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Transparency, Class Bias, and Redistribution: Evidence from the American States

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Abstract

This study employs state-level panel data between 1978 and 2000 to explore the relationship between transparency, media market penetration, class bias in voter participation, and welfare effort in the United States. I present empirical evidence that the effect of transparency—operationalized as state fiscal transparency—on state welfare effort is conditional on class bias in voter participation. Specifically, I present evidence that in states where transparency and class bias increased over time, state welfare effort significantly declined. These results are robust to the inclusion of controls for other determinants of redistribution that traditionally vary with geography such as governor partisanship, legislator ideology, citizen ideology, gross state product (GSP), and state demographic characteristics, and are robust across several alternate model specifications. My findings suggest that increased transparency does not automatically improve the condition of socioeconomically-disadvantaged citizens and that transparency may have welfare-reducing effects in societies with increasing participatory gulfs between the most and least advantaged citizens.

Keywords

transparency, welfare, inequality, panel data

Transparency is generally perceived as a powerful and necessary tool for reducing democratic deficits and inequalities between groups in society (Gurria 2017). Good governance reformers argue that greater openness enhances governance quality by forcing those in positions of power to be accountable to the public who—in democratic societies—possess the capacity to punish legislators for self-enrichment through

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elections. In an effort to forestall being sanctioned by the public for such self-enrichment, legislators should theoretically propose more equitable divisions of resources in governments with greater levels of transparency.

The argument that greater transparency incentivizes good governance and greater redistribution by politicians seems quite straightforward, and from both theoretical and empirical perspectives, it enjoys some support (Kaufmann, Mastruzzi, and Zavaleta 2002; Reinikka and Svensson 2004; Rose-Ackerman 2005). However, it is not immediately obvious that the effects of transparency on governance outcomes are unequivocally positive. Indeed, in some cases, transparency interventions have been shown to restrict the range and quality of deliberation in legislative negotiation (Binder and Lee 2015; Mansbridge 2009) and to reduce the performance and responsiveness of legislators (Malesky, Schuler, and Tran 2012). Moreover, some studies suggest that transparency interventions funnel benefits to groups in societies that already enjoy states of relative advantage to the detriment of the poor (Peisakhin and Pinto 2010; Ross 2006). In these instances, transparency interventions may have no discernible effect on reducing inequalities between the wealthiest and poorest citizens. It is, thus, imperative for scholars to accurately identify the conditions under which transparency interventions promote the well-being of the most disadvantaged groups in society and the conditions under which transparency interventions leave the disadvantaged behind.

In this article, I investigate the effect of transparency on state welfare effort in the context of U.S. states. I posit that the success of transparency and accountability initiatives in promoting redistributive effort is not conditioned by the extent to which information is diffused in a society as has recently been argued but is conditioned by the extent to which the capacity to sanction legislators is stratified by socioeconomic status. Previous studies have demonstrated that the information released by transparency initiatives is disproportionately consumed by higher-income citizens (Dugdale et al. 2005; Nam and Sayogo 2011); that higher-income citizens are more likely to vote than the poor, and thus, to sanction legislators for perceived nonresponsiveness (Franko, Kelly, and Witko 2016; Verba, Schlozman, and Brady 1995); and that to the extent that they are responsive to their constituents, legislators are more responsive to the preferences of higher-income citizens than they are to the preferences of lower-income citizens (Gilens and Page 2014) and to the preferences of voters over nonvoters (Calvo 2007; Peters and Ensink 2015). I, thus, hypothesize that the effect of transparency on state welfare effort will not be significantly conditioned by media market penetration but will be significantly negatively conditioned by class bias in voter participation.

To test these hypotheses, I use state-level panel data between 1978 and 2000 to explore the relationship between transparency, media market penetration, class bias in voter participation, and welfare expenditures in the United States. Using a series of "between-within" panel models, I present evidence that the effect of transparency on state welfare effort is conditioned by changes in electoral class bias within a state over time. Specifically, states in which both transparency and class bias in voting increased over time experienced statistically significant decreases in state welfare effort. These results are robust to the inclusion of controls for other determinants of redistribution that traditionally vary with geography such as governor partisanship, legislator ideology, citizen ideology, gross state product (GSP), income inequality, unemployment rate, and other state demographic characteristics. To the author's knowledge, this is the first study within the state politics literature to explicitly examine the relationship between fiscal transparency and state welfare effort. Moreover, this study methodologically improves on previous cross-national research on the contingencies of transparency by leveraging panel data and a robust between-within model specification that permits simultaneous estimation of heterogeneous longitudinal and cross-sectional effects.

My findings suggest that increases in transparency may have welfare-reducing effects in societies with ever widening participatory gulfs between the most and least advantaged citizens. The implications of these findings for policy practitioners and program sponsors are clear: transparency and accountability initiatives should move beyond the release of information alone and do more to develop the informational and mobilization capacities of the least well off if the ultimate goal of reducing between group inequality is desired.

The Transparency-Accountability Relationship

Good governance reformers and civil society campaigners often promote transparency—defined as "an openness of the governance system through clear processes and procedures and easy access to public information for citizens" (Kim et al. 2005, 649)—as an effective way to enhance democratic accountability, democratize political voice, and redress unequal power relations (Keohane and Nye 2003; Stiglitz 2001). Articulating this widely held belief, then-Senator Barack Obama remarked in a 2007 press release that

> the more people know about what's going on in Washington, and how their tax dollars are being spent, and who's raising money for who, the less likely it is that major decisions will be hijacked by lobbyists and special interests. (Obama 2007)

This argument—that transparency improves the quality of democratic representation by reducing the disproportionate political voice of advantaged groups such as lobbyists and interest groups—is premised on the belief that the release of information to the public will deter those in positions of power—primarily legislators and other power-brokers in society—from using their positions of power to engage in self-enrichment (Meijer 2014; Mohtadi and Roe 2003). Specifically, this argument holds that legislators in transparency-rich, democratic environments should pursue more egalitarian policies and, by implication, favor redistributive policies because they are motivated by concerns about their re-election prospects—a relatively straightforward instance of a classic principal-agent problem (Besley 2006; Hollyer, Rosendorff, and Vreeland 2011; Mayhew 1974). Stated differently, in more open societies in which a greater proportion of citizens of different socioeconomic statuses can directly observe the behavior of legislators, legislators face strong electoral incentives to not self-enrich and to instead distribute resources more evenly than they potentially would otherwise (Kolstad and Wiig 2016).

