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Distributing Discipline:

Race, Politics, and Punishment at the Frontlines of Welfare Reform*

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ABSTRACT

Numerous studies have confirmed that race plays an important role in shaping public preferences toward both redistribution and punishment. Likewise, studies suggest that punitive policy tools tend to be adopted by state governments in a pattern that tracks with the racial composition of state populations. Such evidence testifies to the enduring power of race in American politics, yet it has limited value for understanding how disciplinary policies get applied to individuals in implementation settings. To illuminate the relationship between race and the application of punitive policy tools, we analyze sanction patterns in the TANF program. Drawing on a model of racial classification and policy choice, we test four hypotheses regarding client race and sanctioning. Our study does not support a simple story in which racial minorities are always more likely to be targeted for discipline. Rather, we find the impact of race to be contingent on local politics, administrative decentralization, and other client characteristics.

INTRODUCTION

Punishment is a core technology of governance, central to the modern state. It is usually most visible in criminal justice systems, where authorities penalize violations of law by imposing fines, stays of imprisonment, and even sentences of death. Yet as a tool of governance, punishment has far broader application. Across diverse societal domains, disciplinary systems establish behavioral incentives (Ehrlich 1996) and shape the ways individuals understand and regulate their own conduct (Foucault 1979).

In recent years, punishment has taken on greater significance for students of American politics, as governments at all levels in the United States have come to rely more heavily on punitive tools. This is evidenced by the increasing harshness of criminal justice policies, and the impact that these policies have had on the use of imprisonment. This has contributed to a level of mass incarceration that is unprecedented in American history and unrivaled by other nations (Western 2006). Yet, this custodial turn has not been isolated to criminal justice policies. Rather, it has coincided with a broader shift toward paternalist principles in U.S. social policy (Starobin 1998). Policies for low-income populations, in particular, have been redesigned around the idea that the state has a legitimate interest in ensuring that socially marginal groups practice appropriate behavior. Today, public aid programs are more directive in setting behavioral expectations, supervisory in monitoring compliance, and punitive in responding to infractions (Wacquant 2001).

As disciplinary systems have come to play a more central role in poverty governance, the study of administrative practice has taken on greater urgency (Mead 2004). Because public benefits are now more contingent on recipient behavior, policy implementers have had to reorganize their operations around the tasks of identifying and penalizing rule violators. Indeed, it is inevitably at the frontlines of policy implementation – in concrete organizations such as courts, schools, prisons and welfare agencies – that the politics of punishment culminates in actual decisions to impose penalties on some but not others.

To understand discipline as an element of governance, then, political scientists must do more than explain why mass publics or political jurisdictions vary in the extent to which they embrace punitive policy designs. We must investigate how such designs are actually put into practice and, in so doing, clarify the social, political, economic, and organizational forces that shape decisions to discipline. It is here, in the operation of administrative discretion, that the state's capacity to punish collides with the core question of distributive politics posed by Harold Lasswell (1936): "who gets what, when, how?"

This article presents an analysis of discipline at the frontlines of welfare reform, focusing on how patterns of punishment are affected by the interplay of race and politics. In the United States, there is a long history of social-policy tools being applied to racial groups in unequal ways (Lieberman 1998; Katznelson 2005). Today, however, the age of Jim Crow is a distant memory; *de jure* discrimination is banned; and norms of racial equality are widely embraced (Mendelberg 2001). In this post-civil-rights context, how and why does race matter for disciplinary action in local welfare offices? Under what conditions do client racial characteristics become significant influences on decisions to impose penalties?

In recent years, numerous studies have confirmed that racial attitudes remain an important determinant of public preferences regarding both social programs and systems of punishment (Gilens 1999; Peffley, Hurwitz, and Sniderman 1997). Likewise, studies suggest that punitive policy tools tend to be adopted in the American states in a pattern that tracks with the racial composition of state populations (Fellows and Rowe 2004; Hero 1998; Jacobs and Carmichael 2002; Soss et al. 2001). Such evidence testifies to the enduring power of race in American politics, yet it has limited value for understanding how disciplinary policies actually get applied to individuals in implementation settings.

To illuminate the relationship between race and the application of punitive policy tools, we analyze sanction patterns in the Temporary Assistance for Needy Families (TANF) program. When federal lawmakers passed welfare reform in 1996, they required states to develop procedures for sanctioning TANF clients. Sanctions are penalties that suspend all or part of a family's benefits for failing to comply with a program requirement. Although they have existed as a policy tool in the welfare domain for many years, they have taken on a more central programmatic role since 1996, as clients have confronted stricter work obligations, narrower exemption criteria, expanded behavioral requirements, and stronger penalties for noncompliance (Hasenfeld, Ghose and Larson 2004). Sanctions impose highly visible, direct material consequences for behavioral infractions. They provide the primary mode of disciplinary action in state TANF programs and, as such, are among the most potent tools that case managers have for influencing client behaviors and pursuing performance goals.

Prior research has shown that states with higher proportions of nonwhite recipients have been significantly more likely to adopt the strictest sanction policies under welfare reform (Soss et al. 2001; Fellowes and Rowe 2004). Government records suggest that such racial dynamics may characterize the sanction implementation process as well. Using data from the U.S. Government Accountability Office (GAO), one can examine the relationship between the racial composition of the welfare caseload and sanction usage in the states that have adopted the strictest sanction policies. Doing so, we find a strong relationship between the percentage of TANF recipients who are black or Latino and the extent to which states impose sanctions (see Figure 1). Figure 1 suggests a relationship between race and sanctioning outcomes, yet these aggregate state-level data are extremely limited. They do not allow us to determine if states are actually sanctioning black and Latino recipients more often than whites, and they tell us nothing about how, when, and why client race might matter for the use of penalties under welfare reform. To pursue such issues, one must examine individual-level data within a research design that maximizes our ability to understand if and why racial disparities exist in sanction implementation.

[Figure 1]

Drawing on a general model of racial classification and policy choice, we begin our investigation by developing the logic underlying four hypotheses regarding client race and sanctioning in the TANF program. We then proceed to test our hypotheses by using two very different, yet complimentary datasets. We begin with an analysis of longitudinal data on TANF clients from the state of Florida's Welfare Transition (WT) program, a welfare-to-work program funded by the TANF block grant. As the centerpiece of federal welfare reform in 1996, state TANF programs have been at the heart of the recent turn toward paternalist social policy. The Florida WT program, in turn, has been at the leading edge of state efforts to use sanctions as a tool to motivate behavioral compliance and change (see Figure 1). We supplement this analysis with a cross-sectional analysis of TANF sanctioning using a national dataset, which allows us to provide a partial replication of our results from Florida. Our findings are consistent across both analyses and suggest that race has played an important role in sanction implementation under welfare reform. However, our results do not support a simple story in which racial minorities are always more likely to be targeted for discipline. Rather, our findings support a more nuanced account in which the impact of race is contingent on local politics and administrative decentralization, as well as other client characteristics.

POLITICS, CHOICE, AND THE IMPLEMENTATION OF TANF SANCTIONS

A variety of studies have used administrative or survey data to analyze the client characteristics that correlate with a higher likelihood of being sanctioned. Their findings converge on the conclusion that sanctioned participants tend to resemble long-term welfare participants across a variety of characteristics such as marital status, age, family size, education level, job experience, and most importantly for our purposes, race (Wu et al. 2006; Pavetti, Derr and Hesketh 2003; Hasenfeld, Ghose and Larson 2004; Kalil et al. 2002; Mancuso and Lindler 2001; Koraleck 2000; Westra and Routely 2000;).

The existing literature is creative and sophisticated in its use of econometric methods. Yet it is also limited by its attention to only one side of a two-sided transaction. With only a few exceptions (e.g. Keiser, Meuser, and Choi 2004), previous studies have trained their attention on clients and asked, at least implicitly, “who is likely to get sanctioned?” As a result, these studies largely ignore the fact that welfare sanctions arise, not just from client characteristics and behaviors, but also from policy choices made by a series of political actors such as state and local representatives, program directors, supervisors, and case managers. The rate and incidence of sanctioning depend, at least in part, on the decisions such actors make as they set policy, organize administrative practice, and apply general rules to specific instances. In short, research to date has largely failed to address sanctions as tools of governance that may or may not be deployed depending on what officials choose in particular jurisdictions and cases.

The sanctions deployed at the frontlines of welfare agencies are products of choice at numerous levels of governance. Under federal welfare reform, states confronted a range of options in designing their sanction policies. Seventeen adopted “immediate full-family sanctions,” which suspend all benefits for all members of the TANF family as penalty for the first instance of noncompliance. Fifteen chose “gradual full-family sanctions," which can eventually rise to a "full-family" impact, but only after a progression of penalties. The remaining states selected "partial sanctions," which reduce only a portion – typically the adult portion – of the family's benefits (Pavetti, Derr and Hesketh 2003). Studies suggest that these choices have been quite consequential, and as a result, the decline in the welfare caseload has been as much as 25 percent greater in states with immediate full-family sanctions, compared to states with the least punitive sanction policies (Rector and Youseff 1999).

The politics of policy choice is equally crucial at the local level, where a variety of officials hold responsibility for interpreting state sanction policy and specifying it through rulemaking. Many states have practiced "second-order devolution" where they devolved primary authority over TANF policy down to county government officials or regional workforce boards (Gainesborough 2003). In all states, however, local TANF officials are likely to have substantial discretion over program elements that affect the rate and incidence of sanctioning. Their local policy choices typically establish the process by which clients are informed of TANF rules and penalty procedures; the mode of monitoring participation in required activities; the steps for initiating a sanction, including how and when clients are notified of an impending sanction and the steps needed to avoid it; and the procedures and requirements for "curing" a sanction and returning a client to the TANF rolls. Not surprisingly, given the wide scope of this discretion, field studies find that, even when TANF offices operate under the same state guidelines, sanction philosophies and practices tend to vary dramatically across local communities (Pavetti et al. 1998).

State and local policy choices combine to create the settings for more proximate decision processes in the interactions between case managers and clients. As frontline workers, case managers exercise significant discretion in interpreting program rules and applying penalties (Myers et al. 2006; Pavetti, Derr and Hesketh 2003: 6). Their decisions can influence sanction patterns in a variety of ways. From the outset, case managers are responsible for assessing client needs and capabilities. They identify which clients should be relieved of certain work requirements or told to attend special classes. In doing so, the case manager establishes particular sanctioning parameters for particular clients. And of course, ultimately, it is the case manager who must decide whether to initiate sanction procedures in response to an infraction and whether the circumstances of the infraction justify a "good cause" exception to the rules.

In short, sanctions can be understood, not just as events that happen to some clients more than others (as in most of the current literature), but also as outcomes of governmental decision-making processes. From this perspective, the analytic focus shifts from an investigation of client characteristics *per se* to an investigation of how client characteristics become relevant to administrative decisions and interact with the environmental forces that systematically shape administrative behavior. In what follows, we pursue exactly this approach, investigating how client racial characteristics combine with stereotype-consistent cues and local political environments to influence frontline decisions to impose sanctions.

RACIAL CLASSIFICATION AND THE LOCAL POLITICS OF PUNISHMENT

Our analysis is based on a simple but general model of race and social policy choice called the Racial Classification Model (RCM). Elaborating on the work of Schneider and Ingram (1993, 1997), the RCM identifies how and when racial classifications should affect target-group constructions and, hence, policy design and implementation patterns. The model does not preclude but, crucially, does not assume the operation of racial animus, racial threat, or racial group loyalty. It is a minimalist cognitive model of policy decision making that focuses solely on the necessity of social classification and consequences of group reputation. Here, we apply the model to policy implementation settings by deriving four hypotheses regarding TANF sanctions. The RCM consists of three basic premises.

1. To be effective in designing policies and applying policy tools to specific target groups, policy actors must rely on salient social classifications and group reputations; without such classifications, they would be unable to bring coherence to a complex social world or determine appropriate action.

2. When racial minorities are salient in a policy context, race will be more likely to provide a salient basis for social classification of targets and, hence, to signify target differences perceived as relevant to the accomplishment of policy goals.

3. The likelihood of racially patterned policy outcomes will be positively associated with the degree of policy-relevant contrast in policy actors' perceptions of racial groups. The degree of contrast, in turn, will be a function of (a) the prevailing cultural stereotypes of racial groups, (b) the extent to which policy actors hold relevant group stereotypes, and (c) the presence or absence of stereotype-consistent cues.

The RCM asserts that policy actors try to choose courses of action that they expect to be effective given what they believe about the specific groups they aim to address. In this process, social group characteristics can serve as proxies for more detailed information about a policy's intended target group. When race is salient to a policy area, as in the case of welfare (Gilens 1999), racial classifications can serve precisely this function, regardless of whether racial animus is present and regardless of the decision makers' own racial identities. In such cases, group reputations can frame interpretations of ambiguous policy-target behaviors and cue broad assumptions about what kinds of policy actions are likely to be effective. The effects of such reputations, however, will depend on their specific policy-relevant content as well as situational factors that may strengthen or weaken their utility as information proxies.

