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**Does Intolerance Dampen Dissent?
Macro-Tolerance and Protest in American Metropolitan Areas***

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ABSTRACT

Political tolerance has long been regarded as one of the most important democratic values because intolerant political cultures are believed to foster conformity and inhibit dissent.

Although widely endorsed, this theory has rarely been investigated. Using multilevel regression with poststratification to measure levels of macro-tolerance in U.S. metropolitan areas, and event data to measure rates of protest, we test whether cultures of intolerance do indeed inhibit public expressions of dissent. We find that they do: levels of macro-tolerance are positively and strongly associated with higher rates of protest in American metropolitan areas. Our findings have implications for the study of political tolerance, for normative theories of free speech and other civil liberties, and for scholarship on protest and collective action.

Keywords: Political tolerance; Dissent; Protest; Political Culture; Multilevel regression with poststratification

POLITICAL TOLERANCE is widely recognized as one of the most important democratic values (e.g., Sullivan, Piereson, and Marcus 1982). At its heart, it posits that every adult citizen ought to be allowed to participate fully in the political process, regardless of the content of her or his beliefs and opinions. Tolerant political cultures allow a wide range of ideas to be publicly expressed, freely circulated, and debated. Intolerant political cultures, in contrast, disapprove of the public expression of opinions that are judged to be objectionable. Scholars have long thought that this disapproval creates a creeping conformity and a reluctance to dissent. Conformist political cultures of this sort are widely believed to have adverse consequences for democratic politics (e.g., Gibson 1989; Gibson and Gouws 2003; McClosky 1964; Prothro and Grigg 1960; Sullivan, Piereson, and Marcus 1982).

However, despite a large literature addressing aspects of political tolerance, only a few studies have directly examined the question of whether intolerance does indeed constrain the public expression of political views. The primary purpose of this article is to investigate this rarely tested hypothesis. Our overriding expectation is that intolerant political cultures discourage dissent and reduce levels of protest. This is a cross-level rather than a micro-level conjecture: feeling free to express one's views and participate in politics is a function of the level of intolerance in the geographic context in which one lives (e.g., Gibson 1995). Gibson (1989, 1992, 2008) – relying on the work of Noelle-Neumann (1984) – has suggested that intolerance can create a spiral of silence in which individuals are reluctant to express their views because they fear the reactions of those around them. Damage to democratic deliberation can thus be dealt by ordinary citizens as well as repressive governments. When intolerance dominates a political culture, even majorities can be cowed into political quiescence (as in the infamous “silent majority” in the 1950s).

Our hypothesis implies a macro-level relationship between tolerance and protest, but macro-level measures of both of these variables have been lacking. Political tolerance is typically conceived of, and measured as an attribute of individuals, at the micro-level. In response, we use the method of multilevel regression with poststratification (MRP; Park, Gelman, and Bafumi 2004) along with existing survey datasets to measure the levels of political tolerance in 365 U.S. metropolitan areas over the period from 2007 to 2011. Protest data suffer from the same deficit: although scholars have previously gathered measures of protest incidence in American cities (e.g., Spilerman 1970; Olzak, Shanahan, and McEneaney 1996), contemporary data are scarce. As such, we use event data from the GDELT project to measure the rates of protest in U.S. metropolitan statistical areas over the same time period covered by our measures of tolerance. These new measures of political tolerance and protest in metropolitan areas allow us to test whether cultures of intolerance do in fact dampen dissent.

The article proceeds as follows. In the next section, we provide an account of why a tolerant political culture should facilitate dissent. The following section examines the extant empirical literature on protest, and explains why we think contextual levels of tolerance ought to also influence the incidence of protest. Then, we describe our measurement of macro-tolerance and rates of protest. Tests of our hypothesis show that macro-tolerance does indeed affect the rate at which citizens protest across U.S. metropolitan areas: a move from a below-average to an above-average level of macro-tolerance is associated with a striking 77 percent increase in the rate of protest. We conclude with some observations about tolerance and protest in contemporary world politics.

THEORY

Tolerance and Dissent

Democracy requires both majority rule and respect for minority rights. Particularly important are rights relating to expression and contestation (Dahl 1971), which strengthen democracy by facilitating citizen choice (Gibson 1992), encouraging deliberation and contestation (Ackerman and Fishkin 2004), and allowing all voices to be heard (Young 2000).

However, as the large literature on political tolerance has abundantly illustrated, many citizens in nearly all societies support limiting the rights of expression and contestation for certain unpopular groups (e.g., Erisen and Kentmen-Cin 2017; Duch and Gibson 1992; Gibson and Gouws 2003; Peffley and Rohrschneider 2003). In present day American society, examples of such groups are neo-Nazis, the Ku Klux Klan, and radical Muslims (Gibson 2008; Kalkan, Layman, and Uslaner 2009); in the 1950s, communists were similarly as disliked (Stouffer 1955). But more mundane groups also face opposition to the free expression of their points-of-view. For example, Gibson (2005) found that 48% of Americans are less than completely tolerant of atheists. And social commentators (e.g., Lukianoff and Haidt 2015) have argued that college campuses have become intolerant of conservative political views, particularly on cultural or “moral” issues.

There are two mechanisms whereby intolerance may hinder collective action and protest (Gibson 1989; 1992; 2008). First, public intolerance toward a group may spur the government to pass legislation directed against the activities of that group. This has happened many times in American history (see Goldstein 1978). But of arguably greater importance is the direct effect of public intolerance. A climate of intolerance may discourage citizens from dissenting or expressing unpopular political views. And the more these views are left unexpressed, the less popular they appear to be. This dynamic creates what Noelle-Neumann (1984) refers to as a

“spiral of silence,” in which people learn that it is simply best to “keep their mouths shut.”

Moreover, as the logic of the spiral metaphor suggests, an intolerant political climate can result in the silencing not only of extreme and odious views, but other minority and dissenting voices as well.

According to the spiral of silence theory, political tolerance works as a macro-level force that shapes political outcomes: it is a *culture of intolerance* surrounding those with dissenting opinions that matters for whether these views are expressed (Gibson 1995). Yet despite this theoretical justification for conceptualizing tolerance as a macro-level variable, existing literature almost entirely treats political tolerance as a micro-level attribute, and focuses mainly on its conceptualization, measurement, and origins (for a review, see Gibson 2006). Only a few studies have examined the consequences of cultures of intolerance.