To this point, Malena et al. (2004, 5) note,

> Given the difficulty, inability or unwillingness of governments to deliver essential services to their citizens—especially the poorest, enhanced accountability initiatives that allow greater articulation of citizens' demands and increased transparency of public decision-making increase the effectiveness of service delivery and produce more informed policy design.

Empirically, some studies have found a positive association between transparency, political voice, and public goods provision specifically to the poor, though it is worth noting that this work has almost exclusively been conducted at the cross-national level (Kaufmann, Mastruzzi, and Zavaleta 2002; Rose-Ackerman 2005). In addition, agency-based formal models developed by Ferejohn (1999) and Shi and Svensson (2003) predict that greater transparency produces more responsive politicians and generates public confidence in legislators because it alleviates information asymmetries between voters and politicians. Thus, according to this perspective, greater transparency and accountability jointly work to empower the poorest citizens, disciplining politicians by forcing them to be responsive to the redistributive demands of the least well off.[1]

However, these arguments in defense of greater transparency that constitute the conventional wisdom make three assumptions about the relationship between transparency and redistributive outcomes that warrant greater scrutiny: (1) that information release alone automatically produces accountability and responsiveness on the part of elected officials, (2) that access to information released by transparency interventions is uniformly distributed across the population, and (3) that mobilization and sanctioning (e.g., accountability) capacities are similarly uniformly distributed across the general population.

Recent studies cast doubt on this conventional wisdom and suggest that increases in transparency are not always associated with commensurate increases in accountability.[2] As Fox (2007, 665) notes, "transparency initiatives which mobilize the power of shame have no purchase on the shameless," and as others have noted, it is not a given that information release is sufficient to induce meaningful behavioral changes among legislators (McGee and Gaventa 2011; for exception, see Haley and Fessler 2005).

Transparency might not automatically translate into increased accountability for any number of reasons. When those in positions of power are indifferent to citizen preferences, as Fox (2007) notes, then increased information exposure may have little effect on good governance outcomes and by extension redistributive behavior. For instance, previous research has found that the export of transparency initiatives from electoral democracies to more authoritarian regimes does not considerably improve delegate performance or the quality of citizen representation (Malesky, Schuler, and Tran 2012). However, in more democratic societies, the empirical evidence suggests that legislators are not altogether indifferent to citizen demands even if they are more attentive and responsive to the demands of some groups of citizens over others (Gilens and Page 2014). Additionally, Lindstedt and Naurin (2010) argue that transparency may not generate accountability in countries where educational institutions are not well-developed, where media circulation is limited, and where electoral democracy is weak. If citizens have difficulty figuring out how to physically access information released by transparency interventions, then transparency may exert little effect on accountability to the extent that access is a necessary precondition for and motivator of political mobilization (Burnett, Jaeger, and Thompson 2008; Jaeger and Bertot 2010). Relatedly, even when the information released by governments or other agencies is physically accessible, if this information is complex or cognitively inaccessible, then transparency may have little effect on accountability (Jaeger and Bertot 2010). Finally, information release alone—however anger-inducing—may do little to reduce other significant barriers to collective action among the already disadvantaged (Klanderrmans 1996). Even when citizens know how to physically access information released by transparency initiatives, are able to understand this information, and are able to perceive that their group has been wronged in some way (e.g., by getting the short stick of redistributive pork), if there are institutional restrictions on their ability to articulate or translate their demands into meaningful electoral sanctioning, then nominal increases in transparency might not necessarily be associated with similar increases in accountability and, in turn, redistribution (Lindstedt and Naurin 2010). Thus, when considered independently of information access and mobilization capacities, it is not immediately clear what effect transparency should have on a given state's level of redistributive effort. This generates the following hypothesis:

> **Hypothesis 1:** There will not be a statistically significant relationship between transparency—considered unconditionally—and state welfare effort.

It is important to stress—and a point that has often been overlooked in previous studies—that the strength of the relationship between transparency, accountability, and governance outcomes, to the extent that there is one, may be determined nontrivially by who demands accountability. That there exists significant variation across social groups—specifically, social groups stratified by income (Solt 2008)—in the extent to which they monitor the behavior of their legislators and are likely to access information released by transparency initiatives (Bertot et al. 2006a; 2006b) suggests that the relationship between transparency and redistribution may be conditioned by these cleavages.

Many studies suggest that more affluent individuals are more likely to actively search for information (Organisation for Economic Co-operation and Development [OECD] 2016), actively monitor the behavior of their legislators (Livingstone and Markham 2008), and keep up with current events and the news (Druckman 2005) relative to less affluent citizens. Some evidence also suggests that these disparities in information access also hold, specifically, when we consider information released through transparency initiatives. For instance, Jaeger and Bertot (2010) present evidence that lower-income individuals, those without frequent access to the Internet, and those with lower levels of motivation and cognitive ability made little use of *e*-transparency websites. If it is the case that wealthier, more affluent individuals are the ones disproportionately consuming media (and by implication, monitoring the behavior of their representatives and being exposed to the information released by transparency interventions), then I expect that increased media market penetration—while previously hypothesized to amplify the equalizing distributional effects of transparency—will not necessarily be associated with commensurate increases in state welfare effort.[3] Stated differently, to the extent that media consumption is stratified by income, increased media market penetration—while ostensibly reducing costs to information access by increasing information availability—may only end up magnifying the political voice of the middle class at the expense of the lower class since consumption is disproportionately skewed toward more affluent segments of society. This generates the following hypothesis:

> **Hypothesis 2:** As media market penetration increases over time, increases in transparency will be unrelated to state welfare effort.

The ability to sanction legislators through the vote—a central component of accountability (McGee and Gaventa 2011; Gaventa and McGee 2013)—may promote more egalitarian distributional outcomes. For instance, Malesky, Schuler, and Tran (2012) present evidence that where the prospects of electoral sanctioning are trivial, transparency exerts weak effects on reducing corruption. However, the effect of voting on good governance outcomes may be conditioned by the type of citizen who actually votes. While the ability to vote within the United States is de jure not stratified across socioeconomic lines, it is de facto stratified along socioeconomic lines. As Verba, Schlozman, and Brady (1995) note, wealthier and more educated citizens are more likely to participate in politics because they have more time to, more money to dispense with, and cultivated a wider array of usable civic skills.