In applying the RCM to TANF sanction decisions, we conceptualize sanctions as tools for motivating welfare clients, stimulating work effort, and enforcing responsible behavior. Accordingly, local policy actors should be more likely to organize and implement sanctions in a stringent fashion when TANF clients are perceived as less motivated and responsible in their own right – i.e., when clients are perceived as needing a stronger external stimulus to follow program rules and achieve welfare-to-work goals. In this context, client race should affect sanctioning patterns to the extent that contrasts between racial-group reputations convey information about motivation, work effort, and personal responsibility.

Combining these assumptions with research on group stereotypes in the United States, we can derive our first hypothesis from the RCM. Relative to white Americans, black Americans remain strongly associated with low work effort and motivation, socially irresponsible behavior, and preferences for welfare reliance (Gilens 1999; Schuman et al. 1997). Stereotypes of Latinos occupy a midpoint, less negative than blacks but more negative than whites (Fox 2004). Accordingly, the RCM suggests:

(H1) The Simple Disparity Hypothesis: All else equal, TANF officials will be more likely to sanction black clients than white or Latino clients, and more likely to sanction Latino clients than white clients.

Taken in isolation, H1 might be read as predicting that client race will have a significant and invariant effect in all circumstances. The third premise of the RCM, however, points to a more situational effect, with the impact of client race depending on both the degree to which policy actors hold relevant group stereotypes and the presence or absence of stereotype-consistent cues. These elements of the RCM allow us to move beyond simple disparities to specify and test more nuanced hypotheses.

Two individuals who are perceived as members of the same racial group may nevertheless be associated with quite different group reputations. Research on intersectionality, for example, emphasizes that the meaning of one category of social identity can be altered when combined with another (e.g., when woman is combined with black as opposed to white, see Crenshaw 1991; Hancock 2006). Likewise, social cognition researchers have shown that perceivers tend to distinguish "subtypes" of racial groups (such as "ghetto blacks" vs. "black businessmen") and to attribute negative global-group traits to these subtypes to very different degrees (Richards and Hewstone 2001; Devine and Baker 1991). As a result, race-of-target effects will often be contingent on additional characteristics that strengthen or weaken the individual's connection to the racial group's prevailing reputation. Eberhardt et al. (2006), for example, find that black defendants convicted of killing white victims are more likely to receive the death penalty if they are perceived as having a "stereotypically black appearance."

Indeed, research suggests that a variety of stereotype-consistent cues can enhance race-based disadvantages. In some instances, this effect functions to widen already-existing racial disparities. In others, it creates disparities where none had otherwise existed. The former scenario is well illustrated by Pager's (2003) influential field experiment exploring the effects of race and "the mark of a criminal record" on hiring outcomes. Pager (2003) finds that black job applicants are already disadvantaged relative to whites in the no-felony condition of her experiment, yet the attribution of a felony conviction actually reduces black applicants' job prospects to *a greater degree* than the prospects of already-advantaged white applicants. By contrast, the latter scenario is illustrated by a recent study of how racial cues affect preferences for political candidates. Valentino, Hutchings, and White (2002: 86) find that

> When the black racial cues are stereotype-inconsistent, the relationship between racial attitudes and the vote disappears…. [Likewise] the presence of black images alone… does not prime negative racial attitudes…. The effect emerges only when the pairing of the visuals with the narrative subtly reinforces negative stereotypes in the mind of the viewer.

Building on this research, the RCM suggests that, regardless of whether simple racial disparities exist, the presence of a trait that is consistent with minority-group stereotypes should increase the odds of a sanction more for minority clients than for white clients. In the present analysis, perhaps the most relevant trait of this sort is long-term welfare usage (Gilens 1999). For several decades now, welfare “dependency” has been a prominent stereotype associated with poor racial minorities – interpreted in various quarters as a distinguishing feature of “underclass” culture, a pathology akin to addiction, and a clear sign of an individual’s unwillingness to work (Schram 1995). To stay on the welfare rolls for a comparatively long period of time is, in the context of welfare-to-work programs, to mark oneself in a policy-relevant and racially-inflected way. Accordingly, the RCM suggests:

(H2) The Time-Contingent Disparity Hypothesis: The effects of race on sanctioning will grow stronger, and thus the racial disparities predicted under H1 will grow larger, as TANF participation spells increase in length.

The third premise of the RCM also predicts that the production of racial disparities will be contingent on the extent to which policy actors hold relevant group stereotypes. When such stereotypes are viewed with skepticism, racial-group reputations will be perceived as poor proxies for more detailed information and, hence, as unreliable guides to decisions regarding policy targets. By contrast, when such stereotypes are accepted as largely valid, differences in group reputation should convey more useful information and, thus, the likelihood of racially patterned outcomes should increase.

Ideally, one would like to test this component of the RCM by utilizing direct measures of stereotype acceptance at the individual level. In the absence of such measures, one must seek out a suitable proxy by asking *where* in American society one is likely to find greater acceptance of stereotypes regarding racial-group orientations toward work and welfare. As an empirical matter, one answer to this question is that such stereotypes are more likely to be found in politically conservative communities than in politically liberal communities. There is, of course, no logical reason why conservatives should hold more negative views of racial minorities. Indeed, there are good reasons to distinguish between the two when trying to explain public opposition to various policies designed to advance egalitarian goals (Sears, Sidanius, and Bobo 2000). Nevertheless, two basic empirical observations emerge as uncontroversial in the existing literature. First, conservatives are more likely than liberals to oppose welfare and to hold negative views of welfare recipients (Gilens 1999; Cook and Barrett 1992). Second, conservatives are more likely than liberals to hold negative stereotypes of African Americans, especially in attributing to this group a preference for receiving welfare rather than working (Domke 2001; Federico and Sidanius 2002; Gilens 1999; Glaser 1994; Johnson and Marini 1998; Oliver and Mendelberg 2000).

This relationship can be easily documented with recent data from the National Election Study (NES). In 2004, the NES asked respondents to evaluate the degree to which different racial and ethnic groups (including blacks, Latinos, and whites) were either “hardworking” or “lazy.” Figure 2 presents the difference in the average assessment of each group’s perceived “laziness” (on the vertical axis), by respondents’ liberal-conservative identification (on the horizontal axis), thus allowing a direct examination of the relationship between ideological orientations and group stereotypes.

[Figure 2]

The results are clear. Conservatives are more likely than liberals or moderates to view blacks and Latinos as lazy, compared to whites. As a result, the perceived gap between group reputations grows consistently larger as one shifts from the liberal to the conservative end of the ideological spectrum. One need not engage the thorny causal questions of why this relationship exists to pursue our present analytic goals. It suffices to say that there is an empirical basis for assuming that, in more conservative political environments, one is more likely to find negative views of welfare reliance and acceptance of stereotypes asserting that racial minorities differ from whites in preferring welfare over work.

Combining this observation with the RCM leads to the expectation that racial disparities in sanctioning will be larger in more politically conservative communities. A variety of mechanisms may underlie this relationship. Because case managers tend to be drawn from local communities, one would expect them to be more conservative (and hence, more likely to perceive racial-group differences in welfare-work orientations) when their offices are located in more conservative environments. The same selection dynamics are likely to operate for TANF supervisors, program directors, and governing board members. As a result, racial disparities in sanctioning may arise more often in conservative communities, not only due to the attitudes of individual case managers, but also due to the possibility that racial classifications may guide the officials above them who set local operating procedures and manage TANF implementation. Finally, because welfare agencies are "open systems" that must respond to their political environments, conservative and liberal communities may produce different patterns of program implementation because of the ways that racialized understandings inform political pressures, standards of legitimacy, and agendas for action. For all these reasons, the RCM suggests:

(H3) The Ideology-Contingent Disparity Hypothesis: The effects of race on sanctioning will be stronger, and thus the racial disparities predicted under H1 will be larger, in conservative political environments.

Finally, we offer a fourth hypothesis which predicts that the causal mechanisms assumed by the RCM will have the strongest effect on implementation outcomes in administrative environments which offer greater discretion to local decisionmakers and frontline workers. As discussed above, many states have devolved significant authority in TANF implementation to local governments or regional workforce boards – a process known as second-order devolution (SOD) (Gainsborough 2003). Under SOD, states grant administrative authority to local governments just as the federal government grants authority and responsibility to state governments in designing and implementing TANF (Nathan, 1997; Adkisson, 1998). Although the division of state-local responsibility varies across SOD states, it is generally agreed that in SOD states, local TANF administrators and frontline staff enjoy greater freedom from state control and thus have the opportunity to exercise significantly greater discretion in the implementation of TANF (Fording, Soss and Schram 2007). Therefore, it is likely that the effects of racial classification may be enhanced in SOD states, and as a result, racial classification may more likely to result in racial disparities in sanctioning outcomes. This leads to our final hypothesis:

(H4) The Decentralization-Contingent Disparity Hypothesis: The effects of race on sanctioning will be stronger, and thus the racial disparities predicted under H1 will be larger, in states which have chosen to engage in second order devolution.

To date, we are aware of only one study that has examined the effects of both race and the political environment on TANF sanctioning. Keiser, Mueser and Choi (2004) analyze administrative data on sanctioning in Missouri, finding that after one controls for relevant client-level and county-level factors, blacks are on average 23 percent more likely to be sanctioned than white recipients. Moreover, the authors find that the magnitude of racial differences varies across seven geographic regions in a pattern that they interpret as politically meaningful. Yet, as a test of the RCM's predictions, Keiser, Meuser, and Choi's analysis has significant limitations.

Because of small sample sizes, the authors could only examine how racial effects varied across a small number of geographic groupings (seven), all of which were defined at a high level of aggregation and some of which were non-contiguous. As a result, they were unable to test whether differences in racial effects correlated with any direct measure of political ideology (H3). In addition, because Keiser, Meuser, and Choi relied on binary logit analysis rather than event history analysis, they were unable to test hypotheses related to spell length (H2) and their analysis of racial disparities offered only a limited ability to control for differences in at-risk periods across groups (see Wu et al. 2006). Finally, by examining a single state operating in a centralized administrative environment, not only is their analysis limited in external validity, it cannot test for differences in racial effects due to variation in administrative decentralization (H4).

Our analysis follows Keiser, Mueser and Choi (2004) in that we examine the individual and contextual determinants of sanctioning using administrative data. However, we build on their analysis in several important ways. First, by examining sanctioning in the state of Florida, where the minority population is not only sizeable but widely dispersed, we are able to examine racial effects across a large number of geographic contexts at a lower level of aggregation that is both administratively and politically meaningful (the county). As a result, we can directly test the interaction of local political ideology and client racial characteristics. Second, because Florida is a racially diverse state, we are able to examine disparities in sanctioning across blacks, whites, *and* Latinos. Third, by modeling sanctions using an event history design, we are able to control for group differences in at-risk periods and directly test the hypothesis that racial disparities will grow across the length of the welfare spell. And finally, by extending our analysis to national data on TANF sanctions, we are not only able to test for the effects of administrative decentralization, but we are able to provide a level of external validity that is absent in prior studies of sanctioning.

RACE AND SANCTIONING IN THE FLORIDA TANF PROGRAM: AN ANALYSIS OF ADMINISTRATIVE DATA

We have selected the Florida WT program for our study, not because it is typical of all state TANF programs, but because it provides close to an ideal setting for analyzing how race and politics affect local differences in the use of penalties. Since 1996, Florida, as a practitioner of "second-order devolution," has constructed one of the most decentralized TANF programs in the country. Frontline services have been contracted out to public, non-profit, and for-profit providers throughout the state, and primary authority over the WT program has shifted down to 24 local public/private "Regional Workforce Boards" (RWBs). These RWBs are responsible for strategic planning, policy development, contracting, and oversight of local one-stop delivery systems. Several of the regions encompass more than one county; and the regional boards set policy in a way that allows for county offices to have some discretion in implementing policies. The regional boards are overseen, not by state agencies, but by a statewide public/private partnership called Workforce Florida, Inc. (WFI). The Florida Department of Children and Families (DCF), a conventional state agency, receives the federal TANF block grant and maintains responsibility for eligibility determination. But otherwise, Florida stands out among American states for its emphasis on local control and privatization within a work-oriented TANF program (Botsko et al. 2001: 7).

Florida also scores high on factors that raise the importance of sanction decision processes. After 1996, Florida adopted "some of the strictest time limits and work requirements in the nation" and broadened the pool of clients subject to sanctions by creating "few possibilities for exemptions" (Botsko et al. 2001: 4). The sanctions themselves also fall at the strong end of the continuum, resulting in an immediate, full-family loss of TANF benefits and a reduction of Food Stamp benefits to the fullest extent permitted by federal law (Botsko et al. 2001: 6). Moreover, as we can see in Figure 1, Florida employs sanctions at an extremely high rate compared to other states with full-family sanctions. Thus, it is not surprising to find that Florida DCF identified sanctions as the most common cause of TANF case closings in fiscal year 2003, accounting for 31 percent of closings vs. 21 percent for increased earnings.

