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Article

Political fragmentation and widening disparities in African-American and white mortality, 1972–1988

Yonsu Kim [a,*], Tim A. Bruckner [b]

[a] Planning, Policy, and Design, University of California Irvine, 300 Social Ecology I, Irvine, CA 92697-7075, USA
[b] Program in Public Health, University of California Irvine, 635 E. Peltason Drive, Irvine, CA 92697-3957, USA

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ABSTRACT

Objective: During the 1970s and 1980s in the U.S., population movement, urban sprawl and urban governance reform led to a proliferation of local, autonomous jurisdictions. Prior literature examines how this creation of local governments, also referred to as political fragmentation, contributes to economic growth and social inequality. We examine the impact of political fragmentation on health equity by testing the hypothesis that the mortality disparity between whites and African-Americans varies positively with political fragmentation.
Methods: We retrieved mortality data from the multiple cause-of-death file and calculated total number of local governments per 1000 residents in a county to measure the degree of political fragmentation. We focused on 226 U.S. counties with population size greater than 200,000 and restricted the analysis to four distinct periods with overlapping government and mortality data (1972–73, 1977–78, 1982–83, and 1987–88). We applied generalized estimating equation methods that permit analysis of clustered data over time. Methods also controlled for the age structure of the population, reductions in mortality over time, and confounding by county-level sociodemographic variables.
Results: Adjusted coefficients of fragmentation are positive and statistically significant for both whites (coef: 2.60, SE: 0.60, $p < 0.001$) and African-Americans (coef: 5.31, SE: 1.56, $p < 0.001$). The two-fold larger positive coefficient for African-Americans than for whites indicates a greater racial disparity in mortality among more politically fragmented urban counties and/or time periods.
Conclusions: From 1972 to 1988, political fragmentation in large urban counties moves positively with the racial/ethnic gap in mortality between whites and African-Americans. We discuss intervening mechanisms through which political fragmentation may disproportionately affect mortality among African-Americans.

1. Introduction

After World War II, population growth and rapid urbanization in the U.S. led to dramatic political and geographic changes in urban areas. Scholars have explored how changes in the urban landscape affected economic growth as well as social and racial inequalities (Akai & Sakata, 2002; Bollens, 1986; Oates, 1985; Schneider, 1986; Stansel, 2008; Weiher, 1991). We build on this literature and focus on a relatively unexplored aspect: whether, and to what extent, increasing decentralization of political authority, in the form of governmental fragmentation, corresponds with rising health disparities between whites and African-Americans.

Political fragmentation refers to the process of redistributing functions, powers, or people away from a central authority by incorporating autonomous entities such as municipalities and special districts (Judd & Swanstrom, 2009). Decentralization of urban areas by population movement and urban sprawl in the U.S. led to a proliferation of local jurisdictions that established autonomous entities such as municipalities (Judd & Swanstrom, 2009; Morgan & Mareschal, 1999). Along with local governments, smaller jurisdictional boundaries resulting from fragmented governance allow residents to make locational decisions, considering the quality of schools, crime rates, racial composition and other public services (Weiher, 1991). Empirical studies report that political fragmentation accelerates spatial income and racial segregation (Bischoff, 2008; Miller, 1981; Morgan & Mareschal, 1999).

Hutson, Kaplan, Ranjit, and Mujahid (2012) used a cross-sectional analysis to examine the relation between fragmented governments and health disparities for large metropolitan statistical areas (MSA). The Authors examined data in the late 1990s and

* Corresponding author.
E-mail addresses: yonsuk@uci.edu (Y. Kim), tim.bruckner@uci.edu (T.A. Bruckner).

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report that the number of local governments varies positively with a disparity between white and African-American mortality. Their study provides a "proof of concept" of an association between political fragmentation and health disparities. However, Hutson and colleagues employed a fragmentation metric only for one time period in 1997. The pace of political fragmentation occurred rapidly in the 1970s and 1980s but significantly slowed in the 1990s. Such differences indicate that we cannot know the external validity of Hutson and colleagues' findings in understanding the health implications of urban landscapes that evolve over time. In addition, MSAs do not define legal or administrative municipal boundaries; rather, they define economically and socially integrated areas. Given that many local agencies are established primarily at the county level, Hutson and colleagues' choice of MSA as the unit of analysis may obscure meaningful county variation in political fragmentation that occurs within an MSA.

We build on the approach by Hutson and colleagues and assess whether results generalize to the dynamic period in the 1970s and 1980s. Specifically, we test whether fragmentation from 1972 to 1988 precedes an increase in African-American relative to white mortality rates. We improve upon earlier work in two ways. First, the structure of our analysis aims to exploit within- and across-county differences in fragmentation. Second, we control for confounding by secular improvements in mortality and the changing age structure of the U.S. population.

2. "White flight" and political fragmentation

In the 1960s and 1970s, metropolitan areas witnessed drastic growth in outer areas. The middle class moved from inner cities to suburban areas. Such urban sprawl and central city decline contributed to an outward movement of the economic base and employment opportunities (Jargowsky, 2002). This economic and demographic shift precipitated inequality between urban and suburban jurisdictions in access to public services, health care, affordable housing, education, infrastructure and job opportunities (Hutson et al., 2012). In addition, federal policies accelerated these inequalities by providing fewer opportunities for ethnic minorities. The Federal Housing Administration (FHA), the Veterans Administration (VA), and the Federal-Aid Highway Act helped affluent whites relocate to the suburbs and encouraged racial segregation (Cashin, 2010; Jackson, 1985; Judd & Swanstrom, 2009).