The evidence on policy responsiveness in the United States suggests that legislators are more responsive to the preferences of more affluent citizens, presumably because of the expectation among legislators that affluent constituencies are more likely to monitor and mobilize (Gilens and Page 2014; Hacker and Pierson 2010; Martin 2003). For instance, research in political economy suggests that wealthier jurisdictions with higher rates of voter turnout receive greater shares of federal expenditures (Martin 2003) and that middle-class and more affluent citizens—vis-a-vis their organizational resources—are well equipped to pressure legislators to pursue inegalitarian policies that benefit their groups to the detriment of the poor (Hacker and Pierson 2010; Jacobs and Soss 2010).

Describing the effect that these political inequalities can have on governance outcomes, Peisakhin and Pinto (2010, 262) note,

> It is far from given that greater transparency should always result in a drop in corruption levels. For instance, it seems reasonable that in highly hierarchical societies where the power gulf between the poor and government officials is very wide, greater transparency should benefit first and foremost the middle classes, while the underprivileged keep paying the same bribes.

More generally, these empirical findings are consistent with antipluralist (Lowi 1969; McConnell 1966; Schattschneider 1960) and power-relations (Piven 2006) accounts of rising inequality in the United States, which argue that ill-informed and fragmented publics (generally, more economically disadvantaged citizens) do not utilize legislative and political processes to advocate for their interests. Because lower-income citizens score lower on measures of political interest and engagement, members of more cohesive, economically advantaged pressure groups are able to exert significantly greater influence on politicians. When this occurs, policy outcomes may not be majoritarian in character and, instead, simply mirror the preferences of the middle and upper classes (Jacobs and Soss 2010; Soss and Jacobs 2009).

Accordingly, I anticipate that increases in transparency will do little to improve the social welfare of the most disadvantaged in society, specifically in states with greater participatory gulfs between the most and least disadvantaged (e.g., states where wealthier citizens have a higher likelihood of voter turnout relative to poorer citizens). This generates the following hypothesis:

> **Hypothesis 3:** As class bias in voter turnout increases over time, increases in transparency will be associated with decreases in state welfare effort.

To sum, these arguments represent a compelling theoretical challenge to the argument that transparency initiatives can, in isolation, increase the welfare of the most disadvantaged. When variation in information accessibility and mobilization capacities mirror existing status asymmetries between groups, then it seems reasonable to suspect that to the extent that they are responsive to constituent preferences, legislators will strategically allocate resources to reflect the interests of those exerting pressure through strong accountability mechanisms. Thus, while transparency interventions may have the effect of reducing absolute levels of resource concentration at the top of the income distribution (which from a normative equalitarian perspective may be desirable ends), the redistributive benefits of transparency interventions may accrue only to already advantaged segments of the population and do little to improve the welfare of the least well off. Only when transparency interventions are accompanied by policies that significantly reduce barriers to information access among the most disadvantaged and reduce the costs of mobilization do I expect transparency to exert significantly positive effects on levels of public welfare expenditure.

Data and Variables

The analysis that follows is based off a panel dataset with 23 time periods (1978–2000) for 47 contiguous states in the United States.[4] My dataset includes 1,081 observations in total. I first define my outcome variable, followed by brief descriptions of my explanatory variables, and then model specification. Details on descriptive statistics are available below in Table 1 and on data sources for these variables in Online Appendix A. I choose to analyze U.S. states because there is considerable variation in transparency, media market penetration, class bias, and welfare expenditures across the time period.

Table 1. Descriptive Statistics of Model Covariates.

Variable	*M*	*SD*	Minimum	Maximum	*N*
Transparency	0.469	0.194	0	1	1,104
Media Market Penetration	1.147	0.152	0.71	1.568	1,104
Class Bias	0.321	0.062	0.036	0.578	1,104
Gini Index	0.537	0.043	0.439	0.656	1,104
Unemployment	6.077	2.158	2.1	17.9	1,150
GSP (Adjusted)	1.475	1.755	0.095	13.77	1,104
Log (Population)	14.994	0.984	12.974	17.342	1,104
% Population > 65	0.121	0.022	0.026	0.185	1,150
% Population Black	0.185	0.135	0.005	0.713	1,150
Democratic Governor	0.535	0.496	0	1	1,104
Citizen Ideology	46.741	14.988	9.250	93.912	1,106
Legislator Ideology	56.438	19.877	6.514	95.58	1,106
Democratic Chamber Control	0.649	0.42	0	1	1,081
Divided Government	0.554	0.497	0	1	1,081

Note. GSP = gross state product.

Welfare Effort

I measure welfare effort as the log of per capita indigenous state expenditures on public welfare programs in constant 2000 dollars. Expenditures under this heading include cash assistance paid directly to needy persons under the categorical programs (e.g., Aid to Families with Dependent Children) and under any other welfare programs; vendor payments made directly to private purveyors for medical care, burials, and other commodities and services provided under welfare programs; and provision and operation by the government of welfare institutions including nursing homes not directly associated with a government hospital.

This measure of welfare expenditures— alternatively known as state welfare effort—is a relatively standard approach to assess how much a state prioritizes redistributive policies (Barrilleaux, Holbrook, and Langer 2002; Case, Rosen, and Hines 1993; Dye 1984; Jennings 1979). I log the state welfare effort variable for two reasons: first, model residuals are skewed-right when I regress the nontransformed welfare effort variable on the set of covariates, and second, to ease aid of interpretation (Wooldridge 2015).[5]

Transparency

My measure of transparency is derived from Alt, Lassen, and Rose's (2006) index of fiscal budget transparency. Alt, Lassen, and Rose measure budget transparency as a continuous index based on nine items reflecting different aspects of budget transparency. The nine items included in this measure are as follows:

1. Budget is reported using generally accepted accounting principles (GAAP).
2. Multiyear expenditure forecasts are used.
3. Budget cycle is annual.
4. Revenue forecasts are binding.
5. Legislative branch has or shares responsibility for revenue forecasts.
6. All appropriations are included in a single bill.
7. Appropriation bills are written by nonpartisan staff.
8. Open-ended appropriations are prohibited.
9. Budget requires published performance measures.

The transparency score is the ratio of items coded yes to the total items answered for a given year and, thus, ranges between 0 and 1. Alt and Lowry (2010, 386) note that this operationalization of transparency signals information "about fiscal policy or the cost of public spending (or even policy choice), rather than open access to the policy process or personal character or past records of candidates." As my primary outcome of interest is state redistributive spending, this particular operationalization of transparency is arguably the most relevant.