The selection of Florida also aids our analysis because it offers significant variation on the two most critical independent variables in our study: race and local political environment. Florida is one of the most racially diverse states in the country, with sizeable black and Latino populations, and the state's TANF population displays even more diversity. Between January 2000 and March 2004, 36.2% of TANF adults were black, 33.7% were white (non-Latino), and 28.5% were Latino. In addition, Florida is a politically diverse state, a fact clearly reflected in recent presidential elections. Over the last three presidential elections, the average Democratic share of the two-party vote across Florida's 67 counties has been approximately 44%, with a healthy standard deviation of 9.2%. The most conservative counties have supported Republicans by a strong majority, with the Republican vote share as high as 75% in some counties (e.g. Okaloosa, Santa Rosa, Clay). The most liberal counties have likewise supported Democrats by a significant margin, with Broward and Gadsen counties leading the way (66% and 69% Democratic vote share, respectively). In combination with Florida's heavy emphasis on sanctioning and decentralized approach to welfare provision, this variation in race and ideology provides an ideal setting for a study of the joint effects of race and ideology on local sanction implementation.

Data and Methods

Our sample consists of individual-level administrative data for all new adult TANF clients who entered WT during the 24-month period from January 2001 through December 2002,[1] supplemented with contextual data indicating how local implementing environments vary across the state's 67 counties. Thus, our entire period of analysis extends from January 2001 (first cohort enters) through November 2003 (12th month of spell for last cohort). We follow each of the 24 cohorts for up to a maximum of 12 consecutive months, ending our observations of the case at the spell's termination or at the 12-month mark, whichever comes first. We restrict our attention to the first TANF spell for each individual during

[1] We define "new" TANF clients as those clients who have spent at least twelve continuous months without TANF benefits.

this period, defined based on continuous months of TANF receipt. As defined, and accounting for a small percentage of cases for which values on some variables are missing, our total sample size exceeds 74,000 individuals who were subject to over 28,000 sanctions across approximately 200,000 person-month observations.

As our data consist of variables that are measured at two levels of analysis (individual and county), we examine the determinants of sanction usage by employing a discrete-time multilevel event history analysis of the initiation of a sanction (Barber et al. 2000). Our dependent variable is $Sanction_{tjk}$ - a dichotomous variable which indicates whether or not client *j*, residing in county *k* has been sanctioned in month *t*. We estimate our model using the logit link (i.e., as a hierarchical generalized linear model, or HGLM), and therefore the effects of the independent variables are additive and represent the change in the log odds of sanction. The independent variables include individual-level measures capturing client effects, and county-level measures capturing community-context effects.[2] The individual-level (level 1) model is represented below in equation 1.

$$\begin{aligned}
Sanction_{tjk} = {} & \beta_{0k} + \beta_{1k} Black_j + \beta_{2k} Latino_j + \beta_{3k} (Black_j * Month\ of\ Spell_{tj}) \\
& + \beta_{4k} (Latino_j * Month\ of\ Spell_{tj}) + \beta_{5k} Number\ of\ Children(2)_j \\
& + \beta_{6k} Number\ of\ Children(3\ or\ more)_j + \beta_{7k} Age\ of\ Youngest\ Child(3\ months\text{-}2\ yrs.)_j \\
& + \beta_{8k} Age\ of\ Youngest\ Child(3\text{-}4\ years)_j + \beta_{9k} Age\ of\ Youngest\ Child(5\text{-}11\ yrs.)_j \\
& + \beta_{10k} Age\ of\ Youngest\ Child(12\ or\ more)_j + \beta_{11k} Education(H.S.)_j + \beta_{12k} Education(>H.S.)_j \\
& + \beta_{13k} Male_j + \beta_{14k} Citizen_j + \beta_{15k} Age_j + \beta_{16k} Single\text{-}Parent_j + \beta_{17k} Earned\ Income_j + \beta_{18k} Month\ 2_{tj} \\
& + \beta_{19k} Month\ 3_{tj} + \beta_{20k} Month\ 4_{tj} + \beta_{21k} Month\ 5_{tj} + \beta_{22k} Month\ 6_{tj} + \beta_{23k} Month\ 7_{tj} + \beta_{24k} Month\ 8_{tj} \\
& + \beta_{25k} Month\ 9_{tj} + \beta_{26k} Month\ 10_{tj} + \beta_{27k} Month\ 11_{tj} + \beta_{28k} Month\ 12_{tj} \qquad [1]
\end{aligned}$$

An important feature of a multilevel model is that the coefficients for the level 1 variables (the β's in equation 1) are permitted to vary across our level 2 units (counties). Based on theoretical

[2] We provide detailed variable descriptions, including data sources and descriptive statistics for each variable used in the analyses that follow in an unpublished Appendix, available at www.xxx.xxx.

expectations, the county-level variables therefore enter the level 2 model as explanatory variables for these effects. Our level 2 model is reflected below in equations 2-5 and reflects our theoretical expectations concerning the effects of county-level variables on the intercept (β_{0k}) of equation 1 (i.e., the mean rate of sanctioning across counties), and on the slope values for *Black* and *Latino* (i.e. β_{1k} and β_{2k}).

$$\beta_{0k} = \gamma_{00} + \gamma_{01}\textit{Local Conservatism}_k + \gamma_{02}\textit{Percent Black}_k + \gamma_{03}\textit{Percent Latino}_k + \gamma_{04}\textit{Annual Wage}_k + \gamma_{05}\textit{Unemployment Rate}_k + \gamma_{06}\textit{Poverty Rate}_k + \gamma_{07}\textit{Population}_j + \gamma_{08}\textit{TANF Caseload}_j + \varepsilon_{0k} \quad [2]$$

$$\beta_{1k} = \gamma_{00} + \gamma_{01}\textit{Local Conservatism}_k + \varepsilon_{1k} \quad [3]$$

$$\beta_{2k} = \gamma_{00} + \gamma_{01}\textit{Local Conservatism}_k + \varepsilon_{2k} \quad [4]$$

$$\beta_{pk} = \gamma_{p0} \qquad \text{for p = 3-28} \quad [5]$$

Level 1 Hypotheses

To test the simple disparity hypothesis, we classify clients as belonging to one of three mutually exclusive racial/ethnic group combinations: black, Latino, and white (non-Latino).[3] We then include the dichotomous variables *Black* and *Latino* in our model of sanction initiation, where we expect the coefficient values in equation 1 will be positive for each of these variables (β_1, $\beta_2 > 0$), and the coefficient for *Black* will be larger than the coefficient for *Latino* ($\beta_1 > \beta_2$). We test the time-contingent disparity hypothesis by including two interaction terms in equation 1, *Black*Month of Spell* and *Latino*Month of Spell*, where we expect that the coefficients for these variables will be positive (β_3, $\beta_4 > 0$).

Based on past research on sanctions and welfare implementation, we include a number of other variables to control for variation in clients' individual characteristics. These include variables measuring the client's sex (*Male*), citizenship status (*Citizen*), and age (*Age*). The age of the youngest child in the TANF family is measured by a series of categorical variables, as is the number of children in the TANF family. We include the client's marital status (*Single-Parent*), and two indicators of human capital (*Earned Income* and *Education*). All of these variables have been found to be important determinants of

[3] We omit a very small percentage (<2%) of cases classified as "other race" by the state.

individual sanctioning outcomes in past research (Wu et al. 2006; Hasenfeld, Ghose and Larson 2004; Kalil et al. 2002; Mancuso and Lindler 2001).

Finally, we include a series of dummy variables for each month of the spell to model the baseline hazard of sanction. This strategy has the advantage of leaving the shape of the baseline hazard function unspecified, which in this sense is analogous to the continuous-time formulation of the Cox proportional hazards model (Beck, Katz and Tucker 1998).

Level 2 Hypotheses

We model the intercept of equation 1 (β_{0k}) as a function of several different features of local political, economic and social environments. Due to the theoretical importance of local ideology for our analysis we rely on two alternative measurement strategies. First, for each of Florida's 67 counties we coded election results for 18 ideologically-relevant constitutional amendments that appeared on the ballot throughout the entire state between 1996 and 2004. Based on a factor analysis of support for all 18 amendments, we used factor scores to create an index of county conservatism that runs from 0 (most liberal county) to 1 (most conservative county). This index is labeled *Local Conservatism* and serves as our primary measure of the local political environment in Florida. As an alternative measure, we rely on the Republican share of the two-party presidential vote, averaged over the last three presidential elections. Because partisanship tends to be imperfectly related to political ideology (Miller 1999), we believe our amendment-based measure to be a more valid measure of local ideology. However, as we describe below, our key results are consistent regardless of the indicator we use, enhancing our confidence in the validity of our conclusions.[4]

In addition to local political ideology, we consider one additional dimension of the local political environment: the local racial context. Previous studies have often found that racial context has a significant impact on racially relevant policy outcomes, either through the effects of a "racial threat" felt by the white majority (Key 1949), or the effects of increased minority political power (Keech 1968).

[4] The simple correlation between our measure of county conservatism and our measure of the Republican vote share is reasonably strong at .65. Details on the construction of our amendment-based measure of county ideology are provided in our unpublished Appendix.

Because there is reason to suspect that either effect might exist (see Keiser, Mueser and Choi 2004), we test for effects of community racial composition by including the percentage of the county population that is black and Latino, respectively (*Percent Black*, *Percent Latino*).

Several additional measures capture the effects of local labor markets and employment opportunities, which we expect to affect sanctioning in one of two ways. First, where employment opportunities are relatively numerous and attractive, TANF clients may be more likely to work enough hours to avoid falling out of compliance with TANF rules. Alternatively, local labor market conditions may also influence the sanction decisions of case managers, who may be less inclined to sanction clients when job opportunities are less numerous or less attractive. To capture such effects, we include the county unemployment rate (*Unemployment Rate*), the county poverty rate (*Poverty Rate*), the level of urbanization (as measured by county population - *Population*), and the annual local wage in food service/drinking establishments (*Annual Wage*).

We also include a measure of the county *TANF Caseload*, expressed as a proportion of the county's population. As the caseload size increases, we might expect that, all else equal, administrative pressures to reduce the caseload would result in an increase in sanctioning. Alternatively, as the caseload size increases, if the number of case managers remains fixed, individual case managers may have less time to closely monitor TANF clients for violations of rules, thus resulting in a lower rate of sanctioning.

Finally, we test the ideology-contingent disparity hypothesis by including *Local Conservatism* as a predictor for the effects of *Black* (β_{1k}) and *Latino* (β_{2k}), as reflected in equations 3 and 4. As we expect racial disparities to increase in conservative counties, we expect *Local Conservatism* to be positively related to the effects of *Black* (β_{1k}) and *Latino* (β_{2k}), and thus γ_{11} and $\gamma_{21} > 0$.

Results

Table 1 presents results for two versions of our event history model that differ only in the measure used to capture local ideological environment. For each version, we report both the coefficient values and the associated odds ratios reflecting the proportional increase in the risk of sanction given a one-unit increase in the independent variable.

We begin by examining the dynamics of sanctioning across the TANF spell, as reflected by the effects of the month-of-spell dummies in equation 1 (which collectively reflect the baseline hazard of sanctioning). For ease of interpretation, we do not report these results in Tables 1 and 2, but instead present a graphical display in Figure 3. The vertical axis of the figure is the odds ratio of sanction, based on the coefficient estimates for equation 1. Thus, for each month of the TANF spell reflected on the horizontal in Figure 3, the associated odds ratio reflects the relative odds of sanction, compared to the initial odds at month 1, for a client who has been on TANF throughout all of the preceding months, but has not yet been sanctioned. As can be seen, the risk of sanction steadily decreases throughout the TANF spell. Indeed, by month 5 of the spell the client's risk of sanction has decreased by nearly 50% compared to month 1.

[Figure 3]

Next, we examine the results for our control variables. As expected, we find that sanctions are significantly related to clients' individual traits in both specifications of our model. Specifically, TANF sanctions are significantly more likely to be applied to the small number of men in the program, relative to the large majority of women. The probability of being sanctioned is also higher for clients who are younger, who are heads of two-parent families, who have older children, who are citizens, and who possess less human capital (as measured by education level).

[Table 1]

Moving to our contextual variables, sanctioning appears to be significantly heavier in high-poverty counties with large populations, yet significantly lower in counties with large TANF caseloads. We find weaker effects for other aspects of the local economic context: neither unemployment rates nor local wage levels prove to be consistent predictors. Nor do we find higher levels of minority presence in the community to affect local sanctioning rates. In sum, these results are largely consistent with the results of past studies, and therefore give us greater confidence in the results we report below (Wu et al. 2006;

Keiser, Mueser and Choi 2004; Hasenfeld, Ghose and Larson 2004; Kalil, Seefeldt and Wang 2002; Mancuso and Lindler 2001; Westra and Routely 2000;).