For example, Gibson (1988) found that state-level political tolerance is unrelated to government repression of communists, while Gibson (1989) found a surprising positive correlation between levels of tolerance and state restrictions on campus protest during the Vietnam era. He explained this positive effect of tolerance on repression by noting that tolerance increased the incidence of campus protest, which then generated subsequent backlash from state governments.

Gibson (1992) also examined the effects of contextual tolerance (the level of tolerance in the respondent’s immediate community) on individual perceptions of freedom of expression in their communities, finding a positive relationship. In a later study (1995), he showed that this positive effect of tolerance on perceived freedom holds even for black Americans when tolerance includes tolerance of racists. These results provide evidence that tolerant communities do indeed increase the perceived freedom of individuals living therein.

We extend that line of research by examining the effects of intolerant contexts on macro-level political behavior rather than on micro-level perceptions. If we draw out the logic of the spiral of silence to its full extent, the theory holds that intolerant cultures should be expected to dampen not only micro-level *proclivities* to dissent, but macro-level *occurrences* of dissent as well. In particular, we hypothesize that the level of political intolerance of a geographic area should be negatively related to the rate at which citizens living in that area engage in political protest.

If not at the micro-level, then which level of aggregation is best suited to testing our hypothesis? The effects of cultures of intolerance typically play out on a wider stage than neighborhoods, but require a smaller scale than states. Although there is a long tradition of research into the political cultures of the U.S. states, considerable heterogeneity exists among them, as election-watchers frequently remind us. We ultimately settle on the metropolitan statistical area (MSA) as our unit of analysis. Although these are fairly large areas, spanning many counties in the largest cases, they are communities that exist in the perceptions, imaginations, and identities of most of their inhabitants – indeed, by definition, MSAs are communities that exist even in the commuting behaviors of their inhabitants.¹ Because their inhabitants interact behaviorally, we expect that MSAs should be propitious geographic units for examining the effects that people’s cultural values play in constraining or facilitating the

¹ MSAs are the counties adjacent to an urbanized core area of at least 50,000 people that have “a high degree of social and economic integration with the core as measured through commuting ties” (https://www.census.gov/geo/reference/gtc/gtc_cbsa.html, accessed 10/19/2017).

behavior of others.²

Tolerance and Protest

The most prominent way through which citizens of democracies can dissent is to participate in or organize protest. A long tradition of research in political behavior distinguishes between conventional, institutionalized actions related to the electoral process – such as volunteering for a campaign, contacting a politician, or voting – from non-routine, “unconventional” forms of political action – such as demonstrations, strikes, vigils, marches, and boycotts (e.g., Barnes and Kaase 1979; Dalton 2004; Muller, Jukam, and Seligson 1982). Although one may cast a “protest vote,” the act of voting itself is entirely orthodox and culturally appropriate. Unconventional political action is different. Not only is a protest a manifestation of disagreement over policy or procedure, but these acts are themselves challenging and contentious, and publicly so.

Scholars have identified two sets of factors that work as determinants of protest incidence and protest participation: those that “push” individuals into protest and those that “pull” them in (McAdam 1986). Push factors drive individuals to organize and take part in protest by shaping their motivations, which can galvanize a large pool of would-be participants. Perhaps the longest-standing push factor associated with protest is grievances (Muller, Jukam, and Seligson

² MSAs are also an attractive unit of analysis for efforts to explain protest incidence inasmuch as the study of protest in the U.S. – if not also protest itself – has primarily been an urban preoccupation. Significant studies that use cities or metropolitan statistical areas as the units of analysis include Eisinger (1973), Okamoto (2003), Olzak and Shanahan (1996), Olzak, Shanahan, and McEneaney (1996), and Spilerman (1970).

1982), including both absolute (Massey and Denton 1993) and relative deprivation (Eisinger 1973). Heighted competition over resources, especially between ethnic groups, has also been shown to trigger protest (Okamoto 2003; Olzak, Shanahan, and McEneaney 1996). Researchers have furthermore shown that left or liberal ideological leanings also push individuals to participate in protest (Dalton, Van Sickle, and Weldon 2010; Schussman and Soule 2005; Verba, Schlozman, and Brady 1995).

Pull factors, in contrast, represent structural forces that promote the organization of or participation in protest. They facilitate collective action by reducing its costs. Pull factors might take the form of individual resources, such as education or money, that enable individual protest participation (Schlozman, Verba, and Brady 2012). Or they might take the form of networks that aid in the organization of protest and the recruitment of participants (McAdam 1986). Two likely network sources are civic organizations (Putnam 2000) and certain types of religious congregations (McAdam 1986; Schlozman, Verba, and Brady 2012). Macro-tolerance is clearly a pull, rather than a push, factor because, according to the theory of the spiral of silence, macro-tolerance facilitates (or inhibits) protest by increasing (or decreasing) perceptions of having the freedom to dissent.

We are aware of only one study that attempts to test the effects of tolerance on protest. Rapp and Ackermann (2016) show that, while individual-level “social intolerance” increases political participation, national-level intolerance decreases participation. This national-level effect is consistent with the hypothesis we are testing. However, we offer a clearer test of the hypothesis because we use metropolitan area political intolerance as our independent variable. Political tolerance is conceptually closer to the “culture of conformity” described by scholars such McClosky (1964) and Sullivan, Piereson, and Marcus (1982) than it is to social intolerance

(Gibson 2006), and the metropolitan level of analysis is geographically more proximate to the individuals who decide whether or not to protest.

Finally, a fair amount of research has linked protest to other aspects of political culture. Inglehart (1977), for example, showed that post-material values were associated with the rise of new social and protest movements in Western democracies during the 1960s and 1970s. Muller, Jukam, and Seligson (1982) found that protest participation is triggered by a lack of diffuse support for the political system, which they call “political alienation.”

Yet these explanations differ fundamentally from the hypothesis we are proposing. The theories of these and other democratic theorists are best regarded as *individual* ideological or value commitments that *push* people into protest. In contrast, we regard tolerance as a contextual factor that pulls individuals into protest. We now turn to our strategy for measuring contextual levels of tolerance – or “macro-tolerance,” as we refer to it.

MEASURING MACRO-TOLERANCE

The measurement of public opinion within small geographic areas is best accomplished using a large-sample survey with sufficient respondents in each area. Surveys of this type (such as the Cooperative Congressional Election Study) are, however, hugely expensive, and thus rare. Nor have they included measures of political tolerance.