Media Market Penetration

Media market penetration is a variable that aggregates newspaper circulation and cable reception by state and year derived from Alt and Lowry (2010). This measure aggregates newspaper and cable consumption into a single index to adjust for the fact that over the time period covered, newspaper subscriptions declined and cable consumption increased. More details on the construction of this variable are available in Online Appendix A.

Class Bias

The measure of class bias in voter participation is a relatively new measure derived from Franko, Kelly, and Witko (2016) and captures disproportionate participation rates across income groups (Blakely, Kennedy, and Kawachi 2001; Wichowsky 2012), providing an empirical measure of how much richer citizens participate in elections relative to poor citizens within their state. Specifically, the class bias variable indicates the ratio of expected voter turnout of the top 10% of income earners in a state relative to the bottom 10% of income earners in a state. The variable is bound between −1 and +1, with negative values meaning that low-income individuals are more likely to vote than high-income individuals, 0 meaning that low- and high-income individuals have an equal probability of voting, and positive values meaning that the high-income individuals vote at a higher rate than lower-income individuals. A value of 0.32 (the average of the measure across all years and states), for example, indicates that the richest 10% of citizens in a state are 32% more likely to vote than the poorest 10% of citizens in a state. More details on the construction of this variable are available in Online Appendix A.

Other Explanatory Variables

I include controls for the level of economic inequality within a state (Gini index), a state's unemployment rate, standardized GSP in real 2000 dollars, population size, the proportion of the state's population that is black, the proportion of the state's population that is over 65 years old, which political party controls the governorship, state citizen ideology, state legislator ideology, which party controls the legislative chamber, and whether there is divided government in the legislative chambers of a state. More detailed descriptions of these variables are available in Online Appendix A.

Econometric Models

I present the result of a series of "between-within" panel models developed by Brandon Bartels (2008) and discussed in recent political science scholarship by Allison (2009) and Bell and Jones (2015). I use the between-within estimator (hereafter, BWE) to escape some of the well-documented bias and efficiency limitations of conventional fixed effects and random effects approaches to modeling panel data. For instance, while fixed effects estimation is able to control for the problem of unobserved heterogeneity at the cluster level, fixed effects estimation generally generates inefficient estimation of slow-evolving variables and, by necessity, excludes time-invariant variables from the model altogether (Beck 2001; Plumper and Troeger 2007). Alternatively, while random effects estimation allows the researcher to include time-invariant variables within the model, random effects estimation makes the potentially problematic assumption that longitudinal and cross-sectional effects are statistically equivalent, raising the issue of cluster confounding (Skrondal and Rabe-Hesketh 2004; Zorn 2001). If we have reason to suspect that longitudinal and cross-sectional effects of a particular variable on an outcome are distinct and desire efficient estimation of slowly evolving variables, then conventional approaches to panel data estimation—specifically, pooled ordinary least squares (OLS), fixed effects, and random effects—may generate misleading, biased coefficients.[6]

The BWE is able to overcome these limitations, accounting for unobserved cluster-level heterogeneity, producing efficient estimation of slowly evolving and time-invariant variables, and explicitly decomposing variation at both cross-sectional and longitudinal levels, which allows for the explicit modeling of cross-sectional and longitudinal effects.[7] To address the problem of cluster confounding, the first step in between-within estimation involves transforming time-variant variables in the model (Skrondal and Rabe-Hesketh 2004; Zorn 2001). First, one computes a cluster mean $\overline{X}_i$ across all time periods, j, where i indexes the relevant geographic unit (state). In the case of the article's model, the cluster mean $\overline{X}_i$ is the average level of transparency in a particular state across all 22 years. To generate the within-cluster transformation, one subtracts the longitudinal average of the covariate within i from each observed instance of the covariate within each time period j, producing a time-variant measure, X_{ij}. This represents the longitudinal within-cluster deviation from the cluster mean and is given by equation 1:

Table 2. Correlation Matrix.

	(1)			
	Welfare/capita	Transparency	Media market penetration	Class bias
Welfare/Capita	1	—	—	—
Transparency	0.151**	1	—	—
Media Market Penetration	0.445**	0.138**	1	—
Class Bias	−0.046	−0.0590	−0.0326	1

*p < .10. **p < .05.

$$X_{ij}^{W} = X_{ij} - \overline{X}_{i}. \tag{1}$$

This group mean-centering decomposes the variation at the longitudinal and cross-sectional levels and, in so doing, satisfies the controversial assumption that there is no correlation between the group mean centered within transformation and the cross-sectional random effect. The next step in between-within estimation involves specifying a random intercept model that includes both the within- and between-cluster transformations of the relevant covariates. This generates the following simplified model:

$$WelfareEffort_{ij} = \beta_0 + \beta_1\left(x_{ij} - \overline{x_j}\right) + \beta_2\overline{x}_j + \beta_3 z_j + \left(\mu_j + \epsilon_{ij}\right). \tag{2}$$

In this model, β_0 is the intercept term, β_1 represents the within-cluster effect of the covariates for a given state on state welfare effort, and β_2 represents the between-cluster effect of the covariates on state welfare effort. Conceptually, one can consider the within-effect to be analogous to a longitudinal effect and the between-effect to be analogous to a cross-sectional effect. Thus, the within-coefficient expresses for a given state the effect of longitudinal *changes* in transparency on state welfare effort holding the state's level of transparency constant, whereas the between-coefficient expresses the association of cross-sectional *levels* of transparency with state welfare effort. As I am most interested in the effect of changes in transparency on state welfare effort, I primarily discuss the within-effects and their respective interactions in the following.

Results

I first examine a series of bivariate relationships between the main variables of interest as shown in Table 2. Welfare expenditures per capita correlate positively with both the level of budget transparency and media market penetration, as the conventional wisdom suggests.