Turning to an examination of our primary hypotheses, we begin with the simple disparity hypothesis predicting that (1) the coefficients for *Black* and *Latino* should be positive (reflecting significant black-white and Latino-white disparities in sanctioning), and (2) the coefficient for *Black* should be larger than the coefficient for *Latino* (due to the presumed larger gap in group reputations between blacks and whites). As specified, equation 1 models the effects of *Black/Latino* as interactive, and therefore conditional on spell duration and the local political environment. To ease interpretation, we transformed *Local Conservatism, Republican Vote Share* and *Month of Spell* by subtracting their means prior to estimation, so that the coefficients for *Black* and *Latino* in Tables 1 and 2 reflect the effects of race and ethnicity for clients who reside in a politically moderate county, and who are at a typical point in the TANF spell (month 3). As can be seen from the coefficient estimates in Table 1, we find mixed support for the simple disparity hypothesis. The coefficients for *Black* are statistically insignificant in both specifications, and while the coefficient for *Latino* is statistically significant, the direction of the effect is the opposite of what is expected and suggests that Latino clients are sanctioned at a rate that is lower rate than whites and blacks alike.

By contrast, the results in Tables 1 offer strong support for the time-contingent disparity hypothesis, as revealed by the significant interactions between race/ethnicity of client and a simple counter variable (1-12) representing the month of the current TANF spell (*Black*Month of Spell*, *Latino*Month of Spell*). This interaction between race/ethnicity and month of spell is extremely robust across both specifications of our model and underscores the importance of employing a longitudinal design, such as event history analysis, to study racial dynamics in TANF sanctioning.

We now turn our attention to the ideology-contingent disparity hypothesis, which predicts that sanctioning disparities are conditioned by the political environment in which TANF is implemented. Our test of this hypothesis is based on the estimation of cross-level interactions between the racial/ethnic status of the client and our measures of local ideology (*Local Conservatism* in Model 1 and *Republican*

Vote Share in Model 2). The results reported in Table 1 strongly support our hypothesis. The effects of racial/ethnic status are indeed mediated by the ideological orientation of the local political environment. Not only is this the case regardless of which measure of local political ideology we use, we also find the mediating effect of ideological climate to be statistically significant for both black and Latino clients (although the effect is only weakly significant for Latino clients in Model 1). For both groups (blacks and Latinos), movement from a liberal local environment to a conservative environment raises the probability of being sanctioned, not just in absolute terms, but also relative to that of white non-Latinos.[5] Given the complexity of the results presented in Table 1 (due to interactions between race, time and ideology), we present a series of graphical interpretations of these effects below in Figures 4 and 5.[6]

In Figure 4 we present three graphs of the predicted probability of sanction for "typical" black and white clients, for months 3, 6 and 9 of the TANF spell.[7] In each graph, the probability of sanction is plotted against our measure of local conservatism, thus allowing us to see how the risk of sanction varies across the entire range of local ideological context for both black and white clients. Figure 4 reveals several interesting features of the relationship between race and sanctioning. We can see that as we move from month 3 to month 9 in of the TANF spell, the overall risk of sanction decreases for both blacks and whites (as reflected by the decreasing intercept values of the curves in the graphs). This is expected given the pattern of the baseline hazard reported in Figure 1. However, this decrease in the rate of sanctioning is not consistent across racial groups as whites display a significantly greater reduction in the risk of sanction than blacks as spell duration increases. This pattern is anticipated given the significant interaction between race and month of spell in our results.

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[5] One possible explanation for these results is that it is not ideology but rather the size of the minority population that mediates the effect of client race on sanctioning. To explore this possibility, we tested for interactions between racial/ethnic status of client and the black and Latino percentage of the county population (i.e. *Black*Black%* and *Latino*Latino%*). Neither term was statistically significant.

[6] All illustrations are based on the results from Table 1. Similar patterns are observed when we use *Republican Vote Share* as our measure of local ideology.

[7] We define a "typical" client as a U.S.-born 31 year-old single woman with 2 children (aged 5-12 years), less than 12 years of education an average level of earned income, who resides in an average county (reflecting mean values on all of the contextual variables).

[Figure 4]

Figure 4 also allows us to see how black-white disparities in sanctioning vary across the level of local conservatism. In all three graphs, we see that as we move from the most liberal to the most conservative environments, the probability of a black client being sanctioned, relative to that of a white client, increases. However, the pattern of racial disparities changes in a substantively meaningful way as the spell length increases. For clients in the third month of a TANF spell, we see that in liberal counties, white clients are significantly more likely to be sanctioned than black clients for a large majority of clients. Only in the most conservative counties does the predicted probability of sanction for blacks exceed that of whites. As the length of the spell grows longer, however, black clients become consistently more likely to experience a sanction than their white counterparts. Indeed, by the 9th month of the spell, black clients are predicted to be sanctioned more than whites in every county, and at a rate that is approximately 70% higher than that of whites in the most conservative counties

Figure 5 consists of three identically constructed graphs comparing predicted probabilities for Latino clients and white clients. As can be seen, we observe several similarities between Figures 4 and 5 in the pattern of racial/ethnic disparities. The risk of sanction for a Latino client, relative to that of a white client, increases as spell length increases, and as one moves from a liberal to a conservative political climate. However, the magnitude of these changes in patterns of Latino-white disparities across these two contexts seems smaller than what we observe for black clients in Figure 4. This is largely consistent with our theoretical assumptions regarding relatively smaller reputational gaps between Latinos and whites, (see Figure 2), and in this sense provides additional support for our hypotheses.[8]

[Figure 5]

[8] We have replicated the analyses presented in Table 1 using two alternative estimation techniques used to analyze multilevel data – logistic regression with clustered standard errors (by county), and a 2-stage regression method. These results confirm the interaction between Black *Client and Local Conservatism*, but unlike the HLM results, also find a significant interaction between *Hispanic Client* and *Local Conservatism*. We do not present these results in this paper as we have the most confidence in the HLM results presented in this paper. However, we do provide these results for interested readers in our unpublished appendix to this paper.

Extending our Results Using National TANF Data

Thus far, our analysis of Florida TANF clients provides strong support for our racial classification model of policy choice. Yet, due our examination of a single state, we are unable to test Hypothesis 4, which predicts that the effects of race/ethnicity will be stronger in states that rely on second-order devolution in TANF implementation. In addition, there may be reason to doubt the external validity of our findings from Florida due to our focus on sanctioning patterns a single state. To address these issues, we now turn to an analysis of sanctioning that incorporates individual-level data from all 50 states, available from the Office of the Assistant Secretary for Planning and Evaluation (ASPE) in the Department of Health and Human Services.

The ASPE data consists of a series of state samples of TANF families collected by all 50 states from 1999 through 2005. These data are used by the federal government for annual state performance evaluations, and consist of two types of samples – active cases and closed cases. In addition to variables measuring clients' personal characteristics, the closed case samples also provide information concerning the reason that the case was closed. Such reasons may include leaving welfare for employment, getting married, and being sanctioned, among others. We therefore rely on these data from the closed case samples to construct our dependent variable. Given that we limit the analysis to closed cases, we expect to find some differences compared to the Florida sample that includes open as well as closed cases. Yet, there is reason to think that sanction exits in this population are affected as we have hypothesized. For our analysis, we pool all closed case samples for the entire 1999-2005 period. We restrict our analysis to all adult TANF recipients who were identified as the head of the household by the state. Using this definition, and accounting for some missing data, our final dataset consists of approximately 195,000 TANF adults, residing in nearly 2,700 different counties. Unfortunately, these data are purely cross-sectional in nature, and therefore we cannot exactly replicate our analysis of sanctioning in Florida using nationally representative data. However, these data do allow us to provide a strong test of the effects of decentralized administration, as well as a limited replication of our tests of the simple disparity hypothesis and the ideology-contingent disparity hypothesis.

Our dependent variable is *Sanction Exit*$_{jk}$ - a dichotomous variable which indicates whether or not client *j*, residing in county *k* left TANF due to a sanction. We model the probability of a sanction exit as a function of both individual level and contextual (county) variables. Therefore, we estimate a (cross-sectional) multilevel model, using a logit link (as our dependent variable is dichotomous). The individual-level (level 1) model is represented below in equation 6.

$$\begin{aligned}
\textit{Sanction Exit}_{jk} = {} & \beta_{0k} + \beta_{1k}\,\textit{Black}_j + \beta_{2k}\,\textit{Latino}_j + \beta_{3k}\,\textit{Number of Children(2)}_j \\
& + \beta_{4k}\,\textit{Number of Children(3 or more)}_j + \beta_{5k}\,\textit{Age of Youngest Child(3 months-2 yrs.)}_j \\
& + \beta_{6k}\,\textit{Age of Youngest Child(3-4 years)}_j + \beta_{7k}\,\textit{Age of Youngest Child(5-11 yrs.)}_j \\
& + \beta_{8k}\,\textit{Age of Youngest Child(12 or more)}_j + \beta_{9k}\,\textit{Education(12 yrs..)}_j + \beta_{10k}\,\textit{Education(>12 yrs..)}_j \\
& + \beta_{11k}\,\textit{Male}_j + \beta_{12k}\,\textit{Citizen}_j + \beta_{13k}\,\textit{Age}_j + \beta_{14k}\,\textit{Single-Parent}_j + \beta_{15k}\,\textit{Earned Income}_j \\
& + \beta_{16k}\,\textit{Public Housing}_j + \beta_{17k}\textit{OASDI}_j + \beta_{18k}\textit{SSI}_j + \beta_{19k}\,\textit{Year2000}_j + \beta_{20k}\,\textit{Year2001}_j \\
& + \beta_{21k}\,\textit{Year2002}_j + \beta_{22k}\,\textit{Year2003}_j + \beta_{23k}\,\textit{Year2004}_j + \beta_{24k}\,\textit{Year2005}_j \qquad [6]
\end{aligned}$$

The county-level (level 2) model is represented below in equations 7-10.

$$\begin{aligned}
\beta_{0k} = {} & \gamma_{00} + \gamma_{01}\textit{Republican Vote Share}_k + \gamma_{02}\textit{Percent Black}_k + \gamma_{03}\textit{Percent Latino}_k + \gamma_{04}\textit{Per Capita Income}_k \\
& + \gamma_{05}\textit{Unemployment Rate}_k + \varepsilon_{0k} \qquad [7]
\end{aligned}$$

$$\beta_{1k} = \gamma_{10} + \gamma_{11}\,\textit{Republican Vote Share}_k + \varepsilon_{1k} \qquad [8]$$

$$\beta_{2k} = \gamma_{20} + \gamma_{21}\,\textit{Republican Vote Share}_k + \varepsilon_{2k} \qquad [9]$$

$$\beta_{pk} = \gamma_{p0} \qquad \text{for p = 3-24} \qquad [10]$$

Level 1 Hypotheses

To test the simple disparity hypothesis, we once again classify clients as black, Latino, and white (non-Latino) and include the dichotomous variables *Black* and *Latino*. We expect the coefficient values will be positive for each of these variables (β_1, $\beta_2 > 0$), and the coefficient for *Black* will be larger than the coefficient for *Latino* ($\beta_1 > \beta_2$).

We also include a number of other variables to control for variation in clients' individual characteristics. Many of these variables are identical (or very similar) to the variables we included in our analysis of Florida TANF clients. These include variables measuring the client's gender (*Male*), citizenship status (*Citizen*), age (*Age*), the age of the youngest child in the TANF family, the number of children in the TANF family, the client's marital status (*Single-Parent*), and two indicators of human capital (*Earned Income* and *Education*). We also take advantage of information in the national TANF dataset and include three additional indicators of client hardship. These include two dichotomous variables indicating whether a former TANF client received disability benefits through the Social Security or Supplemental Security Income programs (*OASDI*, *SSI*), and whether a TANF family lived in public housing (*Public Housing*). Finally, we include a series of dummy variables for the year that the TANF sample was collected.

Level 2 Hypotheses

As in our analysis of Florida TANF clients, we model the intercept of equation 6 (β_{0k}) as a function of local political, economic and social conditions. To measure the local political context in the national data, we rely on the Republican share of the two-party presidential vote, averaged over the last three presidential elections (*Republican Vote Share*). We also control for the local racial context by including the percentage of the county population that is black and Latino, respectively (*Percent Black*, *Percent Latino*). To capture the effects of local economic conditions, we include the county unemployment rate (*Unemployment Rate*) and per capita income (*Per Capita Income*). Finally, we test the ideology-contingent disparity hypothesis by including *Republican Vote Share* as a predictor for the effects of *Black* (β_{1k}) and *Latino* (β_{2k}), as reflected in equations 8 and 9. As we expect racial disparities to be larger in conservative counties, we expect γ_{11} and γ_{21} to be positive.

To test Hypothesis 4 (the Decentralization-Contingent Disparity Hypothesis), we estimate our model separately for the 36 state governments that administer TANF directly, and the 14 states that have devolved significant authority in TANF implementation to local governments or regional workforce boards – i.e., states with second-order devolution (SOD) (Gainsborough 2003).[9] For reasons outlined above, we expect that the effects of racial classification may be enhanced in SOD states, and as a result, racial classification may more likely to result in racial disparities in sanctioning outcomes.