Another approach has been to disaggregate an existing survey sample into geographic units, perhaps combining several survey samples together before doing so. Following the pioneering work of Erikson, Wright, and McIver (1993), Gibson used this method to estimate state-level tolerance by disaggregating the Stouffer (1955) and Nunn, Crockett, and Williams (1978) national surveys. Brace et al. (2002) followed suit by disaggregating 24 years of General

Social Survey data. However, even with large pooled samples, Lax and Phillips (2009) and Pacheco (2011) have shown that estimates produced using this method of disaggregation are unreliable in small states. These difficulties are of course only compounded when one is interested, as we are, in measuring opinion in geographic areas smaller than states.

Faced with these limitations, scholars are increasingly turning to a method of model-based estimation known as multilevel regression with poststratification (MRP). Pioneered by Gelman and Little (1997), MRP utilizes a multilevel model to predict the average opinion of thousands of demographic-geographic categories (such as Hispanic, college-educated women who are aged 18-29 and live in Colorado). These raw opinion estimates are then weighted by the population sizes of each demographic-geographic category and aggregated by geographic unit to create a post-stratified estimate of opinion.

MRP has been extensively used for estimating state-level opinion. These estimates have furthermore been validated using a variety of external data, including election results (Park, Gelman, and Bafumi 2004) and large aggregated polls (Lax and Phillips 2009; Pacheco 2011). In particular, Lax and Phillips (2009) show that MRP can reliably measure state-level opinion using national survey samples of as few as 1,500 respondents.

Few scholars have attempted to use MRP to measure opinions in areas smaller than states, which is what we intend to do. An important exception is Warshaw and Rodden (2012), who test the accuracy of MRP opinion estimates for congressional and state legislative districts compared to both a large, aggregated poll and state referendum results. They conclude that “MRP produces very reliable estimates of congressional districts’ public opinion with a national sample of just 2,500 people, and it yields reliable estimates for state senate districts with a national sample of 5,000 people” (204).

In 2010, the population of the average congressional district (711,000) was similar to the population of the average MSA (692,000). However, MSAs are more variable in size than congressional districts: the smallest MSA (Carson City, NV) has just 55,000 inhabitants, compared with the smallest congressional district (Rhode Island), which has 528,000. Nevertheless, with our sample of survey data ($N = 3,133$) being approximately 25% larger than the minimum ($N = 2,500$) recommended by Warshaw and Rodden (2012) for modeling opinion within congressional districts, we conclude that we can indeed use MRP to measure macro-tolerance within MSAs. We describe our approach below.

Survey Data

Our survey data come from the Freedom and Tolerance Surveys (FATS), which were conducted every year between 2007 and 2011. Since we do not expect a measure of political culture, such as macro-tolerance, to vary much from year to year,³ we aggregate the five annual surveys to produce one sample. Respondents were sampled from all 48 contiguous U.S. states and from 317 out of the 366 metropolitan statistical areas. The sample sizes within MSAs ranged from 1 (various) to 177 (New York-Northern New Jersey-Long Island), with a mean of 10.3. The 87 respondents who refused to provide information for key demographics and the 817 additional respondents who lived outside an MSA were excluded from the dataset. We also removed another 65 respondents who lived in the Washington-Arlington-Alexandria, DC-VA-MD-WV

³ Using a one-way ANOVA, we find no statistically significant difference across survey years in micro-level political tolerance ($F = .91$, $df = 4$, $p = .46$). See below for the measurement of micro-tolerance.

MSA because it proved to have extremely high levels of protest, which is likely due, in part, to measurement error.⁴ We are left with a sample of 3,133 respondents for our analysis.

The MRP procedure requires cross-classifiable population data on the demographic and geographic variables of interest. The joint demographic-geographic distributions are required, not merely the marginal distributions for each of these variables. However, the Census Bureau provides the joint distributions of only three demographic variables at the MSA level. This creates a problem for our methodology as there are at least four demographic variables we would like to include as predictors and post-stratification factors: age, education, race, and gender.

To overcome this obstacle, we use the method of Leemann and Wasserfallen (2017), who propose creating a *simulated* joint distribution matrix using combinations of marginal and joint distributions.⁵ Specifically, we obtain the estimates of the joint gender-age-education-MSA

⁴ We suspect that GDELT incorrectly classifies news reports as describing a protest in the Washington-Arlington-Alexandria, DC-VA-MD-WV MSA to a far greater degree than is the case for other MSAs. The reason is that foreign policy reporters occasionally refer to international interactions using national capitals, e.g., “Washington protested the actions of Teheran.” Such false positives affect only the nation’s capital among the MSAs and are likely responsible for a considerable proportion of the protests we counted as occurring between 2007 and 2011 using the GDELT data. Indeed, even after applying our filters to remove false positives, we count 1,261 protest events per year in Washington DC MSA, compared with 306 per year in San Francisco, which is otherwise our most protest-prone MSA.

⁵ Leeman and Wasserfallen (2017) point out that MRP can be conducted when one knows only the marginal distributions of demographic and geographic populations. To do so, one

distributions from the 2007-2011 American Community Surveys, and then combine these estimates with the joint race-MSA distributions from the same source to create a simulated joint distribution matrix of all five demographic and geographic variables. We ensure that these variables are identically coded in both this post-stratification dataset and in the survey dataset.

Measuring Political Tolerance at the Micro-Level

MRP has thus far been applied to a fairly limited range of opinions: electoral preferences (Gelman and Little 1997; Ghitza and Gelman 2013; Park, Gelman, and Bafumi 2004), partisanship (Pacheco 2011), policy preferences (Lax and Phillips 2009; Warshaw and Rodden 2012), and policy “mood” or ideology (Enns and Koch 2013). While electoral and policy preferences are typically measured with single dichotomous indicators (but see Enns 2016), multiple indicators are required to reliably and validly measure latent constructs such as political tolerance. Before estimating tolerance at the metropolitan level, we need to first measure it at the individual level.

Gibson (2013) describes three approaches to measuring political tolerance: the least-liked, fixed-group, and support for civil liberties methods.⁶ All three have been used to measure

creates a simulated joint distribution under the assumption that the proportion of MSA residents who are, for example, aged over 65, is constant across the categories of the other demographic factor(s). Leeman and Wasserfallen show that this approach produces estimates as accurate as those obtained using classic MRP, especially if demographic factors are only weakly inter-correlated.