Indeed, a naïve comparison of the general trends in transparency and public welfare expenditures across time, as depicted in Figure 1, suggests that transparency and

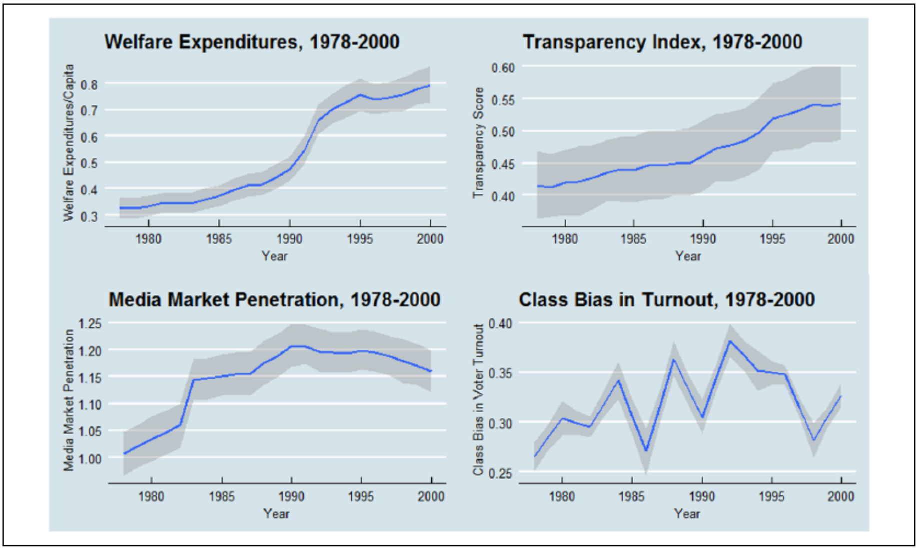

Figure 1. Trends in primary variables, 1978–2000.

welfare expenditures covary precisely in the direction predicted by the conventional wisdom—as transparency increases, so too do state welfare expenditures per capita. Curiously, class bias in voter participation is not significantly correlated with public welfare expenditures per capita, a finding that, at face value, seems inconsistent with previous research on the relationship between redistributive preferences and relative class power (Bobo 1991; Guillaud 2013; Piketty 1995) but which may simply reflect the measure's seasonality and interpolation over nonelection years.

Assessing the Noninteractive Effect of Transparency

To begin a more probative analysis, Table 3 presents the regression estimates for the log of per capita public welfare expenditures at the state level obtained by the BWE with state-clustered robust standard errors included to partially address concerns with heteroskedasticity and serial correlation (Hoechle 2007) as well as year fixed effects and state random effects.[8] The table includes a series of stepwise regressions in which the interactions specified in Hypotheses 2 and 3 are introduced sequentially to the baseline noninteractive Model 1.

It is important to reiterate that the within-coefficients of the covariates produced by the BWE represent the within-state deviations in the covariate over time and, thus, that the regression coefficients presented in Models 1 to 4 do not represent the effect of a one-unit change in the raw measurement of the covariate on the outcome but rather the effect of a one-unit change in the group mean centered transformation of the covariate. Thus, the within-coefficients represent the effect of deviations in a covariate within a state over time on state welfare effort; the between-coefficients represent the cross-sectional effect of

Table 3. BWE Regressions of Public Welfare Expenditures on Transparency and Controls, Baseline Specification, 47 U.S. States, 1978–2000, Balanced Panel.

	Model 1	Model 2	Model 3	Model 4
$Transparency_{wi}$	−0.090	−0.088	−0.097	−0.079
	(−1.29)	(−1.26)	(−1.40)	(−1.10)
$Transparency_{bw}$	0.025	−2.070*	−0.279	1.987
	(0.22)	(−1.77)	(−0.25)	(0.21)
Media Market $Penetration_{wi}$	0.052	0.053	0.055	0.048
	(0.54)	(0.56)	(0.58)	(0.50)
Media Market $Penetration_{bw}$	0.216	−0.736	0.217	2.068
	(0.95)	(−1.28)	(0.95)	(0.42)
Class $Bias_{wi}$	−0.340**	−0.340**	−0.333**	−0.316**
	(−2.88)	(−2.88)	(−2.83)	(−2.57)
Class $Bias_{bw}$	−2.576**	−2.768**	−3.082*	6.238
	(−4.13)	(−4.49)	(−1.65)	(0.36)
$Gini_{wi}$	0.096	0.105	0.089	0.114
	(0.26)	(0.29)	(0.24)	(0.31)
$Gini_{bw}$	−4.348**	−4.980**	−4.325**	−4.995**
	(−3.55)	(−4.02)	(−3.52)	(−4.02)
GSP_{wi}	−0.054**	−0.054**	−0.053**	−0.053**
	(−3.97)	(−3.99)	(−3.92)	(−3.89)
GSP_{bw}	0.077**	0.086**	0.077**	0.089**
	(2.90)	(3.30)	(2.92)	(3.40)
$Unemployment_{wi}$	−0.002	−0.002	−0.003	−0.002
	(−0.56)	(−0.55)	(−0.67)	(−0.60)
$Unemployment_{bw}$	0.003	0.004	0.004	0.004
	(0.18)	(0.22)	(0.21)	(0.23)
Log $(Population)_{wi}$	0.445**	0.442**	0.446**	0.435**
	(5.70)	(5.63)	(5.73)	(5.55)
Log $(Population)_{bw}$	−0.057	−0.066	−0.058	−0.071
	(−1.27)	(−1.51)	(−1.29)	(−1.62)
% Population > 65_{wi}	2.075*	2.089*	2.083*	2.152*
	(1.77)	(1.78)	(1.78)	(1.83)
% Population > 65_{bw}	2.601**	2.965**	2.426**	2.628**
	(2.32)	(2.68)	(1.97)	(2.21)
% $Black_{wi}$	−0.076	−0.067	−0.100	−0.099
	(−0.43)	(−0.38)	(−0.57)	(−0.56)
% $Black_{bw}$	−0.000	0.058	−0.020	0.028
	(−0.00)	(0.36)	(−0.12)	(0.17)
Citizen $Ideology_{wi}$	0.003**	0.003**	0.003**	0.003**
	(2.93)	(2.89)	(2.86)	(2.77)
Citizen $Ideology_{bw}$	0.009**	0.007	0.009**	0.007
	(2.06)	(1.64)	(2.04)	(1.56)
Legislature $Ideology_{wi}$	0.003**	0.003**	0.003**	0.003**
	(3.35)	(3.37)	(3.19)	(3.18)

(continued)

Table 3. (continued)