Results

The results of our analyses are presented in Table 2. Once again, we find that sanctions are significantly related to clients' individual traits, within both SOD and centralized (i.e. non-SOD) states. The probability of being sanctioned is higher for clients who are younger, who have older children (SOD states only), who are citizens, and who possess less human capital (as measured by education level and earned income). These results are generally consistent across SOD and non-SOD states, and they are generally consistent with our results from Florida. However, the effects of several variables diverge from our initial findings from Florida. For example, gender plays no role in sanctioning in SOD states, and in centralized states it is women, as opposed to men, who are more likely to be sanctioned. In addition, while being a single parent has no effect on sanctioning in SOD states, we find that single parents are more likely (rather than less likely) to be sanctioned in centralized states. Finally, we see that clients who receive disability benefits or live in public housing are significantly more likely to have left TANF due to a sanction, perhaps reflecting significant employment barriers among these clients.

[Table 2]

Moving to our contextual variables, sanction exits are significantly more likely to occur in counties with larger black populations, as well as in counties with larger Latino populations (in centralized states). We find weaker effects for the local economic context, although the unemployment rate is positively related to sanction exits in centralized states.

We now turn our attention to the tests of the simple disparity and ideology-contingent disparity hypotheses, and the relative performance of these hypotheses across SOD and centralized administrative environemnts. As in our initial analysis of TANF sanctioning among Florida clients, we centered

[9] These 14 states include Arkansas, California, Colorado, Florida, Maryland, Michigan, Minnesota, New York, North Carolina, Ohio, Tennessee, Texas, Utah and Wisconsin.

Republican Vote Share by subtracting its mean so that the coefficients for *Black* and *Latino* in Table 3 reflect the effects of race and ethnicity in a typical (i.e., politically moderate) county. As we saw in Florida, the results for the national data find the effects of race and ethnicity to be insignificant in such a county, in both SOD and non-SOD states. However, this does not mean to suggest that racial disparities are entirely absent, as evidenced by the results of our test of the ideology-contingent hypothesis. As in our Florida analysis, our test of this hypothesis is based on the estimation of cross-level interactions between the racial/ethnic status of the client and our measure of local ideology (*Republican Vote Share*). Among black clients, the effect of racial status is indeed mediated by the ideological orientation of the local political environment. And consistent with theoretical expectations concerning the nature of the implementation environment, the effect is limited to SOD states. However, we find no interaction between ethnicity (*Latino*) and the local political environment in either sample of states.

In Figure 6, we provide a graphical illustration of the relationship between race and sanctioning in SOD states which offers a clearer understanding of exactly how this relationship is mediated by local ideology. Figure 6 presents the predicted probability of a sanction exit for a typical client, by the race of the client and the local political environment.[10] As we saw in Florida (see Figure 4), black clients and white clients are predicted to be sanctioned at more or less equivalent rates in the most liberal counties. But as we move to the right along the horizontal axis, we see that racial disparities quickly emerge in the anticipated direction. Given the consistency of this finding across two very different datasets and research designs, this result provides important confirmation of the mediating effects of local ideology and administrative decentralization, and ultimately, our Racial Classification Model of policy choice.

[Figure 6]

CONCLUSION

Poverty governance in the United States has been redefined over the past several decades by a turn to more custodial and paternalist policy approaches. Today, low-income Americans confront a web

[10] We define a "typical" client as a U.S.-born 31 year-old single woman with 2 children (aged 5-12 years), less than 12 years of education an average level of earned income, who resides in an average county (reflecting mean values on all of the contextual variables).

of policies that place greater emphasis on behavioral expectations, supervision, and disciplinary action. Leading scholars have called for greater efforts to theorize this "penalization of poverty" (Wacquant 2001) and for more direct analyses of how paternalist rules and penalties are implemented at the administrative frontlines (Mead 2004). As students of American politics turn to these crucial tasks, we do so against a historical backdrop that makes discussion of race and local politics virtually unavoidable. Behavioral regulation and the maintenance of social order have traditionally been state and local functions under U.S. federalism (Lowi 1998), and race relations have often stood at the center of the local political dynamics that have shaped both poor relief and punishment (Lieberman 1998; Zimring 2003).

The contemporary turn toward discipline, of course, occurs in a post-civil-rights era that is distant from earlier eras in which caste-like race relations drove overt practices of discrimination, exploitation, and oppression. Yet it remains tempting to fall back on these earlier eras as simplifying templates as we turn today to the question of how, if at all, race affects the official application of disciplinary policy tools. As Loic Wacquant (1997) has explained, scholarly analyses of race in the U.S. are frequently held captive by "the logic of the trial": actors and institutions are held up to scrutiny and deemed racist or non-racist, discriminatory or unbiased, prejudiced or not. The effect of race is all encompassing or non-existent.

In this paper, we have sought to replace such simplifying logics with a more subtle and contingent account of how race matters for the implementation of sanctions under contemporary welfare reform. The evidence presented here provides striking confirmation that racial classification remains central to the operation of poverty governance in the United States. Yet our racial classification model of policy choice suggests, and our empirical results confirm, that racial disparities in treatment arise to a greater degree in some places than in others, and for some client subgroups more than others. Indeed, under some conditions specified here, black and Latino welfare recipients emerge as no more likely, or even less likely, than their white counterparts to be sanctioned. Minority disadvantage arises in the disciplinary process primarily as a function of community and client characteristics that heighten the degree of contrast associated with racial-group reputations.

In the welfare context, for example, long-term program usage is a discrediting mark that can associate a recipient with dependency, irresponsibility, and lack of effort. The toll imposed by this discrediting mark, however, turns out to vary across racial groups in our analysis. As TANF participation spells grow longer, it is minority, and especially black, clients who become increasingly subject to penalties. This result is consistent, not only with the RCM, but also with a line of experimental research suggesting that welfare reliance has different meanings when attached to blacks and whites (Gilens 1999) and that stigmatizing, stereotype-consistent cues interact with racial status to disadvantage blacks relative to whites (Pager 2003; Valentino, Hutchings, and White 2002). In the Florida WT program, black-white disparities increase over the course of the participation spell, suggesting that blacks are disproportionately tainted by – and ultimately taxed for – the stigma of long-term program usage.

Likewise, we find that racial disparities in sanctioning depend significantly on the local context. Larger disparities emerge in more ideologically conservative jurisdictions, and within administrative structures which delegate significant discretion to local decision-makers. Our interpretation of these results flows, once again, from the basic premises of the RCM. According to this model, racial disparities in policy treatment should emerge to a greater degree in those times and places where officials are more likely to hold stereotypes that distinguish racial groups in policy-relevant ways, and where such officials are afforded a greater opportunity to translate those beliefs into policy outcomes.

In addition to shedding light on when and how race matters for policy choice, these results have significant implications for the growing body of research on TANF implementation. Econometric analyses of TANF outcomes, including sanction rates, have generally ignored the political and organizational context of implementation. Our analysis suggests that such models may be under-specified in ways that produce a distorted picture of sanctioning processes and outcomes. Equally important, our analysis suggests a more troubling account of racial disparities than one often finds in many econometric studies. When prior studies have found racial disparities, it has been quite reasonable to ask whether the group differences were simply due to unobserved heterogeneity: perhaps the racial disparity only indicated that minority and white clients with similar demographic profiles actually behaved in different ways. Our study calls attention to a different pattern of results. We have not reported simply that a racial disparity exists that we cannot explain away based on our collection of demographic controls; we have found that the extent of racial disparities in sanctioning rises or falls systematically depending on the political environment. There is, of course, little reason to suspect that black-white or Latino-white differences in individual behavior would vary across liberal and conservative jurisdictions, or across administrative structures. Thus, the findings point to a racial effect that seems far harder to dismiss as not being “really” about race.

Finally, we end with a broader point about the history of welfare provision. Scholars have often suggested that welfare has followed a “vicious cycle” in which social inequalities have shaped policy choices that, in turn, have recreated social inequalities (Mettler 1998; Lieberman 1998; Schram 2006). Negative images of a target group guide policy design and implementation, and policy designs are then implemented in ways that reinforce negative group outcomes and reputations (Schneider and Ingram 1993; Schram 1995, 2005). The analysis presented here raises the prospect that such dynamics may be at work today in contemporary welfare reform. Racial politics contributed greatly to federal welfare reform in 1996 (Gilens 1999; Soss and LeClair 2004), which devolved substantial policy authority to the states. State policy choices regarding TANF rules and penalties were then significantly influenced by racial composition (Fellowes and Rowe 2004; Fording 2003; Soss et al. 2001). Nationally, these racialized state choices resulted in black and Latino families being disproportionately concentrated in the policy regimes with the toughest rules and sanctions (Soss et al. 2003). And as we have shown, among the states that adopted the toughest sanctions, racial composition appears to have a significant relationship with the rates at which sanctions are applied. Moving further downward to the state and local level, our analysis suggests that race matters in substantial but variable ways for frontline decisions to impose discipline. Today as in the past, public aid for poor women and children remains entangled with the complex interplay of race, politics, and local policy control.

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TABLE 1. Discrete-Time Multilevel Event History Models of Sanction Initiation

		Model 1		Model 2	
		β	Odds Ratio	β	Odds Ratio
Intercept (β_{0k})	Intercept (γ_{00})	-1.198**	0.384	-0.854**	0.426
	Local Conservatism (γ_{01})	0.133	1.142	---	---
	Republican Vote Share (γ_{01})	---	---	-0.003	0.997
	Black% (γ_{02})	-0.005	0.995	-0.008	0.992
	Latino% (γ_{03})	-0.001	0.999	-0.002	0.997
	Annual Wage (γ_{04})	-0.038	0.963	-0.036	0.965
	Unemployment Rate (γ_{05})	0.003	1.003	0.006	1.006
	Poverty Rate (γ_{06})	0.040**	1.041	0.040**	1.041
	Population (γ_{07})	0.295**	1.295	0.243**	1.243
	TANF Caseload (γ_{08})	-0.227**	0.797	-0.235**	0.791
Race/Ethnicity (reference=white)					
Black (β_{1k})	Intercept (γ_{10})	-0.041	0.960	-0.020	0.980
	Local Conservatism (γ_{11})	0.371**	1.449	---	---
	Republican Vote Share (γ_{11})	---	---	0.007**	1.449
Latino (β_{2k})	Intercept (γ_{20})	-0.139**	0.870	-0.101**	0.904
	Local Conservatism (γ_{21})	0.376+	1.458	---	---
	Republican Vote Share (γ_{21})	---	---	0.010*	1.010
Black * Month of Spell (β_{3k})	Intercept (γ_{30})	0.065**	1.067	0.065**	1.067
Latino * Month of Spell (β_{4k})	Intercept (γ_{40})	0.035**	1.035	0.035**	1.035
Number of children (reference = 0 - 1):					
Two (β_{5k})	Intercept (γ_{50})	-0.024	0.976	-0.024	0.976
Three or more (β_{6k})	Intercept (γ_{60})	-0.025	0.975	-0.025	0.975
Age of youngest child (reference = 0 – 2 months):					
3 months – 2 years (β_{7k})	Intercept (γ_{70})	0.567**	1.763	0.567**	1.763
3 – 4 years (β_{8k})	Intercept (γ_{80})	0.656**	1.927	0.656**	1.927
5 – 12 years (β_{9k})	Intercept (γ_{90})	0.740**	2.097	0.740**	2.097
More than 12 years (β_{10k})	Intercept (γ_{100})	0.758**	2.134	0.758**	2.134
Education (reference = <H.S.)					
H.S. Education (β_{11k})	Intercept (γ_{110})	-0.322**	0.724	-0.323**	0.724
More than H.S. Education (β_{12k})	Intercept (γ_{120})	-0.483**	0.617	-0.483**	0.617
Male (β_{13k})	Intercept (γ_{130})	0.190**	1.209	0.190**	1.209
Citizen (β_{14k})	Intercept (γ_{140})	0.154+	1.166	0.154+	1.166
Age (β_{15k})	Intercept (γ_{150})	-0.022**	0.979	-0.022**	0.979
Single-Parent (β_{16k})	Intercept (γ_{160})	-0.200**	0.819	-0.200**	0.819
Wage Income (β_{17k})	Intercept (γ_{170})	-0.007	0.993	-0.007	0.993
Number of Subjects (Level 1)		74,517		74,517	
Number of Failures		28,307		28,307	
Time at Risk (Person-Months)		198,147		198,147	
Number of Counties (Level 2)		66		66	

*p<.05, **p<.01, +p<.10

Note: The analysis was conducted using HLM 6 (HLM2 module). The cell entries are binary logit coefficients, with significance levels determined based on robust standard errors. Each model includes dummy variables for the duration of spell (months 1-12). These results are not shown, but are available from the authors upon request.

