).[6] We selected every bill from this period that had been coded as pertaining to drug and alcohol abuse ("related to alcohol and illegal drug abuse, treatment, education, and health effects") or to illegal drugs ("related to illegal drug crime and enforcement [and] criminal penalties for drug crimes, including international efforts to combat drug trafficking").[7] We then further coded the summary for each qualifying bill to exclude bills solely focused on alcohol and to identify bills that contained measures addressing criminal

5. Legislators may be responsive to drug problems in nearby areas outside their district that receive news coverage and thus prompt fears among their constituents.

6. The Congressional Bills Project labels bill summaries according to the topic coding system of the Policy Agendas Project (PAP). The PAP codebook is available at www.comparativeagendas.net/pages/master-codebook.

7. These PAP categories include bills addressing illegal drugs as well as those addressing abuse of prescription drugs. To ensure that we did not miss a substantial number of opioid-related bills related to legal prescription drugs, we searched the categories of bills "related to prescription drug coverage, programs to pay for prescription drugs, and policy to reduce the cost of prescription drugs" or "related to the regulation and promotion of pharmaceuticals, medical devices, and clinical labs" for the keywords pain, opioid, addict, and substance. These returned only 67 cases during a 44-year study period (1983–2016). We therefore did not include them in our analyses.

or civil penalties or promoting prevention, treatment, and rehabilitation (34 of the bills, or 2.3%, do both).[8] We then merge information on these bills with cosponsorship data from GovTrack.[9]

From these measures, we construct four simple binary measures of bill sponsorship and cosponsorship for each member of the House of Representatives from 1983 to 2016 at the year level.[10] Specifically, for each member of Congress, we measure whether they sponsored at least one prevention- or treatment-oriented bill related to drug abuse ("treatment bill") and whether they sponsored at least one punishment-oriented bill related to drugs ("punishment bill").[11] We then construct analogous measures for legislative cosponsorship, a symbolic but consequential act in which legislators officially indicate their support for a bill that another legislator has sponsored (Koger 2003).

Our primary independent variables are drug-related death rates by year at the congressional district level. To obtain these, we analyze confidential multiple cause of death data from the Division of Vital Statistics at the National Center for Health Statistics. These data provide individual-level records on the causes of death and contributing conditions for every American who dies in a given year. We identify the causes of death for each variable using ICD-9 and ICD-10 codes, which are provided for each death in the data. Following standard practices in the literature, we use a combination of diagnosis and external cause codes (ICD-9) and multiple cause of death codes (ICD-10) to identify cocaine-, opioid-, and methamphetamine-related deaths.[12] We specifically calculate the total number of drug poisoning deaths overall and separately for whites, blacks,

8. We sought to specifically identify bills that increased penalties for illegal drug use or drug abuse. We therefore excluded bills whose summaries specifically mentioned reducing penalties or specifically targeted drug distributors. Intercoder reliability ratings for the codings we employed in this study exceeded conventional norms in blind tests using randomized samples of bills. Results and detailed coding rules are provided in the online-only appendix.

9. The source is James H. Fowler, Andrew Scott Waugh, and Yunkyu Sohn, "Cosponsorship Network Data," jhfowler.ucsd.edu/cosponsorship.htm (accessed October 18, 2019).

10. We consider the set of legislators who served in each Congress during this period with data from the Legislative Effectiveness Project (Volden and Wiseman 2014). Each is considered to serve in both years except for those who left office in the first year of a given Congress because they died, resigned, etc. or entered office in the second year via appointment, special election, etc. (data from Stewart and Woon n.d.; Swift et al. 2000). We follow standard practice in the Congress literature and treat party switchers as new members after a switch and apply analogous year-level exclusions depending on the switch's timing.

11. We use binary measures due to concerns about skew in a small number of variables for the outcome measures and the greater robustness of ordinary least squares (Angrist and Pischke 2009).

12. See the online-only appendix for a detailed list of our coding rules. We note in particular that we observe no evidence of discontinuities in the aggregate time series of overall or drug-specific deaths during the switch from ICD-9 to ICD-10 in 1999 (see fig. 1). We thus pool the data over the study period.

and people from other racial/ethnic groups. We also calculate the total number of deaths related to opioids, cocaine, and methamphetamines. Finally, we calculate the total number of homicide deaths. We then aggregate these county-level totals, which are based on the location of the deceased's residence, by year at the congressional district level and divide them by the district population, transforming them into drug-related death rates.[13]

To consider the role of media coverage in political responsiveness to drug-related deaths, we construct two measures. First, because legislators might respond to media coverage of drug deaths outside of their district, we estimate drug-related death rates at the media market level using data from Gentzkow and Shapiro (2008). In addition, we use the Snyder and Strömberg (2010) measures of congruence between media markets and congressional districts to identify plausibly exogenous variation in coverage intensity that might affect legislator responsiveness to drug-related deaths in their district.