⁶ Briefly, fixed-group measures of tolerance are those in which the questions refer to

“the degree to which citizens will support the extension of civil liberties to all, including groups advocating highly disagreeable viewpoints and ideologies” (Gibson 2013, 46). This suggests that the three approaches could be regarded as tapping a single latent construct we might call “general political tolerance.” However, Gibson (2013) also finds that correlations between the three additive scales are modest, which casts some doubt on the assumption of a single latent construct.

Our approach is something of a compromise between the various measurement techniques. The fixed-group approach is ill-suited for measuring tolerance at the state-level because it asks about groups, such as “people who are against all churches and religion” and “religious fundamentalists,” with which some respondents might actually identify, and toward which others might even be sympathetic. Such groups would not satisfy the “objection precondition” required to validly measure political tolerance. In addition, the extent to which individuals identify with or feel sympathetic toward such groups likely varies considerably by metropolitan area. We therefore do not consider the fixed-group approach further, and instead focus on the question of whether the least-liked and support for civil liberties items can be used together to validly and reliably measure general political tolerance.

Following Sullivan, Piereson, and Marcus (1982), the least-liked items ask respondents

political groups selected by the researcher (e.g., Stouffer 1955). Least-liked measures allow the respondents themselves to identify the groups (e.g., Sullivan, Piereson, and Marcus 1982).

Support for civil liberties refers to attitudes about whether certain types of activities, irrespective of the groups involved, ought to be allowed (e.g., giving certain types of inflammatory speeches; see Davis 2007 and Gibson and Bingham 1985).

whether they would allow three acts of political participation (making a speech, running for office, and holding a demonstration) by groups that they dislike a great deal. The five civil liberties items ask respondents about their support for policies such as the institution of a national identity card, government monitoring of citizens' communications, and police investigations of protest participants.

To test the reliability and validity of a two-factor model, we specify and fit an ordinal confirmatory factor analysis (CFA), with one factor specified for the three least-liked items and the other for the five civil-liberties items. These two factors are then modelled using a higher-order, general tolerance factor. The model fits the data fairly well (Comparative Fit Index = .992; RMSEA = .059; full results are reported in the Online Appendix, Table A5). The two tolerance scales – least-liked and support for civil liberties – thus clearly appear to tap a latent construct of general political tolerance that combines attitudes toward tolerance of everyone with attitudes toward tolerance of hated groups. We therefore use the general tolerance factor scores from the CFA as our outcome variable in the MRP model.⁷

The MRP Model

MRP estimation begins with a multilevel model of individual opinion. Our model is similar to those used in existing research (e.g., Lax and Phillips 2009, Warshaw and Rodden 2012); we model tolerance as a function of a number of demographic and geographic categories (gender, race, age group, level of education, and MSA – further details on the MRP model are available in

⁷ Cronbach's alpha for the entire set of eight items is .72, which indicates that even a simple summated index of the items would have adequate reliability.

the Online Appendix). Our approach is also typical in that the MSA effects are nested within a higher-order geography (census region in our case) and include MSA-level covariates (the percentages of the MSA population that are black, religious, and hold a bachelor's degree).

This model is used to predict levels of political tolerance within each of the 73,000 demographic-geographic "types." The estimate for each type is then weighted by the proportion of the MSA's population that falls within that demographic-geographic type. These estimates are then aggregated by MSA to yield a post-stratified measure of metropolitan-level political tolerance. They are finally transformed to range from 0 to 1.

The resulting measure of macro-tolerance displays a regional pattern within the U.S., with intolerance pervasive in southern MSAs and tolerance generally increasing as one moves north (we include a map of the estimates in the Online Appendix). This geographic pattern is largely consistent with existing state-level estimates of macro-tolerance (Brace et al. 2002; Gibson 1988; 1989). The most tolerant MSAs include northern college towns such as Boulder, CO, Ithaca, NY, and Corvallis, OR. The least tolerant are smaller southern MSAs such as Sumter, SC and Albany, GA, as well as a somewhat larger MSAs in Texas such as Brownsville-Harlingen and McAllen-Edinburg-Mission.

Macro-tolerance is quite distinct from the partisan and ideological orientations that animate the familiar "red-state, blue-state" divide. It is correlated at .32 with the percentage of the MSA that voted for Obama in 2012, for example. While liberal MSAs are somewhat more tolerant, intolerance can and does exist on both the left and right. As such, we can be sure that any observed relationship between macro-tolerance and protest is not merely another manifestation of partisan and ideological divides.

MEASURING PROTEST INCIDENCE

We use data from an automated event data-collection project, the Global Database on Language and Tone (GDEL T), to measure the incidence of protest in MSAs. GDEL T codes political events beginning in 1979 and continuing until the present day, using a corpus of news sources including *Agence France Presse*, *Associated Press*, *Google News*, *New York Times*, and *Xinhua*. Events are identified, parsed, and linked across news stories, with the type and location of action and the type and identities of actors recorded to the greatest extent possible (see Leetaru and Schrod t 2013 for further details).

Scholars have criticized GDEL T event data for its accuracy (e.g., Hammond and Weidmann 2014). Raw GDEL T event counts are thought to include a large number of false positives. The program either misinterprets the language in some news reports, falsely counting an event where none existed, or it fails to link together separate news reports of a single event, thus counting one event more than once (Leetaru and Schrod t 2013).⁸ This propensity toward false positives may have increased over time (Ward et al. 2013).

A few researchers have evaluated such concerns by comparing GDEL T counts of protests to data obtained from other sources. Hanna (2014), for example, compares the monthly GDEL T count of the incidence of protests in the U.S. between 1979 and 1995 with hand-collected data over the same period from the Dynamics of Collective Action (DCA) project (see Earl, Soule, and McCarthy 2003 for further details on this project). She finds that the two time-series are poorly correlated ($r = -.11$). Ward et al. (2013), on the other hand, find that GDEL T counts of

⁸ Ward et al. (2013) provide the illustration of international trade or currency “wars” that might be counted as conflict events.

daily protests in Egypt during the months of November 2011 and November 2012 are strongly correlated ($r = .84$) with the counts of the Integrated Crisis Early Warning System, another automated event-coding dataset. These two tests both examine the longitudinal accuracy of GDELT protest data, but, as we have seen, they arrive at very different conclusions.