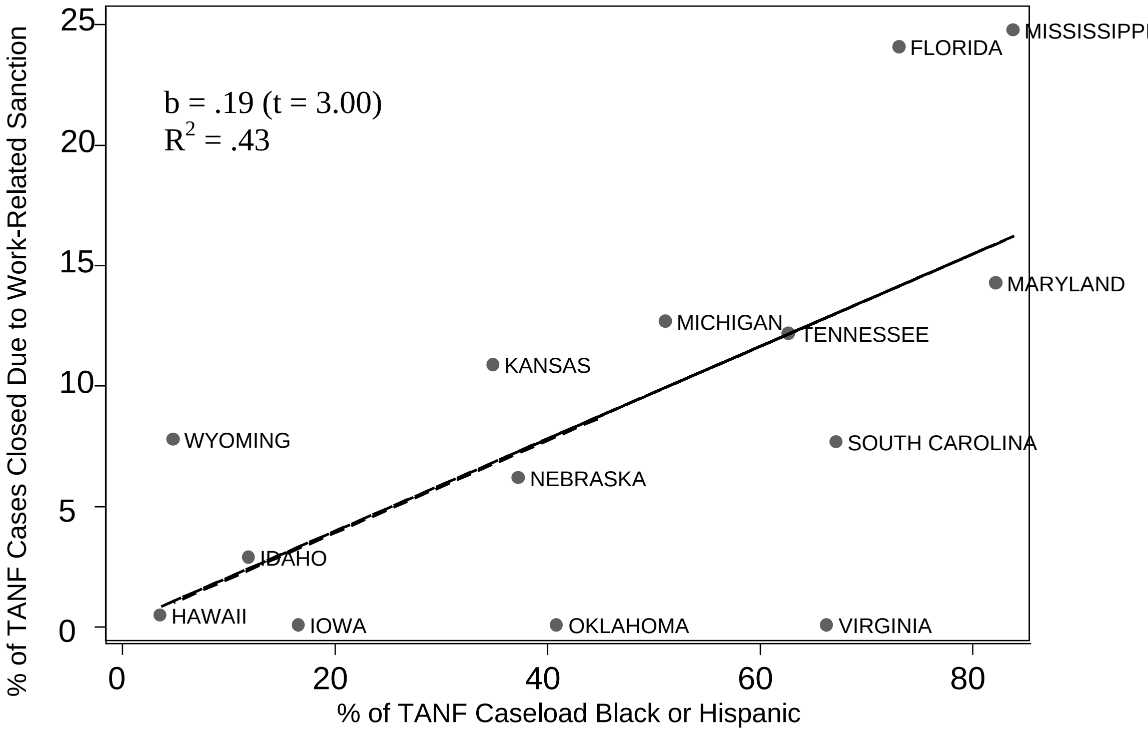
TABLE 2. Multilevel Logistic Regression Models of Sanction Exits, for TANF Clients in SOD and Centralized States

		SOD States	***Centralized States***
Intercept (β_{0k})	Intercept (γ_{00})	-7.909**	-10.840**
	Republican Vote% (γ_{01})	0.032**	0.040**
	Black% (γ_{02})	0.021**	0.023**
	Latino% (γ_{03})	-0.004	0.011**
	Per Capita Income (γ_{04})	0.001	-0.000
	Unemployment Rate (γ_{05})	0.016	0.070**
<u>Race/Ethnicity (reference=white)</u>			
Black (β_{1k})	Intercept (γ_{10})	0.086	0.113
	Republican Vote Share (γ_{11})	0.010*	0.001
Latino (β_{2k})	Intercept (γ_{20})	-0.020	0.030
	Republican Vote Share (γ_{21})	0.001	0.001
<u>Number of children (reference = 0 - one):</u>			
Two (β_{3k})	Intercept (γ_{30})	0.120**	0.082**
Three or more (β_{4k})	Intercept (γ_{40})	0.240**	0.163**
<u>Age of youngest child (reference = 0 – 2 months):</u>			
3 months – 2 years (β_{5k})	Intercept (γ_{50})	-0.336**	0.659**
3 – 4 years (β_{6k})	Intercept (γ_{60})	-0.038	0.724**
5 – 12 years (β_{7k})	Intercept (γ_{70})	-0.072	0.720**
More than 12 years (β_{8k})	Intercept (γ_{80})	0.016	0.689**
<u>Education (reference = <H.S.)</u>			
H.S. Education (β_{9k})	Intercept (γ_{90})	-0.178**	0.252**
More than H.S. Education (β_{10k})	Intercept (γ_{100})	-0.432**	-0.392**
Male (β_{11k})	Intercept (γ_{110})	0.035	-0.208**
Citizen (β_{12k})	Intercept (γ_{120})	0.407**	0.545**
Age (β_{13k})	Intercept (γ_{130})	-0.020**	-0.020**
Single-Parent (β_{14k})	Intercept (γ_{140})	0.042	0.144**
Earned Income (β_{15k})	Intercept (γ_{150})	-0.001**	-0.002**
Public Housing (β_{16k})	Intercept (γ_{160})	0.089	0.183**
OASDI (β_{17k})	Intercept (γ_{170})	0.833**	1.061**
SSI (β_{18k})	Intercept (γ_{180})	1.357**	1.268**
Number of Clients (Level 1)		44,691	151,457
Number of Counties (Level 2)		963	1,708

*p<.05, **p<.01, +p<.10

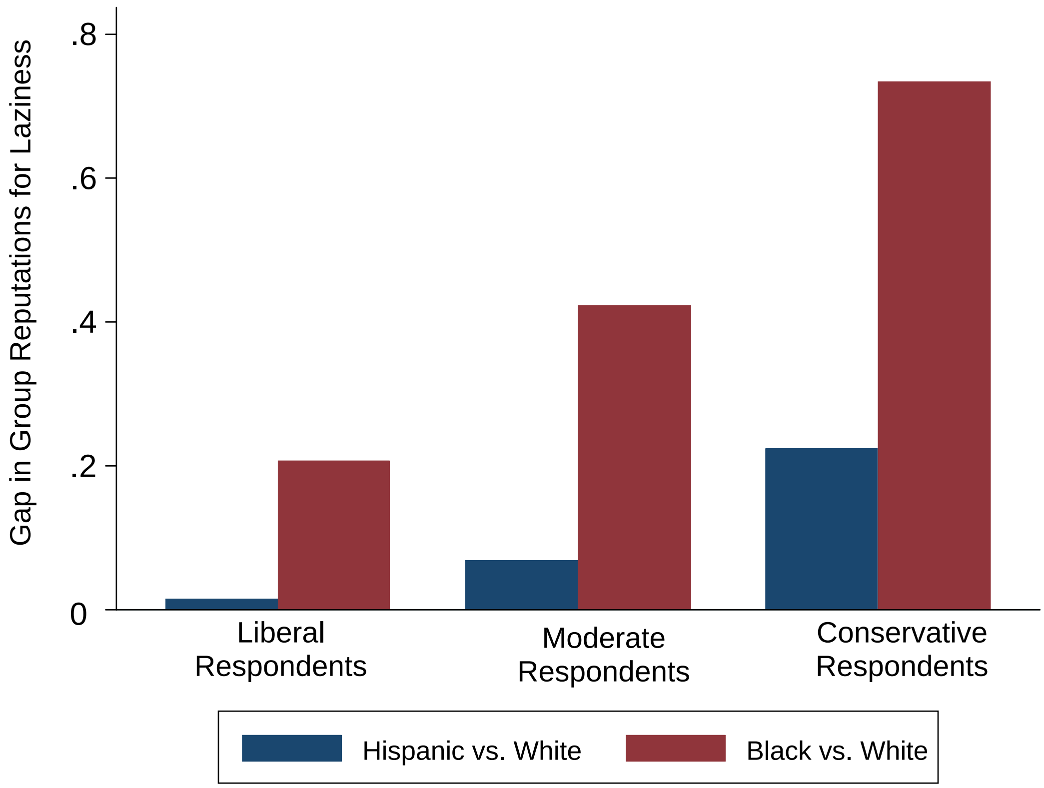
Note: The analysis was conducted using HLM 6 (HLM2 module). The cell entries are binary logit coefficients, with significance levels determined based on robust standard errors. Each model includes dummy variables for the year the state sample was collected (1999-2006). These results are not shown, but are available from the authors upon request.

FIGURE 1. Scatterplot of the Relationship between the Racial/Ethnic Composition of the TANF Caseload and the Work-Related Sanction Rate, FY2002



Note: The sample for this analysis consists of the 14 states which enforced immediate, full-family sanctions in 2002. The horizontal axis is computed as the sum of the percentage of TANF families that are African-American and the percentage that are Hispanic, based on data reported for April, 2002. The vertical axis is the percentage of cases closed due to a work-related sanction during fiscal year 2002. These data are reported in the *Sixth Annual Report to Congress*, published by the Office of Family Assistance in the Department of Health and Human Services (http://www.acf.hhs.gov/programs/ofa/annualreport6/ar6index.htm).

FIGURE 2. Average Gap in Perceived Laziness of Racial Groups, by Ideological Identification of Respondents



Note: The sample for this analysis is taken from 2004 National Election Study and includes respondents of all races. The vertical axis represents the difference in the mean assessment of laziness for each pair of target groups (the mean score for blacks/Hispanics minus the mean score for white), where the laziness scale is coded as follows: 1=hardworking thought 7=lazy.

FIGURE 3. Odds Ratio of Sanction for Florida TANF Clients across the TANF Spell (Baseline = Month 1 of TANF Spell)

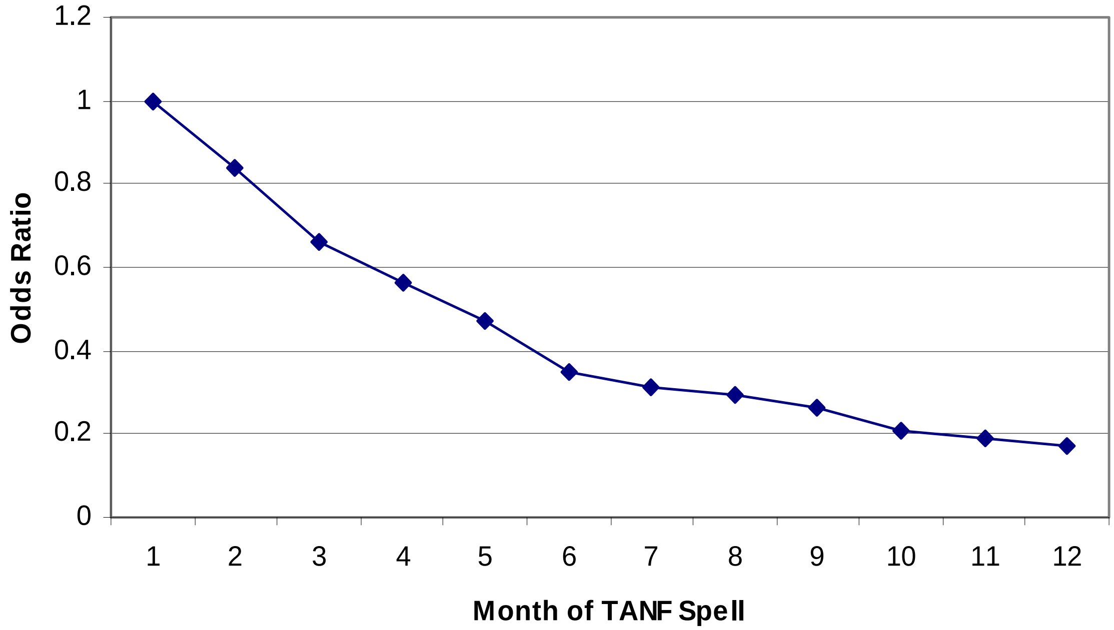


FIGURE 4. Predicted Probability of Being Sanctioned for Florida TANF Clients, by Race, Local Political Environment and Month of TANF Spell