Results

We first present descriptive graphs and statistics for our primary independent and dependent variables, illustrating how drug death rates and legislative policy approaches to drug abuse have varied over our study period.[14] Figure 1 plots annual drug-related death rates by year for all drugs and for opioids and cocaine over the 1983–2016 period. As the figure indicates, drug-related death rates climbed modestly from 1983 to the early 2000s before accelerating in recent years, pushing the mortality rate to .19 per 1,000 people in 2016. This increase was largely driven by opioids. During the crack scare (1983–95), opioids and cocaine were associated with a nearly identical number of deaths despite widespread public and media attention to crack cocaine. Death rates from opioids began to outstrip cocaine death rates in the mid-1990s, however, rising from .02 per 1,000 people in 1995 to .13 per 1,000 in 2016. Opioids now kill far more Americans per year than all drugs did at the crack scare's peak.[15]

13. When counties were split across more than one congressional district, we allocated deaths proportionally with population weights from the most recent census, which covers the 98th Congress and later. We used redistricting data from Carson et al. 2007 to map congressional districts prior to the 1980 census redistricting to counties. We were not able to map 27 districts from this period to counties and thus restricted our main analyses to 1983 and later (results are very similar when including 1979–1982; available upon request).

14. Table A2 in the online-only appendix provides descriptive statistics of the key variables.

15. See figure A1 in the online-only appendix for corresponding race-specific death rates per 1,000 Americans.

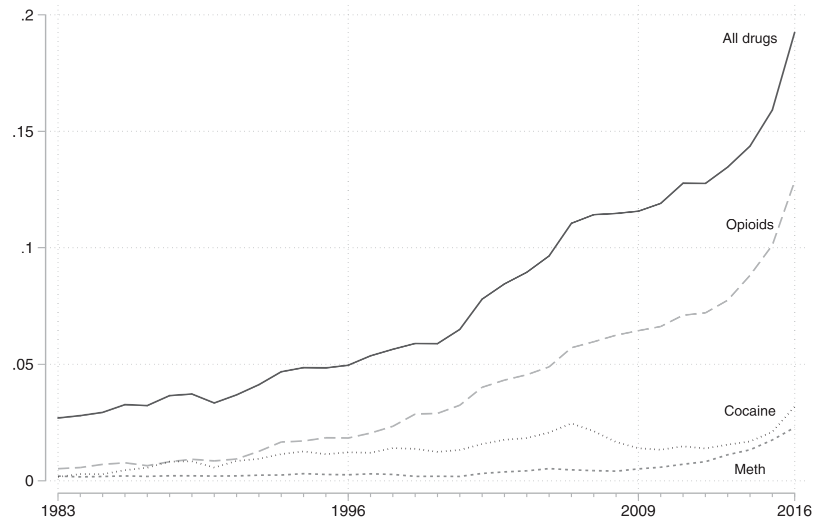

Figure 1 Yearly drug-related death rates (per 1,000 people).

Notes: Total drug-related deaths per year from all illegal drugs and from opioids, methamphetamines, and cocaine. Calculated with data from the National Center for Health Statistics (see the online-only appendix for coding details).

To understand how policy approaches to drug abuse vary over this time period, figure 2 presents smoothed models of over-time variation in legislative policy approaches to drugs. These estimates start in 1979 to show pre-study period trends and avoid extrapolation in the local polynomial fits. The figure shows lawmakers introduced more drug-related bills during the crack scare than later on. When compared with figure 1, which shows that far more people have died of drug poisoning in recent years, this figure demonstrates a striking lack of correspondence between drug mortality and policy responses.[16] While the figure shows that the number of drug-related bills has been increasing during the opioid crisis, the total is still less than in the mid-1980s. These data also indicate that legislators were more likely to sponsor bills that proposed a punishment-oriented approach to drug abuse and addiction than a treatment-oriented approach during the crack scare of 1983–95.[17] This differential was no longer consistently

16. For example, the total number of bills sponsored decreased in the mid-1990s despite the fact that neither cocaine deaths nor overall drug deaths decreased during that period. This decline is likely linked to the decline in media attention to the crack scare around that time (Hartman and Golub 1999).

17. The smoothed year-level estimates in these graphs range from 0 to 0.03, but the yearly data vary from 0.01 to 0.07 for punishment bills (1989, when 30 members [7%] introduced bills) and from 0 to 0.04 for treatment bills (1991, when 16 members [4%] introduced bills).

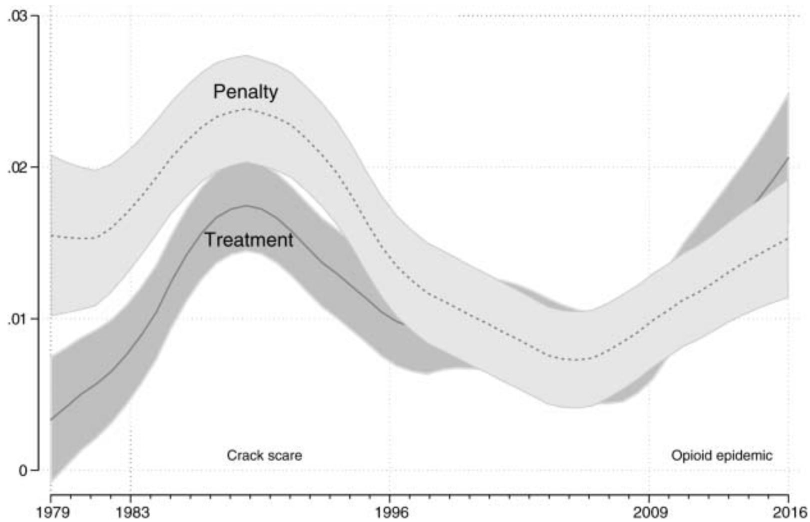

Figure 2 Drug abuse bill sponsorship rates by policy approach.