Our research design is, however, cross-sectional. We have reason to suspect that GDELT provides better cross-sectional measures of protest incidence than it does longitudinal measures, because the false positive rate might be expected to occur in all cross-sectional units under investigation. To verify this intuition, we conducted our own validation exercise. We follow Hanna (2014) in comparing GDELT and DCA data. We diverge in that we are interested in cross-sectional, not longitudinal correlations. Because DCA does not include information on the MSA in which a protest took place, we instead measure and compare the state rates of protest between the two sources of data. Specifically, we compare the total incidence of protest for each of the 48 contiguous states obtained from GDELT with those obtained from the DCA dataset. The timeframe is 1979 to 1995, both inclusive, which is the period that GDELT and DCA overlap. The Spearman's rank-order correlation between the GDELT and DCA measures of state incidence of protest is .80.⁹ This is a very encouraging finding given the different sources of data and slightly different definitions of collective action used in each.

Since it appears that GDELT data can be used to reliably estimate protest incidence in U.S. states, there should be no special problems with using these data to measure protest

⁹ This correlation is obtained using the "filtered" GDELT data; the method of filtering is described in the next paragraph. Using the raw, unfiltered GDELT protest counts, the correlation with the DCA counts is even higher, at .83.

incidence in metropolitan areas.¹⁰ Nevertheless, in an effort to further reduce the prevalence of false positives, we filter the raw GDELT data in various ways. We remove would-be protest events that (1) are not mentioned in the first paragraph of a news report, (2) are mentioned in only one news report, and (3) code the protagonist as being a government, media, or international actor. In addition, events are automatically removed if they lack the necessary geo-code to be located within an MSA. The resulting, “filtered” annual incidence of protest is our main dependent variable.

According to this measure, the annual incidence of protest between 2007 and 2011 ranged from 0 (in seven MSAs: Anderson, IN; Coeur d’Alene, ID; Elkhart-Goshen, IN; Idaho Falls, ID; Jackson, TN; Midland, TX; and Winchester, VA-WV) to 306 in San Francisco-Oakland-Fremont. Incidence, however, is partly a function of the size of the population of the metropolitan area. It is therefore necessary to adjust for the increasing opportunities for protest associated with larger populations by modelling the rate of protest rather than the raw incidence. There are two ways to do this. The first, and perhaps simplest, option is to calculate the rate directly by dividing protest incidence by the population of the MSA. Since the resulting rate is highly skewed, it is desirable to then apply a logarithmic transformation. And since a few MSAs exhibit no protest activity in our five-year window, one must add a small constant before applying the logarithm. We use this method of adjusting our measure of protest when plotting the

¹⁰ Indeed, we might expect that MSAs, being urban areas, exhibit more homogenous levels of media density than states, which vary greatly in their degree of urbanization. As such, newspaper based measures of protest incidence might in fact be more accurate at the MSA-level than the state-level.

data, where continuous measures are more revealing.

When modeling the determinants of protest in regression framework, however, we use a second approach to adjust protest incidence for the increasing opportunities for collective action that come with larger populations. Here, we use negative binomial regression models with incidence of protest as the outcome variable and the log of MSA population included as an offset.¹¹ This transforms the model into a negative binomial rate model (e.g., Gelman and Hill 2007). We prefer the negative binomial rather than the simpler Poisson count model because the incidence of protest is overdispersed, with a variance far in excess of the mean. The negative binomial model includes an additional dispersion parameter, which allows for accurate modeling of such overdispersed count data.

Examining the rates of protest (i.e., incidents of protest per capita), we find that the most protest-prone MSAs are Carson City, NV, and Topeka, KS, with 98 and 80 protests per million residents per year. San Francisco-Oakland-Fremont, CA, the first large urban area on the list, has the third highest rate of protest (71 protests per million residents per year). The seven MSAs with no recorded protests between 2007 and 2011 naturally also have zero rates of protest, and they take up the last seven places on our rankings.

TESTING THE EFFECT OF MACRO-TOLERANCE ON PROTEST

¹¹ Offsets are variables included in regression models with coefficients fixed to a value of one. Doing so in this case allows the log of population to be moved to the left-hand side of the equation. In turn, this permits an interpretation of the model as one of (logged) rate of protest rather than protest incidence, which is arguably a more intuitive interpretation.

We turn now to the primary contribution of this article: testing whether macro-tolerance influences the rate of protest in American metropolitan areas. We begin by examining a scatterplot that depicts the bivariate relationship between macro-tolerance and rate of protest (Figure 1). The scatterplot reveals a positive and moderate relationship between macro-tolerance and protest ($r = .35$): the rate of protest, in other words, is generally higher in metropolitan areas that have more tolerant political cultures, just as we hypothesized.

[PLACE FIGURE 1 ABOUT HERE]

To address the question of whether this relationship persists once we adjust for other known determinants of protest, we specify a negative binomial model of the rate of protest, as shown in Table 1. The primary purpose of this model is to test the robustness of the association between tolerance and protest when we adjust for other factors scholars have found that are also determinants of protest participation and protest incidence. Although we include measures of a number of other determinants of protest, their role here is primarily to serve as controls for possible factors confounding the relationship between the rate of protest and macro-tolerance.

[PLACE TABLE 1 ABOUT HERE]

Our control variables include factors that “push” people into protest, as well as those that “pull” them in. First, among the push factors, is a measure of the ideological color of the MSA: the percentage of the 2012 presidential vote Barack Obama won in the area. Analysis of survey data has shown that a left or liberal ideology is associated with protest participation (Schussman and Soule 2005; Verba, Schlozman, and Brady 1995), although a similar connection has not yet been established at the aggregate level. We also include the Gini coefficient as a measure of the areas’ income inequality, which has long been proposed (but seldom demonstrated) as a determinant of civil conflict, if not necessarily protest (Lichbach 1989). Next, as measures of

heightened competition over material resources in the MSA, we include the average unemployment rate over the 2007 to 2011 period and the rate of population growth between 2000 and 2010 (Olzak and Shanahan 1996).

Our other indicators of push factors relate to processes of competition and threat between racial and ethnic groups. The ethnic fractionalization index measures the diversity of the MSA; we would expect protest to be positively associated with ethnic diversity to the extent that minority groups express greater demands as their numbers grow (e.g., Olzak, Shanahan, and McEneaney 1996; Spilerman 1970) or if the majority group becomes increasingly threatened by an growing minority population (Enos 2016; Giles and Evans 1986). We also include the black-white dissimilarity index, which is used to measure the degree of racial segregation of neighborhoods within a MSA. In more segregated urban areas, groups find each other more threatening, all else being equal (Olzak, Shanahan, and McEneaney 1996; Uslaner 2012).