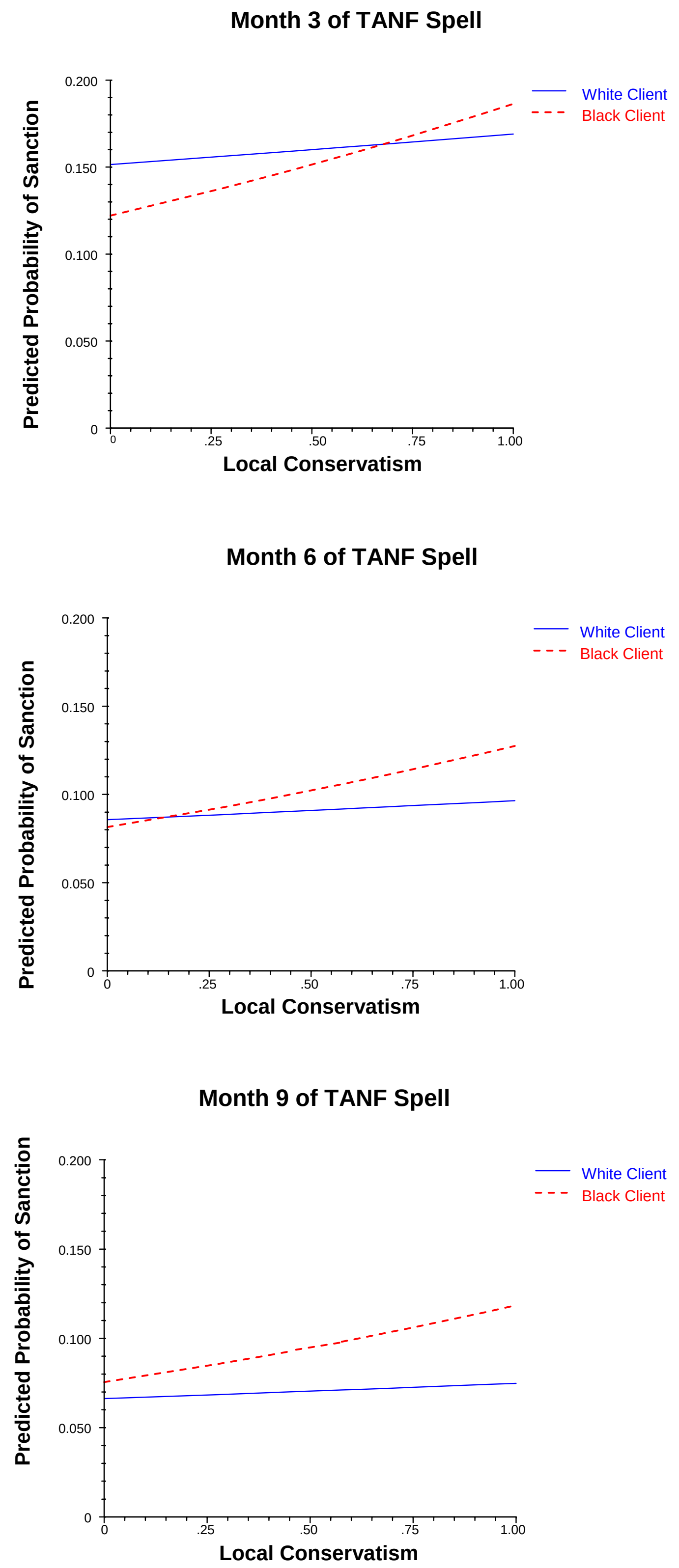


FIGURE 5. Predicted Probability of Being Sanctioned for Florida TANF Clients, by Ethnicity, Local Political Environment and Month of TANF Spell

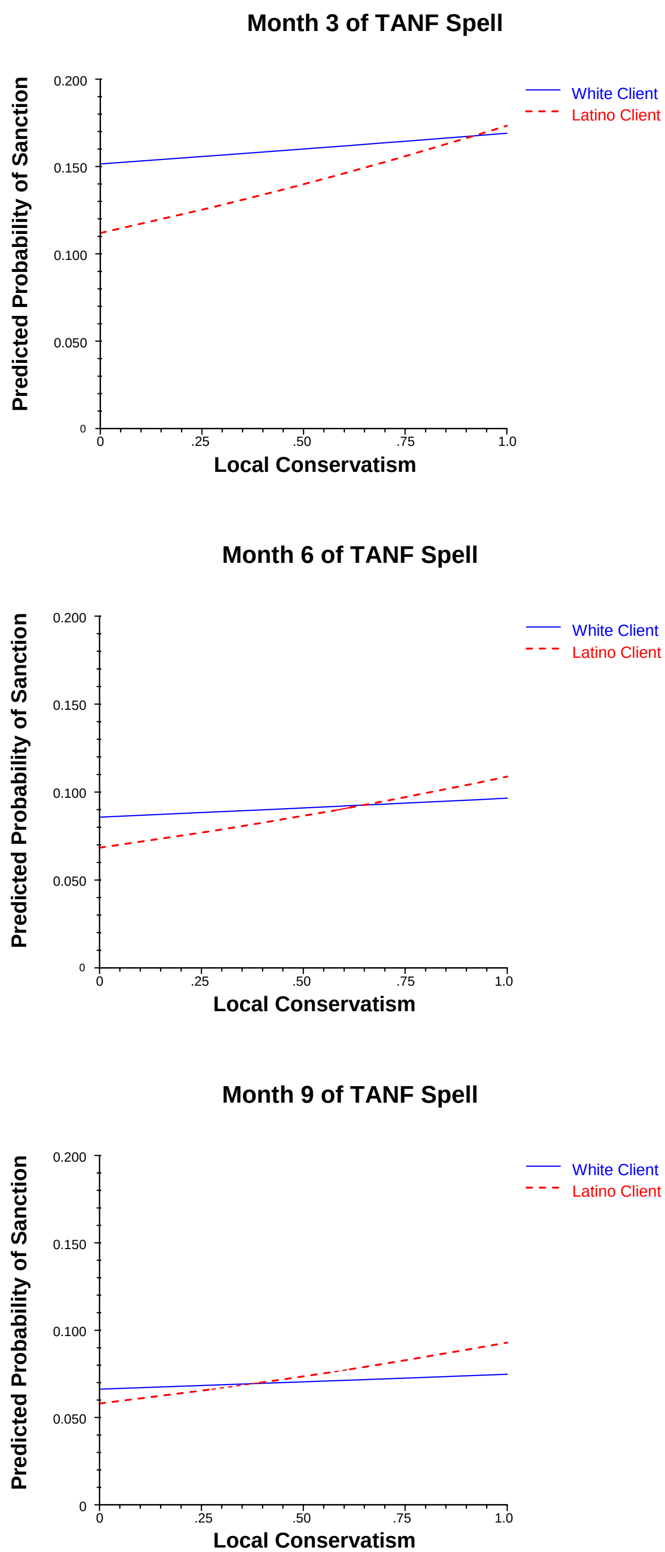
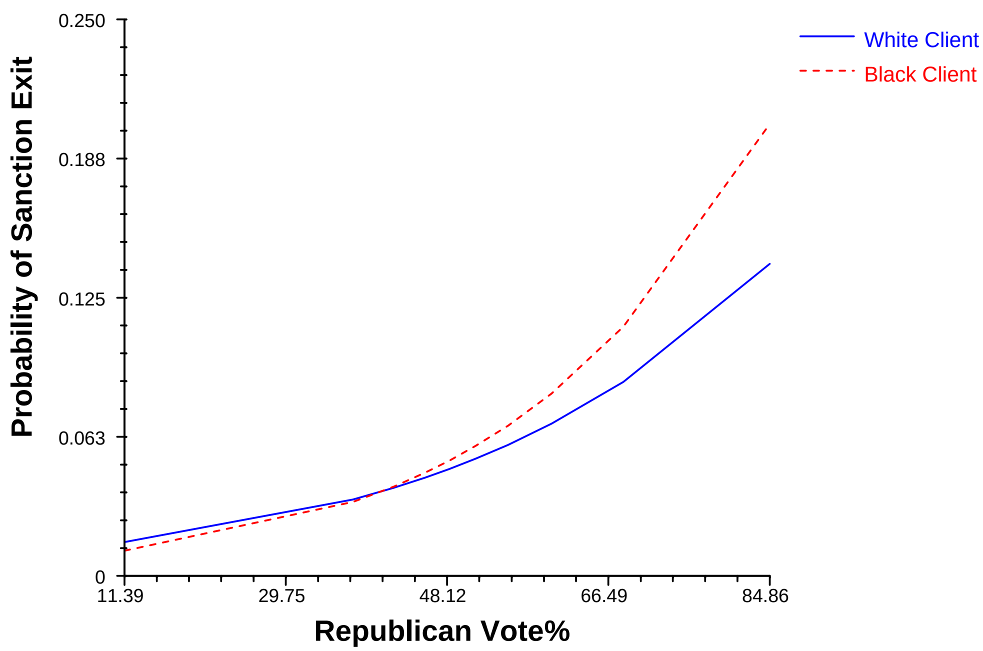


FIGURE 6. Predicted Probability of Sanction in Second-Order Devolution States, by Race of Client and Local Political Environment



Note: Predicted probabilities were calculated based on the results presented in Table 3, for the SOD sample of states.

Unpublished Appendix to

"Distributing Discipline:

Race, Politics, and Punishment at the Frontlines of Welfare Reform"

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Table A1. Variable Definitions, Sources, and Descriptive Statistics for Analyses Presented in Table 1

Independent Variables	Definition	Mean	S.D.	Minimum-Maximum
Individual characteristics:				
Gender	0 = female, 1 = male	.159	.366	0-1
Age	Client age (in years)	31.6	9.2	18-72
Marital status	1 = single parent, 0=otherwise, based on no. of adults in family	.751	.432	0-1
Number of children (ref. = 0 - 1):				
Two	1 = 2 children, 0 = otherwise	.278	.448	0-1
Three or more	1 = 3 or more, 0 = otherwise	.210	.408	0-1
Age of youngest child (ref. = 0 – 2 months):				
3 months – 2 years	1 = 3 months – 2 years, 0 = otherwise	.356	.479	0-1
3 – 4 years	1 = 3 – 4 years, 0 = otherwise	.125	.331	0-1
5 – 12 years	1 = 5 – 12 years, 0 = otherwise	.310	.463	0-1
More than 12 years	1 = more than 12 years, 0 = otherwise	.110	.312	0-1
Wage income	Wage income, from previous quarter, in 1,000s	.530	1.489	0-200
Education (ref. = more than 12 years):				
Less than H.S.	1= less than 12 years, 0 = otherwise	.472	.499	0-1
H.S.	1= 12 years, 0 = otherwise	.346	.476	0-1
Race or ethnicity (ref. = white, non-Latino):				
Black	1 = black, 0 = otherwise	.358	.480	0-1
Latino	1 = Latino, 0 = otherwise	.304	.460	0-1
Political environment:				
County conservatism index	See Appendix B	.465	.220	0-1
County black population (%)	Percentage of blacks in county of client in 2000 (County and City Data Books 2003)	16.2	6.9	2.1 – 57.1
County Latino population (%)	Percentage of Latinos in county of client in 2000 (County and City Data Books 2003)	22.9	21.7	1.5 – 57.3
Socioeconomic environment:				
Annual Wage in food service and drinking places	Average annual income in 1997 for employees in NAICS subsector 722, in 1,000s (County and City Data Books 2003)	12.9	1.9	7.8 – 16.7
County unemployment rate (t-1)	Unemployment rate in county of client, measured each month (Florida Research and Economic Database)	5.6	1.7	1.7 – 19.7
County poverty rate	County poverty rate for all persons in 2000 (U.S. Census Bureau Small Area Income and Poverty Estimates)	13.1	3.5	6.9 – 24.2
County TANF caseload (t-1)	Number of TANF recipients per 100,000 county residents (calculated by authors)	2.28	1.17	.142 - 6.907
County population (in millions)	Total county population in 2000, in 1000's (County and City Data Books 2003)	1072	825	7.02 - 2253

Source: Data on client characteristics were provided by the Florida Department of Children and Families.

Table A2. Variable Definitions, Sources, and Descriptive Statistics for Analyses Presented in Table 2

Independent Variables	Definition	Centralized States Mean	Centralized States S.D.	SOD States Mean	SOD States S.D.
Individual Level Variables					
Race/Ethnicity (reference=white)					
Black	1 = black, 0 = otherwise	0.283	0.450	0.405	0.491
Latino	1 = Latino, 0 = otherwise	0.109	0.312	0.158	0.365
Number of children (reference = 0 - one):					
Two	1 = 2 children, 0 = otherwise	0.280	0.449	0.278	0.448
Three or more	1 = 3 or more, 0 = otherwise	0.193	0.395	0.197	0.397
Age of youngest child (reference = 0 – 2 months):					
3 months – 2 years	1 = 3 months – 2 years, 0 = otherwise	0.371	0.483	0.377	0.485
3 – 4 years	1 = 3 – 4 years, 0 = otherwise	0.140	0.347	0.144	0.351
5 – 12 years	1 = 5 – 12 years, 0 = otherwise	0.285	0.452	0.281	0.450
More than 12 years	1 = more than 12 years, 0 = otherwise	0.101	0.302	0.094	0.292
Education (reference = <H.S.)					
H.S. Education	1= less than 12 years, 0 = otherwise	0.463	0.499	0.319	0.466
More than H.S. Education	1= 12 years, 0 = otherwise	0.050	0.219	0.041	0.198
Male	0 = female, 1 = male	0.160	0.367	0.234	0.423
Citizen	1 = Citizen, 0 = otherwise	0.956	0.206	0.943	0.232
Age	Client age (in years)	31.144	10.223	30.670	8.948
Single-Parent	1 = single parent, 0=otherwise	0.763	0.425	0.755	0.430
Earned Income	Amount of earned income in previous month	278.462	522.513	211.486	447.322
Public Housing	1 = Public housing recipient, 0 = otherwise	0.077	0.266	0.061	0.239
OASDI	1= Social Security recipient, 0 = otherwise	0.022	0.145	0.033	0.178
SSI	1 = SSI recipient, 0 = otherwise	0.010	0.100	0.036	0.188
County Level Variables					
Republican Vote%	Percentage of vote for Republican presidential candidate in general election, 1996 and 2000	44.372	11.765	43.955	13.224
Black%	Percentage of blacks in county of client in 2000 (County and City Data Books 2003)	11.658	16.000	16.349	17.606
Latino%	Percentage of Latinos in county of client in 2000 (County and City Data Books 2003)	7.174	9.257	11.895	15.738
Per Capita Income	County per capita income in 2000 (County and City Data Books 2003)	49.975	18.990	53.966	21.652
Unemployment Rate	County unemployment rate in 2000 (County and City Data Books 2003)	4.429	2.413	4.507	2.234

Source: Data on client characteristics are from TANF Administrative Dataset, available from ASPE.