Notes: Outcome is a yearly binary indicator of sponsorship of one or more bills related to treatment or punishment of drug abuse among members of the House of Representatives (local polynomial fits with bandwidth of three years). Data from the Congressional Bills Project (Adler and Wilkerson n.d.).

measurable after 1995. Indeed, the number of treatment bills has been growing more rapidly than punishment bills since the beginning of the opioid crisis, and the mid-2010s represent the first time that treatment-oriented policy responses became more prevalent than penalty-oriented ones in our data.

Next, we estimate a series of ordinary least squares (OLS) models to evaluate our theoretical expectations. Each model predicts legislative bill (co)sponsorship using drug-related death rates in a legislator's House district. We calculate separate measures by drug type (total drug-related deaths, opioid deaths, and cocaine deaths) and by race of victim (white or black).[18] Each death rate measure is calculated as the total number of deaths per 1,000 district residents and lagged by one year to ensure a plausible temporal relationship between deaths and bill sponsorship.[19] These models include legislator fixed effects to account for time-invariant

18. In this study we focus specifically on white deaths, because they are the majority racial group at the national level, and black deaths, because they are the group critics argue have been treated the worst in drug policy.

19. We normalize by district population to ensure the death rate measures are comparable across districts.

factors such as party that might induce a spurious relationship between drug deaths and (co)sponsorship of drug-related bills.[20] We use legislator fixed effects rather than district fixed effects because districts change over time due to redistricting, and legislators tend to behave quite consistently (see, e.g., Poole and Rosenthal 2007). These fixed effects account for all baseline differences among legislators, allowing us to capture how changes in drug death rates in each legislator's district predicts changes in his or her drug policy responses in the coming year. It is therefore not necessary to control for legislator party or other time-invariant characteristics. We also include fixed effects by census region-year to account for correlated temporal shocks (possibly region specific) that do not vary across legislators or districts and could produce spurious relationships such as changes in national drug policy, differences in the availability of different kinds of drugs over time such as fentanyl, and the growth in support for criminal justice reform in recent years. These fixed effects also account for any (linear or nonlinear) national trends that are correlated across districts. Finally, we separately cluster the standard errors by legislator and region-year to account for any remaining within-legislator or within-region-year correlation (e.g., legislators reintroducing bills they have sponsored repeatedly over time). These fixed effects models are identified using temporal variation in drug-related deaths within (not between) legislators, which we take as exogenous. Per Mummolo and Peterson (2018), we present summary statistics for this identifying within-district variation in table A3 in the online-only appendix.

We begin our analysis by estimating the relationship between district drug-related deaths and subsequent sponsorship and cosponsorship of treatment-oriented bills (tables 1a and 1b, respectively). Our results indicate that drug-related deaths are significantly positively associated with subsequent legislative sponsorship of treatment-oriented bills ($p < .05$; column 1 of table 1a). However, this pattern of responsiveness to drug deaths varies by drug type and victim race. We find that opioid deaths are significantly associated with subsequent sponsorship of treatment-oriented legislation ($p < .01$ in column 2, $p < .005$ in column 5) but cocaine and methamphetamine deaths are not (columns 3 and 4). Similarly, deaths of white drug victims are positively associated with subsequent treatment bills ($p < .005$; columns 6 and 8). By contrast, deaths of black victims are

20. The key identifying assumption of a fixed effects model is that a confounding variable does not vary over time. This assumption would not hold if changes in crime rates are correlated with both drug severity and policy responses. We address this concern below by showing that our results are robust for controlling for district-level homicide rates.

Table 1 Sponsorship/Cosponsorship of Treatment-Oriented Bills by Prior Drug Deaths

(a) Sponsorship of Treatment-Oriented Bills

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Total drug deaths	0.196*							
	(0.078)							
Opioid deaths		0.296**			0.432***			
		(0.112)			(0.135)			
Cocaine deaths			–0.055		–0.683*			
			(0.265)		(0.334)			
Meth deaths				0.853	0.567			
				(0.476)	(0.472)			
White drug deaths						0.274***		0.327***
						(0.082)		(0.080)
Black drug deaths							–0.371	–0.690*
							(0.249)	(0.283)
Legislator fixed effects	Y	Y	Y	Y	Y	Y	Y	Y
Region-year fixed effects	Y	Y	Y	Y	Y	Y	Y	Y
Legislators	1243	1243	1243	1243	1243	1243	1243	1243
Total N	13008	13008	13008	13008	13008	13008	13008	13008

Table 1 *(continued)*

(b) Cosponsorship of Treatment-Oriented Bills

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Total drug deaths	0.096							
	(0.214)							
Opioid deaths		0.560*			1.050***			
		(0.237)			(0.264)			
Cocaine deaths			–0.655		–2.090***			
			(0.553)		(0.680)			
Meth deaths				–0.652	–1.250			
				(1.008)	(0.994)			
White drug deaths						0.170		0.230
						(0.244)		(0.244)
Black drug deaths							–0.561	–0.786
							(0.661)	(0.685)
Legislator fixed effects	Y	Y	Y	Y	Y	Y	Y	Y
Region-year fixed effects	Y	Y	Y	Y	Y	Y	Y	Y
Legislators	1243	1243	1243	1243	1243	1243	1243	1243
Total N	13003	13003	13003	13003	13003	13003	13003	13003