Moving on to pull factors, we include median household income and the percentage of the MSA with a high school diploma as measures of the supply of financial resources and civic skills, which, as Verba, Schlozman, and Brady (1995; Schlozman, Verba, and Brady 2012) have shown, are related to political activism. Protest also requires a supply of people who are “biographically available” to participate (McAdam 1986). We include the percentage of students in the MSA because students have both the time and the inclination for protest (Schussman and Soule 2005). We also include the percentage of the population that is under 18 years of age, since MSAs with greater proportions of children also have greater proportions of parents, who are typically less available to protest (Verba, Schlozman, and Brady 1995). Finally, voluntary organizations have repeatedly been shown to be crucial in recruiting and mobilizing protest participants (McAdam 1982; Putnam 2000). We use data on the number of charitable

organizations and religious congregations per capita.

The first model in Table 1 shows that the positive and significant bivariate relationship between macro-tolerance and protest persists when we control for all these other “push and pull” determinants of protest incidence. The effect of macro-tolerance is substantially significant: a two-standard deviation increase in macro-tolerance is associated with a 77 percent increase in the metropolitan rate of protest.

To illustrate the magnitude of these associations between macro-tolerance and incidence of protest – and to show how they vary as MSA size increases – we plot the predicted effects of macro-tolerance in Figure 2. This figure shows the predicted incidence of protest across the full range of our macro-tolerance variable for two hypothetical MSAs: one a standard deviation above the mean MSA population (or approximately the size of Bridgeport-Stamford-Norwalk, CT); the other, a standard deviation below the mean MSA population (or approximately the size of San Angelo, TX). For the larger MSA, an increase in macro-tolerance from one standard deviation below the mean to one standard deviation above implies an increase from 8.2 to 14.6 protest incidents a year. For the smaller MSA, a similar increase in macro-tolerance corresponds to an increase from 1.0 to 1.8 protest incidents a year.

[PLACE FIGURE 2 ABOUT HERE]

Some of the other covariates in Model 1 also show significant effects. First, ethnic fractionalization has a positive association with the rate of protest, echoing Olzak and Shanahan’s (1996) finding that the size of the non-white population is related to the occurrence of “race riots.” Our result differs in that our measure of protest incidence includes *all* forms of protest, race-related or not. As such, our finding additionally calls to mind Hero and Tolbert’s (1996) argument regarding the centrality of ethnic diversity in affecting state politics. Ethnic

diversity appears to be a key variable for understanding MSA political behavior as well.

Second, the share of the MSA electorate that voted for Obama in 2012 has a positive relationship with the incidence of protest. This effect shows that contextual ideology influences dissent, with liberal metropolitan areas being more likely to protest than conservative ones. Past research has shown the same effect but at the individual level – i.e., left-leaning individuals are more likely to participate in protest (Dalton, Van Sickle, and Weldon 2010; Schussman and Soule 2005). This study is the first to show such an effect for metropolitan protest incidence.

Third, as expected, the share of the MSA population under 18 years of age is negatively related to the incidence of protest. This is likely due to two separate effects: children seldom participate in protest themselves and, as such, MSAs with greater proportions of children have smaller pools of potential protestors. In addition, the presence of more children implies the presence of more parents, and as Verba, Scholzman and Brady (1995) find, having children reduces one's free time and thus also limits political participation.

Finally, we find that the number of charitable organizations per capita is associated with protest activity. A plentiful supply of such organizations provides at least two channels for protest mobilization. First, voluntary organizations teach their members civic skills, e.g., running a meeting or making a speech (Verba, Scholzman, and Brady 1995; Putnam 2000). These skills can then be used to organize and sustain a protest. Second, voluntary associations help facilitate protest participation because they create horizontal ties between members (McAdam 1986). Indeed, perhaps the most powerful predictor of protest participation is simply being asked directly by someone else to take part (Schussman and Soule 2005; Gibson 1997).

However, the most important finding from Model 1 is that macro-tolerance retains a positive and significant effect even in a well-specified model (and across other specifications of

the model).¹² Having controlled for some potential correlates of protest and state politics, we can conclude that the relationship between macro-tolerance and protest is reasonably robust.

Nevertheless, there is still reason to doubt whether these results reflect a causal effect of macro-tolerance on protest. In particular, one might suspect that macro-tolerance may be endogenous to levels of protest. On the one hand, a heightened incidence of protest, particularly if it is aggressive or violent, might dampen popular tolerance for future protest (e.g., Gibson 1989). According to this story, protest can actually *decrease* tolerance via a backlash effect. This, then, also raises the possibility that idiosyncratic attributes of the local community – for example, local elite tolerance, the expected police response to protests, the strength of local interest groups such as the ACLU, etc. – might play a significant role in facilitating or impeding protest. In any particular context, the specific causal processes at play may vary and may be quite complex. On the other hand, participation in protest may actually increase support for civil liberties and thus political tolerance as citizens come to no longer take basic political freedoms for granted (Peffley and Rohrschneider 2003; Duch and Gibson 1992). According to this second story, protest could *increase* tolerance.

To control for the possible endogeneity of macro-tolerance to previous levels of protest, we also include a lagged dependent variable in a second model (Model 2, Table 1): the incidence of protest in the MSA in the five years immediately prior to our 2007-2011 period of analysis. Although useful for bolstering causal conclusions, this is not an ideal specification inasmuch as

¹² We note that our results hold when (1) using the raw, unfiltered measure of protest incidence and (2) removing three MSAs that were influential outliers as defined by their Cook's Distances. These results are reported in the Online Appendix.

including a lagged dependent variable shifts the analysis to one of *change* in incidence of protest. And since macro-tolerance, being a form of political culture, changes slowly over time, we suspect that it is more weakly related to five-year *changes* in protest incidence than to contemporaneous *levels* of incidence.

Nevertheless, as Model 2 demonstrates, macro-tolerance remains positively and highly significantly associated with protest incidence, even when controlling for past levels of protest. These findings mitigate somewhat our concern that the association we observe between macro-tolerance and protest is merely a function of the influence of protest on tolerance.¹³

Finally, to further test the robustness of our hypothesized effect, we conduct a placebo test. In particular, we examine whether macro-tolerance is associated with higher levels of *non-contentious* political behavior as well as higher levels of protest. If so, this would undermine our hypothesis, because, as we argued earlier, intolerance inhibits protest because protest is public, threatening, and somewhat non-normative. Model 3 in Table 2 thus includes an OLS regression model of the proportion of the MSA population who voted in the 2012 presidential election. One can see that macro-tolerance is significantly associated with voter turnout, but *negatively* so. More tolerant MSAs, in other words, have slightly lower rates of voter turnout but higher rates of

¹³ Of the three major micro-level predictors of political intolerance – support for democratic institutions and processes, psychological insecurity, and threat perceptions (Gibson 2006) – only threat perceptions seem likely to be much affected by environmental conditions, at least in the short-term. As Gibson and Gouws (2003) have noted, this means that finding an exogenous instrument by which change in political tolerance can be induced is more difficult than one might imagine.

protest incidence. This analysis demonstrates that the dampening effects of intolerance on political behavior are limited to protest, as we hypothesized.