Table A3. Construction of Index of County Political Ideology

To construct our index of local ideology we collected data on 18 ideologically relevant constitutional amendments that appeared on a statewide ballot for ratification from 1996 through 2004. We computed the percentage of "yes" votes for each amendment, for each county, and conducted a factor analysis using all 18 amendments (thus 18 variables, N=67 counties). The specific amendments are listed in the table below.

Ballot Title	*Election Year*	*Ballot Number*	*To Amend*
Should Two-Thirds Vote be Required for New Constitutionally-Imposed State Taxes/Fees?	1996	Constitutional Amendment 1	Art. XI, sec. 7
Fee on Everglades Sugar Production	1996	Constitutional Amendment 4	Art. VII, sec. 9
Responsibility for Paying Costs for Water Pollution Abatement in the Everglades	1996	Constitutional Amendment 5	Art. II, sec. 7
Preservation of the Death Penalty; United States Supreme Court Interpretation of Cruel and Unusual Punishment	1998	Constitutional Amendment 2	Art. I, sec. 17
Additional Homestead Tax Exemption	1998	Constitutional Amendment 3	Art. VII, sec. 6
Public Education of Children	1998	Constitutional Amendment 6	Art. IX, sec. 1
Basic Rights	1998	Constitutional Amendment 9	Art. I, sec. 2
Ballot Access, Public Campaign Financing, and Election Process Revisions	1998	Constitutional Amendment 11	Art. IV, sec. 5a; Art. VI, subsecs. 1,2,5,7; Art. IX, sec. 4a
Firearms Purchases: Local Option for Criminal History Records Check and Waiting Period	1998	Constitutional Amendment 12	Art. VIII, sec. 5
Florida Transportation Initiative for Statewide High Speed Monorail, Fixed Guideway of Magnetic Levitation System	2000	Constitutional Amendment 1	Art. X, sec. 19
Protect People from the Health Hazards of Second-Hand Tobacco Smoke by Prohibiting Workplace Smoking	2002	Constitutional Amendment 6	Art. X, sec. 20
Voluntary Universal Pre-Kindergarten Education	2002	Constitutional Amendment 8	Art. IX, sec. 1
Florida's Amendment to Reduce Class Size	2002	Constitutional Amendment 9	Art. IX, sec. 1
Animal Cruelty Amendment: Limiting Cruel and Inhumane Confinement of Pigs during Pregnancy	2002	Constitutional Amendment 10	Art. X, sec. 19
Parental Notification of a Minor's Termination of Pregnancy	2004	Constitutional Amendment 1	Art. X, sec. 22
Florida Minimum Wage Amendment	2004	Constitutional Amendment 5	Art. X
The Medical Liability Claimant's Compensation Amendment	2004	Constitutional Amendment 3	Art. I, sec. 26
Authorizes Miami-Dade and Broward County Voters to Approve Slot Machines in Parimutuel Facilities	2004	Constitutional Amendment 4	Art. X, sec. 19

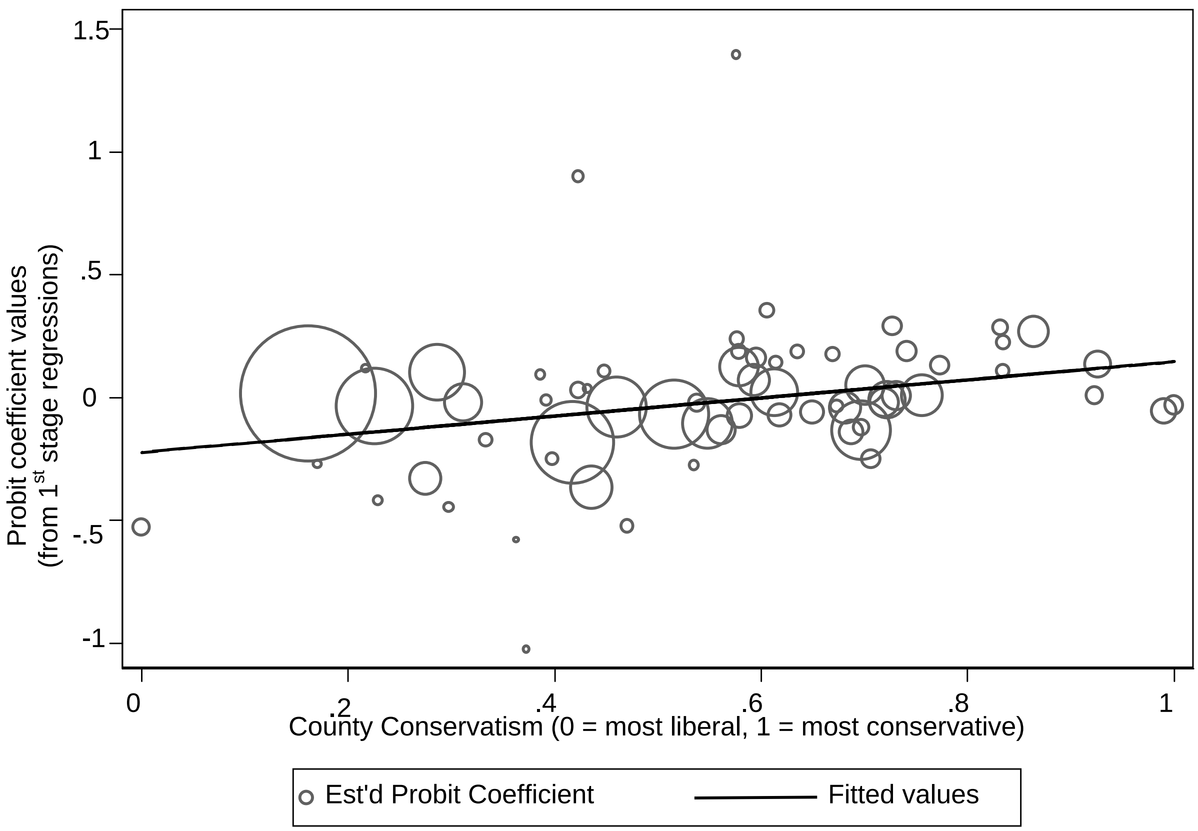
Table A4. Discrete-Time Event History Models of Sanction Initiation [Replication of Table 1 Using Logistic Regression with Clustered Standard Errors]

Independent Variables	Model I Coeff.	Model I Odds Ratio	Model II Coeff.	Model II Odds Ratio
Individual Characteristics				
Male	.1898**	1.2090	.1906**	1.2100
Citizen	.1535	1.1659	.1535	1.1600
Age	-.0129**	.9783	-.0219**	.9783
Single-Parent	-.1965**	.8216	-.1946**	.8231
Number of children (reference = 0 - one):				
Two	-.0191	.9810	-.0195	.9807
Three or more	-.0207	.9795	-.0206	.9796
Age of youngest child (reference = 0 – 2 months):				
3 months – 2 years	.5679**	1.7646	.5672**	1.7634
3 – 4 years	.6529**	1.9212	.6521**	1.9195
5 – 12 years	.7379**	2.0914	.7373**	2.0881
More than 12 years	.7596**	2.1373	.7586**	2.1353
Wage Income	-.0102*	.9898	-.0103*	.9898
Education (reference = >H.S.)				
Less than H.S. Education	.4764**	1.6103	.4784**	1.6135
H.S. Education	.1515**	1.1635	.1526**	1.1649
Race/Ethnicity (reference=white)				
Black	-.0470	.9541	-.0198	.9804
Black * Month of Spell	.0663**	1.0685	.0656**	1.0678
Hispanic	-.1486**	.8618	-.0883	.9154
Hispanic * Month of Spell	.0374**	1.0381	.0371**	1.0378
Political Environment				
Local Conservatism	.2000	1.2214	---	---
Local Conservatism*Black	.4761**	1.6098	---	---
Local Conservatism*Hispanic	.3716	1.4500	---	---
Republican Vote Share	---	---	.0003	1.0003
Republican Vote Share*Black	---	---	.0092**	1.0092
Republican Vote Share*Hispanic	---	---	.0112**	1.0113
Black%	-.0051	.9949	-.0060	.9940
Hispanic%	-.0081*	.9920	-.0096**	.9905
Socio-Economic Environment				
Annual Wage - Food Service/Drinking Places	-.0402	.9606	-.0451*	.9559
Unemployment Rate	.0032	1.0032	.0065	1.0066
ΔUnemployment Rate	-.0067	.9934	-.0038	.9962
Poverty Rate	.0395**	1.0403	.0347*	1.0353
Population	.0003**	1.0003	.0003**	1.0003
TANF Caseload	-.1785**	.8365	-191.0**	.8261
Number of Subjects	74,517		74,517	
Number of Failures	28,307		28,307	
Time at Risk (Person-Months)	198,147		198,147	

*p<.05, **p<.01

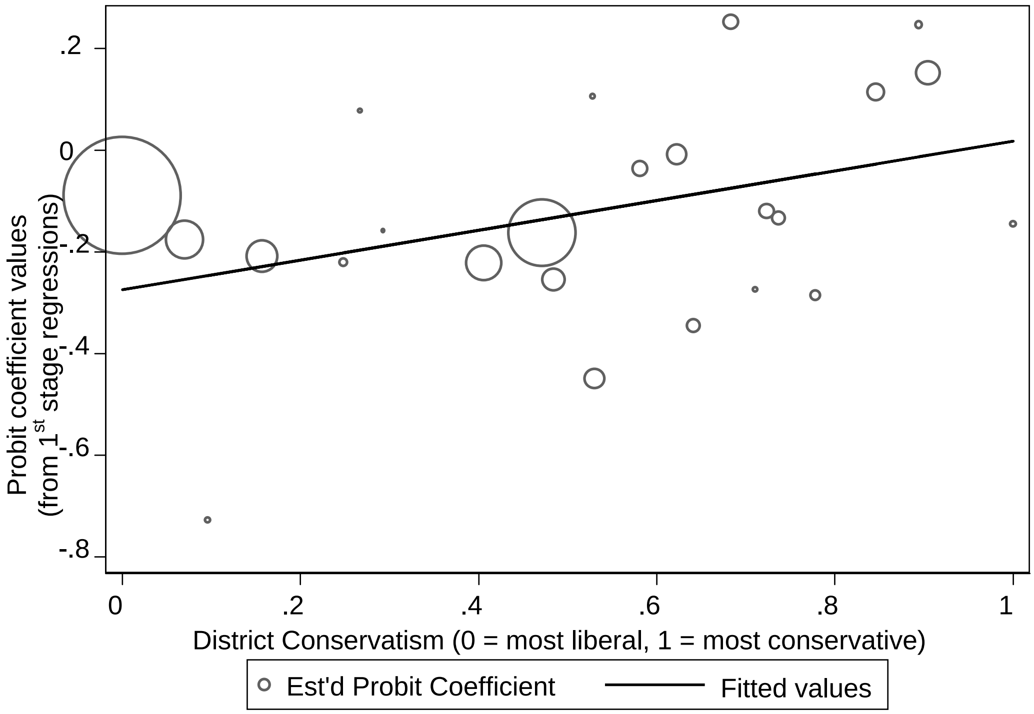
Note: Coefficient values are estimated using binary logit, with standard errors adjusted for clustering at the county level. The sample consists of all new adult TANF recipients in Florida who entered TANF from January 2001 through December 2002. Each model includes dummy variables for the duration of spell (months 1-12) and for calendar month (January-December) to capture potential seasonal effects. These results are not shown, but are available from the authors upon request.

FIGURE A1. The Effect of Race on TANF Sanctioning by Local Conservatism (Second-Stage Results from Two-Stage Multilevel Model)



Note: The vertical axis reflects probit coefficient values for the effect of race of client on TANF sanctioning (equivalent to the effect of *Black* from Table 1), obtained from individual-level (first-stage) probit regressions estimated for 66 Florida counties. The horizontal axis is the county conservatism index. The symbols in the scatterplot are weighted based on the sample sizes from the first-stage probit regressions. The slope of the regression line (b = .38, t = 2.88) was estimated using the weights proposed by Borjas and Sueyoshi (1994).

FIGURE A2. The Effect of Ethnicity (Hispanic) on TANF Sanctioning by District Conservatism (Second-Stage Results from Two-Stage Multilevel Model)



Note: The vertical axis reflects probit coefficient values for the effect of ethnicity of client on TANF sanctioning (equivalent to the effect of *Hispanic* from Table 1), obtained from individual-level (first-stage) probit regressions estimated for 24 workforce districts. The horizontal axis is the district conservatism index, calculated by averaging county conservatism scores within each district (weighting by county population). The symbols in the scatterplot are weighted based on the sample sizes from the first-stage probit regressions. The slope of the regression line (b = .29, t = 2.08) was estimated using the weights proposed by Borjas and Sueyoshi (1994).