* p < 0.05, ** p < .01, *** p < .005 (two-sided *p* values); OLS models with two-way clustering by legislator and region-year (Cameron, Gelbach, and Miller 2011; Correia 2016). Constant is suppressed. Outcome variables are a binary measure of legislative sponsorship or cosponsorship of one or more bills related to treatment of drug use among members of the House of Representatives during the 1983–2016 period. All drug-related death variables are calculated from mortality records as deaths per 1,000 district residents and are lagged by one year.

negatively associated with treatment legislation when entered into the same model as deaths of white victims ($p < .05$; column 8), though we cannot be certain that this negative coefficient reflects a causal relationship.[21] Importantly, we can reject the nulls of no difference between the effects of opioid- and cocaine-related deaths ($p < .05$; column 4) and between the effects of deaths of white and black victims ($p < .005$; column 7). We observe a similar pattern in table 1b of differential responsiveness in cosponsorship of treatment-oriented bills to opioid- and cocaine-related deaths ($p < .005$; column 5), though we find no measurable difference by victim race (column 8).

To interpret the magnitude of these relationships, it is important to note first that the base rate of treatment-oriented drug bill sponsorship is only 1.6%. We must also consider the range of variation in white drug deaths accounting for legislator and region-year fixed effects (Mummolo and Peterson 2018). If white drug deaths increased by two standard deviations, the expected increase in the likelihood of treatment bill sponsorship using the results from column 6 of table 1a is 1.04 percentage points, which represents an increase of 83% in relative terms from the treatment bill sponsorship base rate of 1.26%.[22] An analogous increase of two standard deviations in within-district opioid deaths would generate a 1.24 percentage point increase in the likelihood of treatment bill sponsorship (a 98% increase in relative terms).

The patterns of differential treatment-oriented responses to drug deaths by victim race (table 1a) and type of drug (table 1b) that we describe above do not clearly hold for punishment-oriented bills, however. Table 2a shows no significant association between prior-year drug deaths and sponsorship of punishment-oriented bills regardless of whether we consider total drug deaths (column 1) or disaggregate them by type of drug (columns 2–4) or race of victim (columns 5–7). We therefore do not find evidence that objective conditions at the local level affect the decision to sponsor such bills (unlike the relationship we observe at the national level, which is presented in figure 2). Interestingly, in table 2b, we do observe evidence that legislators are more responsive in cosponsoring punishment-oriented legislation as opioid deaths and white drug deaths increase (columns 5 and 8), which is somewhat inconsistent with our theoretical expectation. Taken together with table 1, this finding suggests that representatives may be

21. The negative association is statistically significant in the model for controlling for white victims ($p < .05$; column 8) but not in the bivariate model (column 7). We therefore cannot rule out the possibility that the column 8 estimate reflects posttreatment bias (the same factors causing black deaths may also be causing white deaths).

22. Such an increase would, if generalized, translate into more than five additional bills that year across 435 House members.

Table 2 Sponsorship/Cosponsorship of Punishment-Oriented Bills by Prior Drug Deaths

(a) Sponsorship of Punishment-Oriented Bills

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Total drug deaths	–0.039							
	(0.067)							
Opioid deaths		–0.077			0.009			
		(0.070)			(0.085)			
Cocaine deaths			–0.389		–0.406			
			(0.362)		(0.437)			
Meth deaths				0.032	0.113			
				(0.397)	(0.408)			
White drug deaths						–0.045		–0.044
						(0.056)		(0.045)
Black drug deaths							–0.060	–0.017
							(0.432)	(0.427)
Legislator fixed effects	Y	Y	Y	Y	Y	Y	Y	Y
Region-year fixed effcts	Y	Y	Y	Y	Y	Y	Y	Y
Legislators	1243	1243	1243	1243	1243	1243	1243	1243
Total N	13003	13003	13003	13003	13003	13003	13003	13003

(*continued*)

Table 2 Sponsorship/Cosponsorship of Punishment-Oriented Bills by Prior Drug Deaths (*continued*)

(b) Cosponsorship of Punishment-Oriented Bills

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Total drug deaths	0.307*							
	(0.139)							
Opioid deaths		0.637**			0.820**			
		(0.162)			(0.224)			
Cocaine deaths			0.366		–0.767			
			(0.521)		(0.676)			
Meth deaths				0.049	–0.611			
				(1.189)	(1.186)			
White drug deaths						0.375*		0.403**
						(0.152)		(0.154)
Black drug deaths							0.035	–0.358
							(0.587)	(0.616)
Legislator fixed effects	Y	Y	Y	Y	Y	Y	Y	Y
Region-year fixed effects	Y	Y	Y	Y	Y	Y	Y	Y
Legislators	1243	1243	1243	1243	1243	1243	1243	1243
Total N	13003	13003	13003	13003	13003	13003	13003	13003

* $p < 0.05$, ** $p < .01$, *** $p < .005$ (two-sided *p* values); OLS models with two-way clustering by legislator and region-year (Cameron, Gelbach, and Miller 2011; Correia 2016). Constant is suppressed. Outcome variables are a binary measure of legislative sponsorship (table 2a) or cosponsorship (table 2b) of one or more bills related to punishment of drug abuse among members of the House of Representatives during the 1983–2016 period. All drug-related death variables are calculated from mortality records as deaths per 1,000 district residents and are lagged by one year.

willing to respond to or address opioid deaths and white deaths using punitive as well as treatment-oriented approaches. However, this conclusion must be treated as tentative—unlike in table 1, we cannot reject the null of no difference in effects with cocaine deaths and black drug deaths, respectively.