DISCUSSION AND CONCLUDING COMMENTS

This article has shown that metropolitan-area political tolerance is associated with heightened rates of protest. These findings have implications for debates around the exercise of civil liberties such as the freedoms of speech and assembly, which have recently leapt back onto the agenda in the U.S. and other long-standing democracies. In particular, concerns about everyday speech and behavior that might be interpreted as prejudiced, “blasphemous,” or politically incorrect or extreme, have made arguments *against* the freedoms of speech and assembly (or at least for limiting these freedoms) more salient and seemingly more legitimate. While our evidence does not suggest where exactly these limitations should lie, if at all, it does bolster the argument *for* putting up with a variety of unwelcomed speech if one’s goal is to legitimize and facilitate dissent and protest.

Indeed, one reason that one might put up with unpopular speech is the link our findings imply between tolerance and the vitality of democracy. While the level of protest expressed in a polity may not correspond in a linear fashion with the health of its democracy, protest is undoubtedly a necessary condition for democratic vitality. Citizens’ ability to freely gather, march, and express their opinions outside of routinized (and often controlled) political channels is what separates democracy from the various authoritarian facsimiles in circulation today. By dampening protest, cultures of intolerance undermine the health of democracy.

Our findings are consistent with the classic theory of the spiral of silence, which holds that intolerance toward those holding unsavory political opinions not only stifles these unpopular

voices, but also risks creating a broader culture of conformity in which all minority and dissenting views are quieted. Our research has not directly investigated that theory, however. The full story of how tolerance relates to dissent may begin with the political culture of the areas in which citizens reside, and may end up with macro-occurrences of protest, but it surely takes a detour through the micro-level perceptions and orientations of the citizens themselves. As such, although we have tried to carefully tailor our conclusions to our macro-level analysis, we certainly recognize that important micro-level linkages remain.

We also fully recognize the limits that result from thinking about a dynamic process in cross-sectional terms. Both political tolerance and protest levels change over time, probably in interconnected ways. Gibson's (1989) cross-sectional research gives a hint about this interaction, in his suggestion that tolerance makes campus-level protests possible, but that campus-level protests then stimulate the passage of repressive legislation. We have shown that a culture of political tolerance is associated with more protest activity, but there is undoubtedly more to be said were we to try to understand how the relationships evolve over time.

Finally, our understanding of protest is also under-theorized regarding institutional influences, particularly the role of law and courts. All metropolitan areas are governed, of course, by the nation's first amendment, but not all local judges embrace those freedoms with equal vigor. Even today, states differ in their promulgation of hate-speech laws, laws that are often used to limit the rights of political dissidents. In the same vein, protests do not take place in gigantic, metropolitan conglomerations; they take place in more finely-grained and localized communities. Throughout American history, it is these local communities that have passed countless ordinances making protest less or (usually) more difficult (e.g., Gibson and Bingham 1985). In this article, we have expanded the model of protest activity to include the effects of

political culture, but the model could certainly stand additional enlargements and elaborations.

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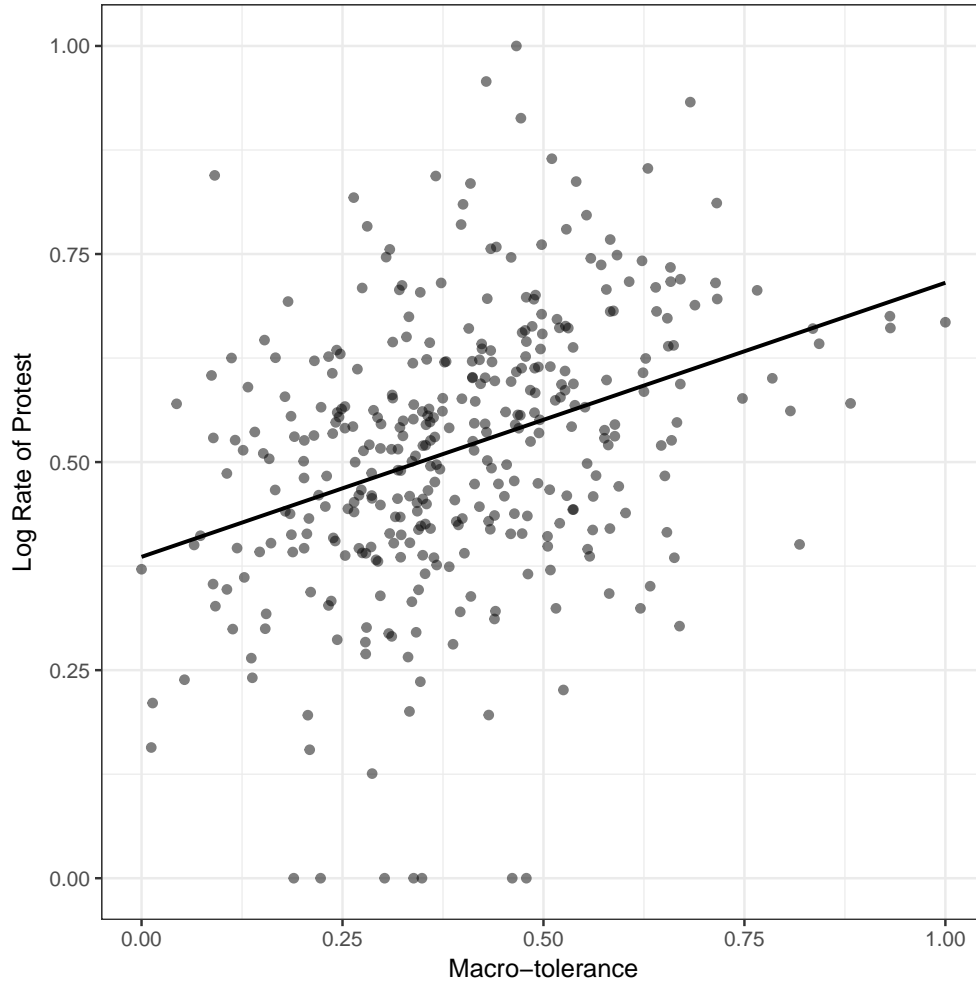
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Figure 1: The Relationship between Rate of Protest and Macrotolerance