	Model 1	Model 2	Model 3	Model 4
Legislator Ideology$_{bw}$	0.001	0.003	0.002	0.004
	(0.18)	(0.49)	(0.23)	(0.57)
Democratic Governor$_{wi}$	−0.046*	−0.047*	−0.044*	−0.045*
	(–1.83)	(–1.87)	(–1.76)	(–1.78)
Democratic Governor$_{bw}$	0.087	0.052	0.070	0.019
	(0.44)	(0.27)	(0.35)	(0.10)
Democrat Chamber Control$_{wi}$	0.023	0.023	0.026	0.026
	(0.86)	(0.85)	(0.94)	(0.96)
Democrat Chamber Control$_{bw}$	0.101	0.015	0.095	0.008
	(0.56)	(0.08)	(0.52)	(0.04)
Divided Government$_{wi}$	−0.004	−0.004	−0.004	−0.002
	(–0.35)	(–0.34)	(–0.33)	(–0.21)
Divided Government$_{bw}$	0.069	0.018	0.063	−0.009
	(0.68)	(0.18)	(0.60)	(–0.09)
Transparency$_{wi}$ × Media Market	—	0.455	—	0.769
Penetration$_{wi}$	—	(0.45)	—	(0.71)
Transparency$_{bw}$ × Media Market	—	1.816*	—	−2.527
Pentration$_{bw}$	—	(1.80)	—	(–0.30)
Transparency$_{wi}$ × Class Bias$_{wi}$	—	—	−2.841**	−3.028**
	—	—	(–2.20)	(–2.20)
Transparency$_{bw}$ × Class Bias$_{bw}$	—	—	0.972	−13.780
	—	—	(0.28)	(–0.46)
Media Market Penetration$_{wi}$ ×	—	—	—	−1.156
Class Bias$_{wi}$	—	—	—	(–0.76)
Media Market Penetration$_{bw}$ ×	—	—	—	−9.319
Class Bias$_{bw}$	—	—	—	(–0.59)
Transparency$_{wi}$ × Media Market	—	—	—	−13.223
Penetration$_{wi}$ × Class Bias$_{wi}$	—	—	—	(–0.84)
Transparency$_{bw}$ × Media Market	—	—	—	14.605
Penetration$_{bw}$ × Class Bias$_{bw}$	—	—	—	(0.54)
Constant	1.487	3.123**	1.680	0.575
	(1.24)	(2.11)	(1.26)	(0.10)
σ_μ Constant	0.124**	0.120**	0.124**	0.119**
	(9.17)	(9.14)	(9.17)	(9.13)
σ_ε Constant	0.142**	0.142**	0.142**	0.141**
	(45.48)	(45.48)	(45.48)	(45.47)
Years Covered	1978–2000	1978–2000	1978–2000	1978–2000
States	47	47	47	47
N	1,081	1,081	1,081	1,081

Note. Cluster robust standard errors in parentheses. Coefficient estimates for state and year indicators suppressed. BWE = between-within estimator; GSP = gross state product.
*$p < .10$. **$p < .05$.

levels of a variable on state welfare effort. In addition, given that the dependent variable is logged, raw coefficients for noninteracted independent variables are interpretable as percentage changes in state welfare effort.

Results provide support for my three hypotheses. In Model 1, transparency—both its longitudinal effect on and cross-sectional associations—with state welfare effort does not achieve statistical significance, consistent with Hypothesis 1. Initially, it does not appear that changes in a state's level of transparency over time or a given state's level of transparency alone are associated with redistributive effort. From Model 1, I additionally note that while changes to or levels of media market penetration do not exert a significant effect on state welfare effort, class bias influences state welfare effort both longitudinally and cross-sectionally in the direction anticipated—in states in which the wealthy are significantly more likely to turnout for elections than the poor, state welfare effort decreases. As an illustration of this point, a one-unit increase in the within-class bias measure (substantively, when a state over time goes from the richest income decile and the poorest income decile voting at equivalent rates to the poorest income decile not voting at all), all else equal, state welfare effort significantly declines by approximately 34%, a finding consistent with antipluralist accounts of rising income inequality.

Assessing the Conditional Effects of Transparency

However, the central logic of my hypotheses is conditional. I argue that whether transparency increases a state's redistributive effort hinges not so much on variation in media market penetration because information is disproportionately consumed by those with higher incomes but rather on the ratio of voter turnout between the richest and poorest citizens. Thus, Models 2 to 4 assess these hypothesized conditional effects.

From Model 2, I find that while there is not a significant within-effect for the interaction of transparency and media market penetration, there is a marginally significant and positive between-effect of the interaction. Tentatively, this suggests that more transparent states with larger media markets have higher levels of per capita welfare spending. However, the statistical significance of this interaction term is an artifact of model extrapolation given the lack of common support in the estimation of marginal effects (Hainmueller, Mummolo, and Xu 2018). When I only examine the interaction effect over observed, within-sample values of the between-transformation of media market penetration, there are not statistically significant differences in the effect of transparency at sample minimum and maximum values.

From Model 3, I find that the longitudinal effect of transparency on state welfare effort is significantly moderated by changes in class bias within states—as the participatory gulf between the richest income decile and poorest income decile in a given state widens over time (e.g., when the wealthy become more likely to vote than the poor over time within a given state), increases in transparency are negatively associated with state welfare effort, all else equal. Perhaps the best way to understand the interaction effect in Table 3—and specifically, in Model 3—is by considering a few extreme cases.

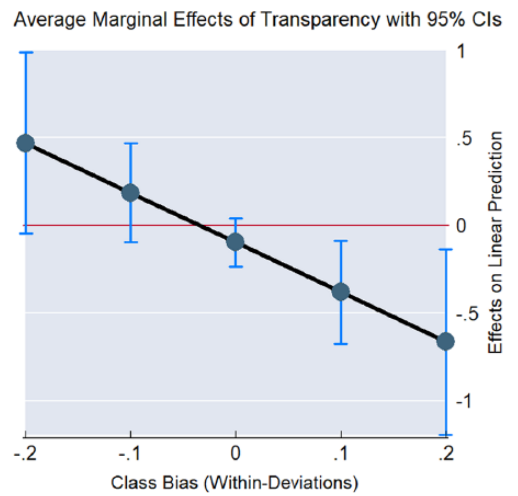

Figure 2. Average marginal effect of changes in transparency on state welfare effort by class bias.

In Figure 2, I plot the marginal effect of fiscal transparency on state welfare effort at varying levels of within-class bias deviations. The plot shows that when class bias in voter turnout does not change within a state over time (when the class bias deviation score equals 0), a one-unit increase in transparency (the extreme instance when a state changes from being completely nontransparent to completely transparent over the 22-year time period) has substantively negligible effects on state redistributive effort. However, in states in which the difference in voter turnout between the top and bottom income deciles increases by 10% over the 22-year time period, a one-unit increase in the transparency deviation score reduces the percentage of welfare expenditures per capita by a significant 33.62% (p value = 0.023). More dramatically, in states in which there is an over-time 20% increase in the difference in voter turnout between the top and bottom income deciles, a one-unit change in the transparency deviation score reduces the percentage of welfare expenditures per capita by an even larger and statistically significant 57.1% (p value = 0.03). Moreover, the statistical significance of this effect holds when I restrict the interaction to regions of common support for both constitutive terms. This finding is consistent with the argument of Hypothesis 3: in states in which wealthier citizens become even more likely to vote than poorer citizens, increases in transparency are negatively associated with state welfare effort.