Broadly, these results suggest that district-level drug deaths increase the (co)sponsorship of treatment-oriented bills, especially for opioid overdoses and when the victims are white, but do not have a strong or consistent effect on punishment-oriented bills. In tables 3 and A4, we examine the extent to which this pattern varies over time. To do so, we estimate versions of previous models predicting sponsorship of treatment- or punishment-oriented bills in which we interact drug deaths with indicators for the crack era, which we define as 1983–95, and the opioid era, which we define as 2009–16.[23] The coefficients on the interaction terms test whether these relationships vary by era compared to the reference period of 1996–2008.

We first consider differences over time in responsiveness to drug deaths with sponsorship of treatment-oriented bills. Consistent with our expectation, table 3 indicates that legislators respond to drug deaths in the opioid era with treatment-oriented legislation ($p < .05$ for the marginal effect in the 2009–16 period), though we cannot reject the null of no difference in effects with the other two eras (column 1). The story becomes clearer when we focus specifically on deaths by drug type. The fully specified model considering opioid, cocaine, and methamphetamine deaths (column 5) shows that legislators sponsored more treatment bills as opioid deaths increased in their district in the interim period (1996–2008; $p < .01$) and especially during the opioid crisis (2009–16 marginal effect; $p < .005$). As a result, we can reject the null of no difference in the relationship between opioid and cocaine deaths only during the opioid crisis ($p < .05$). We also find that white drug deaths were significantly associated with treatment bill sponsorship in the interim period and during the opioid crisis, but not during the crack era (columns 6 and 8). Moreover, this relationship is significantly different from the one observed for black drug deaths in the interim period and opioid crisis ($p < .005$).

For ease of understanding, we plot the relationship between drug deaths and treatment bills by era in figure 3 (based on columns 1–4 and 6–7 of table 3). It shows that the relationship between drug deaths and treatment-oriented bills has become stronger overall (figure 3a). However, this finding

23. By 2009 opioid overuse had become a sufficient enough concern that the Federal Drug Administration launched its Safe Use Initiative (FDA 2019). Heroin deaths started to rise in 2011, and synthetic opioid deaths began to increase in 2014 (Ciccarone 2019).

Table 3 Sponsorship of Treatment-Oriented Drugs Bills by Time Period and Prior Drug Deaths

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Total drug deaths	0.106*							
	(0.050)							
Total × crack era	0.044							
	(0.096)							
Total × opioid era	0.162							
	(0.098)							
Opioid deaths		0.099			0.251**			
		(0.077)			(0.093)			
Opioid × crack era		0.108			−0.289			
		(0.296)			(0.300)			
Opioid × opioid era		0.316*			0.338*			
		(0.127)			(0.148)			
Cocaine deaths			−0.221		−0.603			
			(0.258)		(0.329)			
Cocaine × crack era			0.433		0.713			
			(0.361)		(0.416)			
Cocaine × opioid era			0.256		−0.639			
			(0.348)		(0.407)			
Meth deaths				0.202	0.187			
				(0.479)	(0.499)			
Meth × crack era				0.011	0.128			
				(0.961)	(0.962)			

Table 3 (*continued*)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Meth × opioid era				1.085*	0.550			
				(0.534)	(0.595)			
White drug deaths						0.193***		0.245***
						(0.055)		(0.059)
White × crack era						–0.053		–0.174
						(0.143)		(0.178)
White × opioid era						0.140		0.135
						(0.102)		(0.100)
Black drug deaths							–0.517	–0.803**
							(0.272)	(0.306)
Black × crack era							0.569*	0.639
							(0.281)	(0.339)
Black × opioid era							–0.150	–0.109
							(0.209)	(0.215)
Legislator fixed effects	Y	Y	Y	Y	Y	Y	Y	Y
Region-year fixed effects	Y	Y	Y	Y	Y	Y	Y	Y
Legislators	1243	1243	1243	1243	1243	1243	1243	1243
Total N	13003	13003	13003	13003	13003	13003	13003	13003

* $p < 0.05$, ** $p < .01$, *** $p < .005$ (two-sided p values); OLS models with two-way clustering by legislator and region-year (Cameron, Gelbach, and Miller 2011; Correia 2016). Constant is suppressed. Outcome is a binary measure of legislative sponsorship of one or more treatment-oriented bills related to drug abuse or illegal drugs among members of the House of Representatives during the 1983–2016 period. All drug-related death variables are calculated from mortality records as deaths per 1,000 district residents and are lagged by one year. The crack era is defined as 1983–95 and the opioid era as 2009–16 (reference category is 1996–2008).