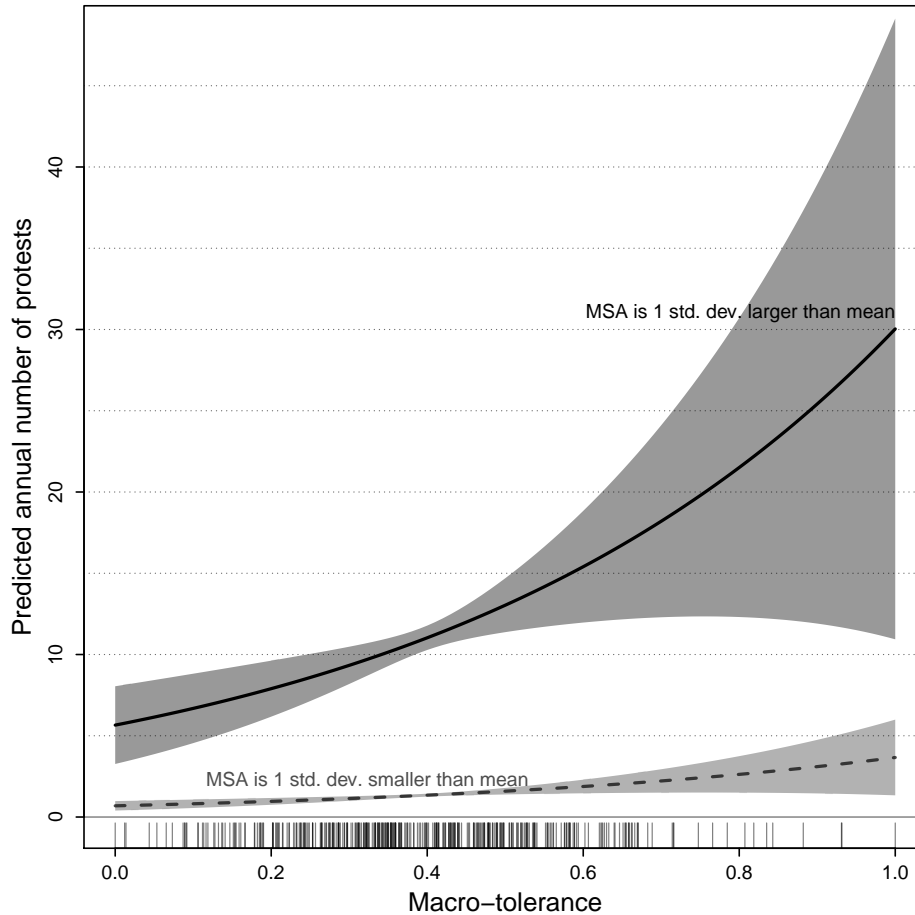
$N = 365$ MSAs. The solid line indicates a least squares regression fit. To calculate the log rate of protest we divide the annual MSA incidence of protest by MSA population in millions, and then find the natural logarithm. A small constant value (1) is added to all cases before taking the logarithm so as to avoid the undefined values that would result from taking the log of zero. We avoid having to take this step in our multivariate analyses of the determinants of protest because we model the protest counts directly. Both variables are standardized to range from 0 to 1.

Table 1: Models of Political Behavior in Metropolitan Statistical Areas

	(1) 2007-11 Protest Incidence	(2) 2007-11 Protest Incidence	(3) 2012 Voter Turnout
<i>Pull factors</i>			
Macro-tolerance (MRP estimate)	1.67 (.54)**	1.35 (.53)*	-.22 (.11)*
Unemployment rate 2007-2011	-.37 (.28)	-.35 (.27)	-.06 (.04)
Median household income	.59 (.43)	.28 (.42)	-.01 (.08)
Percentage students	-.61 (.32)	-.48 (.31)	-.17 (.05)***
Percentage under 18 years old	-.75 (.34)*	-.76 (.34)*	-.04 (.05)
Percentage with high school diploma	-.62 (.42)	-.38 (.41)	.81 (.08)***
Charitable organizations per person	.52 (.22)*	.55 (.21)**	.01 (.04)
Religious congregations per person	.41 (.28)	.33 (.27)	.05 (.05)
<i>Push factors</i>			
Percent Obama vote 2012	.54 (.27)*	.46 (.27)	.18 (.05)***
Gini coefficient	-.26 (.29)	-.41 (.28)	.09 (.04)*
Black-white segregation (dissimilarity index)	-.14 (.25)	-.28 (.24)	.05 (.05)
Ethnic fractionalization	.53 (.23)*	.43 (.23)	-.10 (.04)*
Percentage population increase 2000-2010	-.89 (.45)	-.81 (.44)	.21 (.08)**
Lagged incidence of protest 2002-2006		1.56 (.38)***	
Intercept	3.82 (.52)***	4.02 (.50)***	-.10 (.09)
<i>N</i>	365	365	363
Akaike Information Criterion	2822.83	2808.97	—
Adjusted R ²	—	—	.51
Dispersion	2.76	2.93	—
Residual standard error	—	—	.11

Note: * $p < .05$, ** $p < .01$, *** $p < .001$. The first two columns show the results of negative binomial regressions, with standard errors in parentheses. These models include the log of MSA population as an offset (i.e., with coefficient fixed to 1). The third column shows the results of an OLS regression with robust standard errors in parentheses. All explanatory variables are standardized to range from 0 to 1.

Figure 2: Predicted Effects of Macro-tolerance in Large and Small MSAs



This figure displays the annual predicted number of protests as macro-tolerance varies from its lowest observed level (0) to its highest (1). The solid line shows the predicted incidence for a hypothetical metropolitan statistical area that is one standard deviation larger than the average, and the dashed line shows the predicted incidence for a hypothetical metropolitan statistical area that is one standard deviation smaller than the average. The accompanying shaded regions indicate the respective 95% prediction intervals. The first negative binomial protest incidence model (Table 1) is used to generate these predicted effects. The “rug” of lines at the base of the graph show the observed distribution of metropolitan areas by their level of macro-tolerance.