Robustness Tests

In this section, I present a series of robustness tests for the main analysis. I first assess the robustness of my results to conventional fixed and random effects estimators. I then proceed to note the presence of cluster confounding that arises in these models as a means to justify the use of the BWE. Finally, I conclude by briefly discussing results from a series of alternative model specifications that use different dependent variables and specify other interactions.

I re-estimate Models 1 to 4 using the fixed effects estimator (Table 4) and random effects estimator (Table 5). Results for these conventional panel specifications are given below.

It is worth noting that my substantive results—specifically, the unconditional effect of transparency and the conditional effects of transparency on state welfare effort—are generally robust across model specifications, though the BWE allows for a more nuanced examination of these conditional relationships since it decomposes substantively meaningful within and between variance. Across all noninteractive models (Table 4, Model 5; Table 5, Model 9), transparency by itself does not exert a statistically significant effect on state welfare effort. However, when I investigate the interaction between transparency and class bias in voter turnout, fixed and random effects models generate contradictory results. The fixed effect model estimates (Table 4; Model 7) are consistent with the results generated from the BWE model, finding a significant, negative interaction between class bias and transparency on state welfare effort. Conversely, the random effects model (Table 5; Model 11) does not find a statistically significant effect for this interaction.

One explanation for this particular inconsistency in model estimates may be attributable to the issue of cluster confounding. To diagnose the presence of cluster confounding in the data, I follow Bartels (2008) and re-estimate Model 1 from the BWE with an alternative specification. Instead of including the within-transformations of the time-variant covariates in addition to the between-transformations, I substitute the within-transformations of the covariates with the nontransformed versions of these covariates. Importantly, this alters this interpretation of the between-coefficients, which now represent differences in the within and between estimates. If the coefficients on the revised between covariates achieve statistical significance, this indicates the presence of cluster confounding in the data. I present results of this diagnostic test in Online Appendix C. I find evidence for significant cluster confounding specifically within the class bias, Gini index, adjusted GSP, and population coefficients. The presence of such cluster confounding justifies the utilization of the BWE over pooled OLS and conventional random effects approaches, and may account for the inconsistency between the RE and BWE results.

Finally, to assess the robustness of my main results to additional model specifications, I tried fitting Models 1 to 4 without controlling for a state's level of economic inequality—as measured by the Gini index in the original model specifications—to investigate whether the coefficient on the Gini index absorbed variation that could actually be attributed to variation in class bias in voter turnout; the results were substantively the same. I tried interacting the Gini index and an additional measure of income inequality—income share of the top 10% (Piketty and Saez 2003)—with transparency to investigate whether general economic inequality conditioned the relationship between transparency and state welfare effort. None of the longitudinal

Table 4. Fixed Effects Regressions of Public Welfare Expenditures on Transparency and Controls, Baseline Specification, 47 U.S. States, 1978–2000, Balanced Panel.

	Model 5	Model 6	Model 7	Model 8
Transparency	−0.090	1.501**	0.158	0.493
	(−1.27)	(5.02)	(1.00)	(0.48)
Media Market Penetration	0.052	0.681**	0.039	−0.055
	(0.53)	(4.56)	(0.40)	(−0.11)
Class Bias	−0.340**	−0.325**	0.060	−3.132*
	(−2.83)	(−2.74)	(0.23)	(−1.66)
Gini	0.096	0.201	0.125	0.191
	(0.26)	(0.54)	(0.33)	(0.52)
GSP	−0.054**	−0.048**	−0.055**	−0.046**
	(−3.90)	(−3.54)	(−3.96)	(−3.36)
Unemployment Rate	−0.002	−0.001	−0.002	−0.002
	(−0.55)	(−0.29)	(−0.59)	(−0.40)
Log (Population)	0.445**	0.433**	0.459**	0.432**
	(5.60)	(5.52)	(5.75)	(5.47)
% Population > 65	2.075*	1.743	2.043*	1.674
	(1.73)	(1.48)	(1.71)	(1.41)
% Black	−0.076	0.081	−0.067	0.079
	(−0.43)	(0.46)	(−0.38)	(0.44)
Citizen Ideology	0.003**	0.002**	0.003**	0.002**
	(2.88)	(2.31)	(2.83)	(2.28)
Legislature Ideology	0.003**	0.004**	0.003**	0.004**
	(3.29)	(4.20)	(3.23)	(4.15)
Democratic Governor	−0.046*	−0.072**	−0.045*	−0.071**
	(−1.80)	(−2.83)	(−1.76)	(−2.77)
Democrat Chamber Control	0.023	0.005	0.024	0.005
	(0.85)	(0.17)	(0.88)	(0.18)
Divided Government	−0.004	−0.008	−0.002	−0.008
	(−0.34)	(−0.72)	(−0.22)	(−0.78)
Transparency × Media Market	—	−1.351**	—	−0.440
Penetration	—	(−5.47)	—	(−0.49)
Transparency × Class Bias	—	—	−0.781*	3.464
	—	—	(−1.74)	(1.04)
Media Market Penetration ×	—	—	—	2.527
Class Bias	—	—	—	(1.51)
Transparency × Media Market	—	—	—	−3.097
Penetration × Class Bias	—	—	—	(−1.07)
Constant	−8.216**	−8.806**	−8.539**	−7.963**
	(−6.95)	(−7.53)	(−7.15)	(−6.09)
R^2	.871	.875	.871	.875
N	1,081.000	1,081.000	1,081.000	1,081.000

Note. Cluster robust standard errors in parentheses. Coefficient estimates for state and year indicators suppressed. GSP = gross state product.
$^*p < .10$. $^{**}p < .05$.

Table 5. Random Effects Regressions of Public Welfare Expenditures on Transparency and Controls, Baseline Specification, 47 U.S. States, 1978–2000, Balanced Panel.