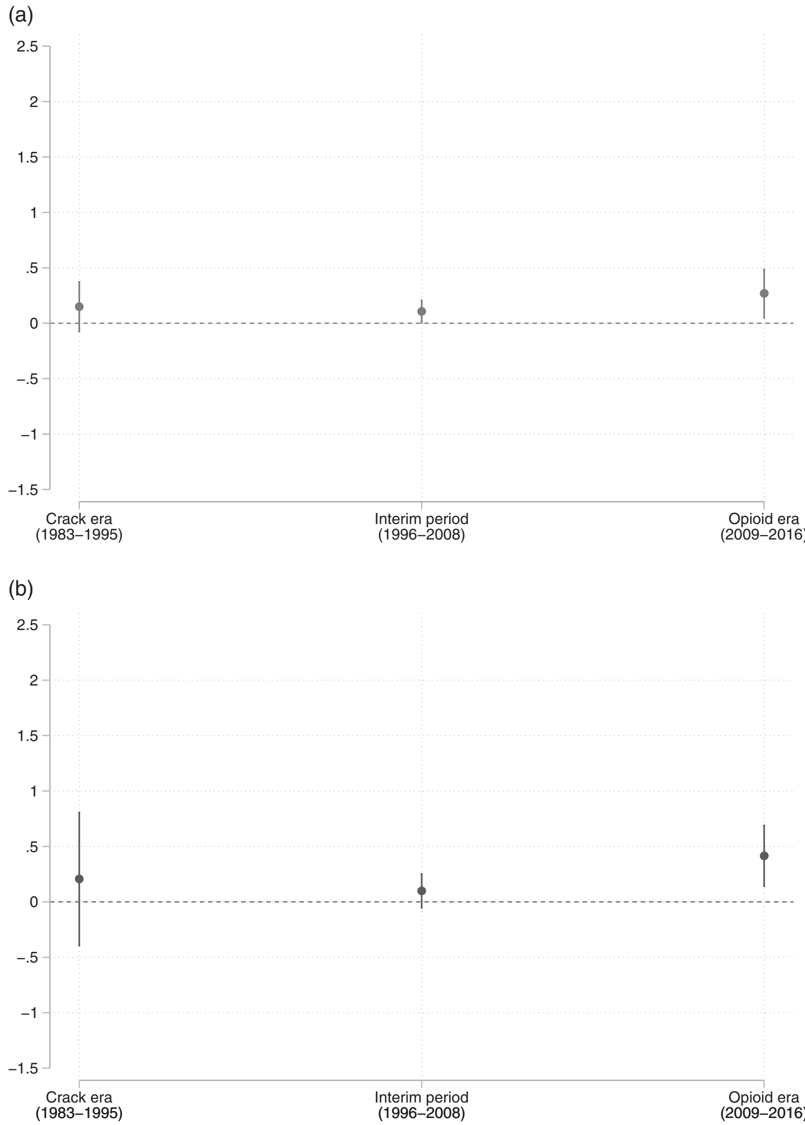

Figure 3 Marginal effects of drug deaths on treatment sponsorship by drug type/race of victim: (a) all drugs; (b) opioids; (c) cocaine; (d) methamphetamine; (e) white victims; (f) black victims.

Notes: Marginal effects of drug deaths on sponsorship of one or more bills related to treatment of drug use among members of the House of Representatives. Figures 3a, 3b, 3c, and 3d correspond to models 1, 2, 3, and 4 in table 3. Figures 3e and 3f correspond to models 6 and 7 in table 3. All drug-related death variables are calculated from mortality records as deaths per 1,000 district residents and are lagged by one year (see the online-only appendix for coding details). Data from the Congressional Bills Project (Adler and Wilkerson n.d.).

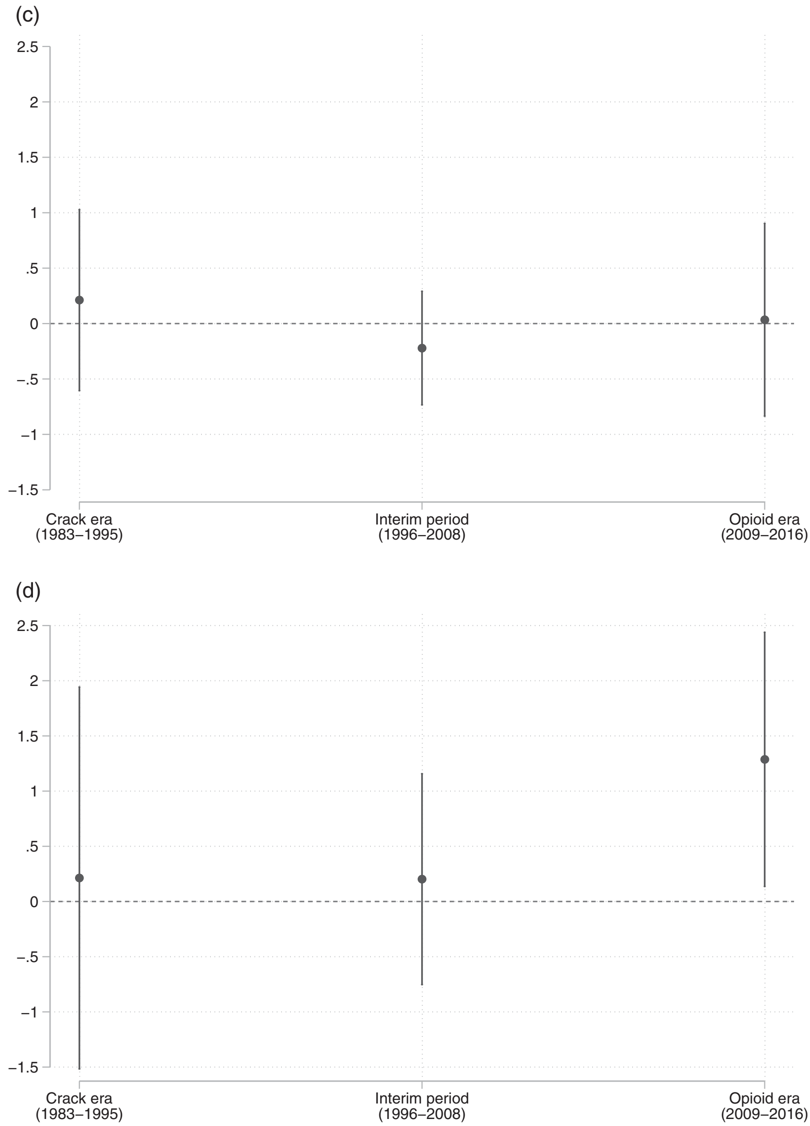

Figure 3 (continued)

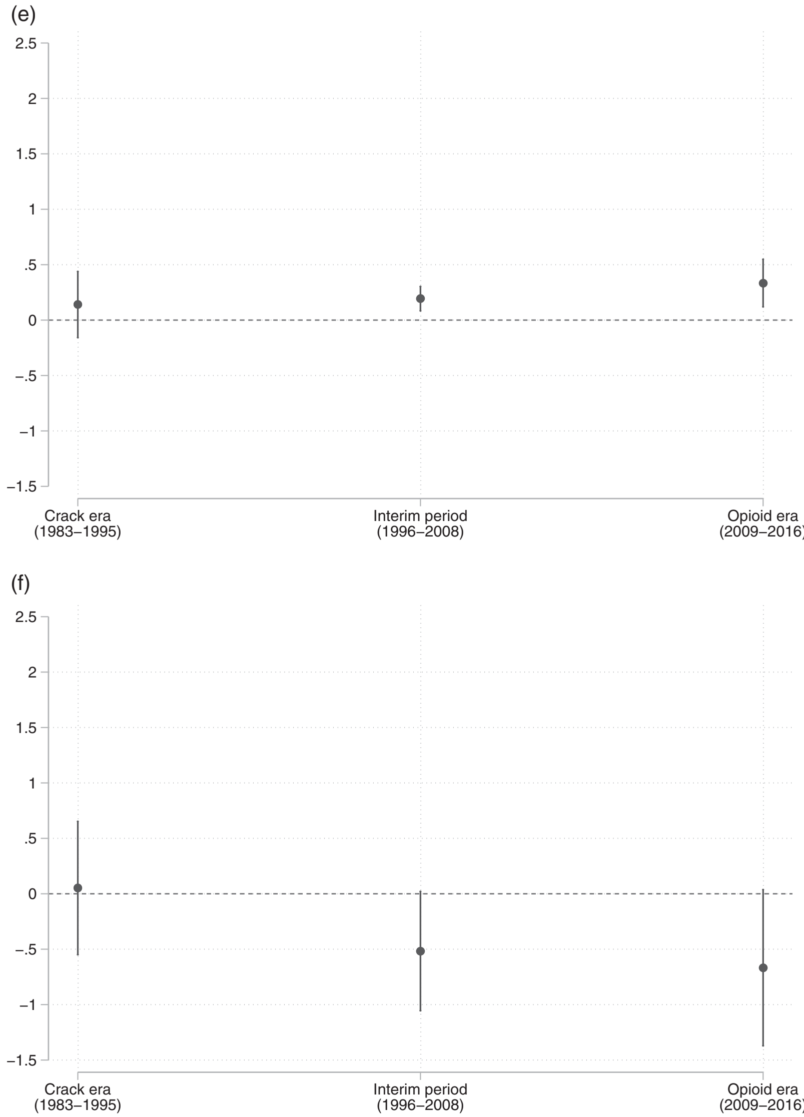

Figure 3 (continued)

holds for opioid-related deaths (figure 3b) and methamphetamine-related deaths (figure 3d), but not those related to cocaine (figure 3c), which are not estimated precisely. The shift over time toward treatment-oriented responses seems to be driven by drugs that are believed to have most affected white communities.[24]

Similarly, there is a noticeable difference in treatment policy responses to drug deaths by victim race. Specifically, legislators are responsive to the deaths of white victims (figure 3e) and not to black victims (figure 3f)—a racial gap in treatment-related drug policy that is widest during the current opioid crisis.

None of the relationships we describe above are observed for punishment-oriented bills, however (table A4). We find no significant association between prior-year drug deaths and bill sponsorship in any era for any measure. These results suggest that sponsorship of punishment-oriented bills is not a response to the local severity of drug use (legislators may instead be responding to other factors such as media coverage or public opinion).