	Model 9	Model 10	Model 11	Model 12
Transparency	−0.054	1.351**	0.090	0.374
	(−0.85)	(4.62)	(0.57)	(0.36)
Media Market Penetration	0.249**	0.832**	0.241**	0.087
	(2.89)	(5.70)	(2.79)	(0.16)
Class Bias	−0.412**	−0.396**	−0.178	−3.385*
	(−3.42)	(−3.32)	(−0.68)	(−1.77)
Gini	−0.298	−0.224	−0.268	−0.198
	(−0.82)	(−0.62)	(−0.74)	(−0.55)
GSP	−0.025**	−0.021*	−0.026**	−0.021*
	(−1.98)	(−1.70)	(−2.06)	(−1.66)
Unemployment Rate	−0.000	0.000	−0.001	−0.000
	(−0.10)	(0.08)	(−0.13)	(−0.03)
Log (Population)	0.122**	0.116**	0.124**	0.119**
	(4.29)	(4.14)	(4.33)	(4.13)
% Population > 65	2.020**	1.727**	2.026**	1.650*
	(2.39)	(2.06)	(2.38)	(1.92)
% Black	−0.054	−0.008	−0.052	−0.013
	(−0.42)	(−0.06)	(−0.39)	(−0.10)
Citizen Ideology	0.005**	0.005**	0.005**	0.005**
	(6.04)	(5.71)	(5.99)	(5.50)
Legislator Ideology	0.004**	0.004**	0.003**	0.004**
	(4.01)	(4.73)	(3.97)	(4.70)
Democratic Governor	−0.068**	−0.089**	−0.067**	−0.088**
	(−2.65)	(−3.47)	(−2.61)	(−3.42)
Democratic Chamber Control	0.009	−0.004	0.009	−0.004
	(0.32)	(−0.13)	(0.34)	(−0.14)
Divided Government	−0.010	−0.013	−0.009	−0.014
	(−0.86)	(−1.18)	(−0.79)	(−1.27)
Transparency × Media Market	—	−1.199**	—	−0.372
Penetration	—	(−4.92)	—	(−0.41)
Transparency ×Class Bias	—	—	−0.455	3.536
	—	—	(−1.00)	(1.04)
Media Market Penetration ×	—	—	—	2.601
Class Bias	—	—	—	(1.52)
Transparency × Media Market	—	—	—	−2.982
Penetration ×Class Bias		—	—	(−1.02)
Constant	−3.559**	−4.169**	−3.668**	−3.352**
	(−7.35)	(−8.43)	(−7.41)	(−4.34)
R^2	—	—	—	—
N	1,081.000	1,081.000	1,081.000	1,081.000

Note. Cluster robust standard errors in parentheses. Coefficient estimates for state and year indicators suppressed. GSP = gross state product.
*p < .10. **p < .05.

interactions were statistically significant, suggesting that changes in class-based inequalities in political power over time—and not simply changes in generalized economic inequality—is what specifically depresses state welfare effort. Finally, I assessed the sensitivity of model results to an alternate operationalization of the dependent variable, examining total state welfare expenditures as a proportion of total state personal income as suggested by Albritton (1990). Results are substantively the same.

It is worth reiterating that across all of these model specifications, longitudinal increases in state fiscal transparency and cross-sectional differences in levels of transparency fail to exert a statistically significant, positive effect on state welfare effort and that in a majority of the specifications, the interaction of transparency with class bias exerts a negative effect on state welfare effort.

Conclusion

There is still considerable academic and policy debate about best practices for implementing transparency initiatives. In this article, I used a dataset of U.S. states over 22 years to estimate the relationship between transparency and state welfare effort and to explore potential factors that have been theorized to condition this relationship—media market penetration and heterogeneous voter participation by income level. My empirical results suggest that (1) transparency—considered unconditionally—does not significantly predict state welfare effort, (2) the effect of transparency on state welfare effort is not moderated by media market penetration, and (3) the effect of transparency on state welfare effort is conditional on the relative sanctioning propensities of the wealthiest and poorest citizens in a state. In states in which wealthy citizens become significantly more likely to vote than poorer citizens over time, increases in transparency considerably depress per capita welfare expenditures.

Although these findings are interesting, future work should investigate the effects of alternative measures of transparency on redistributive behavior and other governance outcomes within the U.S. states. In addition, it would be worthwhile to expand the present analysis to more recent years to examine the effect that Internet access has on redistributive outcomes given questions about the extent to which Internet access reduces costs to information access among the poor and given that the rich and poor search for information online in different ways. To sum, my findings suggest that if one of the overarching goals of transparency interventions is to equalize distributional outcomes, the political voice of the poorest should be magnified. For transparency interventions to be effective at enhancing the welfare of the least well off, it is necessary to ensure that both the ability to access the information released by transparency interventions and the ability to mobilize on behalf of that information are not privileges disproportionately enjoyed by the affluent.

Author's Note

The author will readily share all data and coding information with those researchers who wish to replicate the study. The author is responsible for any errors of omission or commission as well as the interpretations and conclusions made in this article.

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Supplemental Material

Supplementary material for this article is available online.

Notes

1. Most political-economy models of redistributive preferences assume that individuals will support redistributive programs over other, alternative programs when their net income increases under the former as an extension of one's economic self-interest (Guillaud 2013; Meltzer and Richard 1981).
2. Tisné (2010) defines accountability as, "the process of holding actors responsible for their actions. More specifically, it is the concept that individuals, agencies and organizations (public, private and civil society) are held responsible for executing their powers according to a certain standard (whether set mutually or not)."
3. It is worth noting that while it is the case that some interest groups may have a genuine desire to advocate for the preferences of the least well off, historically, most interest groups are not concerned with disadvantaged subpopulations and when they are, they often inadvertently downplay the issues of importance to members of disadvantaged communities (Strolovitch 2006).
4. Excluded from analysis are Alaska, Hawaii, and Nebraska. These states are excluded from analysis because data on relevant independent variables—most notably, the transparency score and class bias measures—were not available for these states for all years in the dataset.
5. In a series of nonreported models, I re-estimate the models from the article's main analysis with a different measure of state welfare effort—total indigenous public welfare expenditures as a percentage of total personal income as suggested by Albritton (1990). Results are substantively similar for this alternate measure.
6. More detail on the advantages of the between-within estimator (BWE) approach is given in Online Appendix B.
7. The STATA command *clustergen* was used to produce within- and between-transformations (Bartels 2008).
8. Year fixed effects and state random effects are suppressed from the main-text due to space constraints.

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