We next consider the robustness of our results. The most plausible threat to our design is that changes in drug deaths are correlated with other crimes and that our models pick up the effects of these non-drug-related crimes. To address this concern, we calculate homicide rates at the district level (the measure of crime incidence that is most consistently measured across districts) and replicate all the models reported in the text above with this measure as a control variable. Our results, which are reported in tables A5–A8 of the online-only appendix, are very similar to those discussed above.

Finally, we consider two possible moderators of the relationships we observe—political parties and media coverage. We first estimate whether the relationship between district-related drug deaths and sponsorship of treatment- or punishment-oriented legislation varies by party identification. Tables A9 and A10 in the online-only appendix show, however, that we cannot reject the null hypothesis of no difference in responsiveness (i.e., the relationship between district-level drug deaths and bill sponsorship) between Democrats and Republicans across all the measures of drug-related deaths considered above (i.e., by race and drug type). Similarly, we observe little consistent evidence that the relationship between drug deaths and legislative responsiveness varies by congruence between media markets and congressional district boundaries (tables A11–A12). Finally, responsiveness to deaths at the media market level is similar to deaths at

24. The estimated marginal effect size is larger for methamphetamines, which likely reflects the reduced level of within-legislator variation in deaths compared to opioids.

the district level (table A13) but these relationships again do not measurably vary by congruence (tables A14–A15), providing little convincing evidence that media coverage drives responsiveness.

Conclusion

The opioid crisis has sparked a public debate over how policy makers have addressed drug abuse in recent decades. Many have suggested that policy responses to the opioid crisis emphasize treatment and have speculated that race explains the changes (e.g., Peterson and Armour 2018), but lack convincing evidence for this claim. Our study provides the first systematic comparison of the federal legislative policy response to the crack scare and the opioid crisis. Despite the massive increase in opioid overdose deaths since 2009, legislators were more likely to sponsor legislation related to drugs during the crack scare. However, we do find that members of Congress have responded to the opioid crisis by proposing more treatment-oriented policies. We focus specifically on legislator responsiveness to local drug deaths, which we find varies by type of drug and race of victim (but not legislator party or factors that affect media coverage). Policy makers appear to respond to district-level drug mortality by increasing the likelihood that they propose treatment-oriented legislation, but this response is driven by responsiveness to victims of opioid overdoses (especially in recent years) and to white drug deaths. Punitive drug policies, by contrast, seem to be unrelated to district-level mortality rates. These results suggest that the political system is differentially sensitive to the suffering of white victims of the opioid crisis and is unusually willing to offer treatment-oriented policies on their behalf.

Importantly, the differential responsiveness by race found in this study is consistent with previous research that has documented various forms of racial discrimination in political representation (e.g., Butler and Broockman 2011; Overby and Cosgrove 1996). Our study contributes to research on the interaction of race and representation by showing how urgent problems such as drug epidemics can generate unequal policy responses depending on which racial groups are affected. Similar forms of double standards and differential treatment likely exist in other issue domains such as welfare or criminal justice (see, e.g., Fellowes and Rowe 2004), though more research is needed to better understand the objective conditions of black and white communities and to compare legislative responses to those conditions. Further research is also needed to determine how to most effectively improve representation for disadvantaged communities and mitigate these inequalities.

Of course, this study has limitations that should be noted. First, we do not consider congressional voting, constituent service, floor speeches, or other forms of potential legislative responsiveness to district conditions. Second, we focus on legislative responsiveness to drug-related deaths, not other harms from drug use. Our reliance on mortality data necessarily highlights opioids due to the greater likelihood of lethal overdoses from its use (especially with synthetic opioids like fentanyl). These data also span the 1999 transition from ICD-9 to ICD-10 cause of death codes, which may reduce the comparability of data between the crack scare and the opioid crisis (though we observe no evidence of discontinuities). Third, future research should seek to extend this approach to study drug policy responsiveness in the states, which plays a critical role in both criminal justice and public health policy under the US federal system. It would also be valuable to examine variation in responses by prosecutors, judges, public health agencies, and other federal and state actors to local-level drug mortality during the opioid crisis. Finally, scholars should devote further attention to possible changes over time in the content of media coverage of drug use, a possible mechanism of support for treatment-oriented policies (e.g., Harbin n.d.).

Still, this study represents an important step toward understanding the forces shaping the policy response to one of the most important issues in contemporary US politics. Given the staggering human toll of the opioid crisis, the stakes could hardly be higher.

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Acknowledgments

We thank Jack Davidson and Lucas Maiman for outstanding research assistance and Devin Caughey, Mia Costa, Jesse Crosson, Alexander Furnas, Sarah Gollust, Gregory Koger, Josh McCrain, Andrew Reeves, Christopher Warshaw, and participants in the Politics of the Opioid Epidemic conference at Brown University for helpful comments. We are also grateful to the Congressional Bills Project, the Legislative Effectiveness Project, and the National Center for Health Statistics for providing the data used in our analysis, and to the Stamps Family Charitable Foundation and the Dartmouth Politics and Law Program for supporting Morgan's work on the project. Finally, we are grateful to our colleague D. J. Flynn, who helped conceive the article and code the legislative bills data.

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