Coupled with urban sprawl, political fragmentation exerted a considerable impact on racial segregation. Newly incorporated government entities allowed middle class communities to segregate from the ethnic minorities or the less affluent, adopted zoning and planning restrictions, and provided tailored public goods and services for residents in that local jurisdiction (Bischoff, 2008; Hart, Kunitz, Sell & Mukamel, 1998). Researchers contend that this process of segregation promotes health disparities through several pathways, including through poverty concentration, insufficient housing, high unemployment rates, and low incomes (Acevedo-Garcia & Lochner, 2003; Hart et al., 1998; Jargowsky, 1997; La Veist, 1989; Massey & Denton 1993; Polednak, 1996; Wilson, 1996).

Our study focuses on urban governance that involves policies of local governments and provision of public goods and services. In accordance with federal and state authorities, many local agencies are established primarily at the county level. The delegation of authority to local governments, especially in the absence of federal or state laws, may lead to great variation of the structure and function of local health agencies across the country (Carter & Slack, 2010). Some counties vary in the extent of their public expenditures (e.g., public schools), taxes and social and fiscal policies. This large variation may affect the structures of economic and social opportunities, as well as the level of generosity of safety net programs. Such county differences may plausibly affect health disparities differentially, even if adjacent counties fall under the same larger metropolitan area.

The literature regarding political fragmentation examines whether fragmented governance enhances economic efficiency and contributes to economic growth. This hypothesis traces its origins to Tiebout's model and public choice theory. Tiebout (1956) argues that local governments provide public goods more efficiently in fragmented governance structures through the competition for residents who "vote with their feet". This argument coheres with public choice theory that stresses economic growth through the interjurisdictional competition in a decentralized organizational structure. Public choice theory implicitly criticizes a lack of competition in the consolidated government (Brennan & Buchanan, 1980; Brueckner, 2011; Kim & Jurey, 2013). Empirical studies assess the relation between fragmented governance and government expenditure, gross product growth, personal income growth, and employment growth. Findings, however, do not converge and remain controversial (Akai & Sakata, 2002; Eberts & Gronberg, 1988; Nelson & Foster, 1999; Raimondo, 1989; Schneider, 1986; Stansel, 2005).

Several studies examine the association between fragmented governance and social equity. Research reports that fragmented settings adversely impact income inequality and racial segregation presumably via increased poverty concentration in core urban areas and incorporation of the affluent community in the suburban areas (Bischoff, 2008; Hill, 1974; Swanstrom, 2001; Weiher, 1991). By contrast, other empirical studies that examine a panel of developing and developed countries report a positive association between fiscal decentralization and infant health (Mills, Vaughan, Smith, & Tabibzadeh, 1990; Robalino, Picazo, & Voetberg, 2001). Our study allows us to capture the variation of local government structures and its impact on health disparities in mortality. We also seek to replicate Hutson and colleagues' findings by utilizing data over four time periods spanning 15 years. We test the hypothesis that fragmented governance at the county level moves positively with widening health disparities in mortality between whites and African-Americans. Given the changes in the urban landscape in the 1970s and 1980s in the U.S., we focus our test on this time period.

3. Methods

3.1. Variables and data

We retrieved data on multiple causes of death from the National Vital Statistics System of the National Center for Health Statistics (NCHS). NCHS constructs the data on the basis of death certificates filed in each state. We obtained death counts at the county level. NCHS provides the entire death count data for the 1970s and 1980s across all counties in the U.S. These death counts include age and race information. Much literature documents the validity of age, race, and county identifiers on these death certificates and uses these files to describe racial disparities in mortality over this time period (Ezzati, Friedman, Kulkarni, & Murray, 2008; Levine et al., 2001; Meara & Culter, 2008; Murray et al., 2006; Satcher et al., 2005).

To measure the degree of political fragmentation, we employed total number of local governments per 1000 residents in a county. Prior literature uses this indicator to measure political fragmentation and more accurately captures the redistribution of political power in local governance than do alternative expenditure or revenue measures (Hawkins & Dye, 1970; Kim & Jurey, 2013). The U.S. Census Bureau conducts a census of local governments for all states for years ending in "2" and "7". This census limited our fragmentation data to the years 1972, 1977, 1982, and 1987. The U. S. Census Bureau classifies local governments into five different types of government. Out of those five types, the U.S. Census Bureau categorizes county, municipal, town, and township governments as general purpose local government. Special districts and school districts governments are referred to as special purpose government. For each census data year, we counted the number of local government entities for five types in each county, aggregated them to derive the total number of local government entities per county, and divided them by the county's total population to obtain a concentration of political fragmentation. A large value of political fragmentation in a county (i.e., more government entities per capita) indicates more fragmentation in urban governance as compared to other jurisdictions.

County governments routinely implement health policies and programs (DeFriese et al., 1981; Mays & Smith, 2009; Halverson et al., 1996). County governments operate their own health agency. Counties plan, establish and implement health policy, taking into consideration various factors that include health care access, social and economic factors, and physical environment in the county. In addition, county government involves management and implementation of revenue and expenditure on health, hospital, and other categories that may affect health status in the jurisdiction (education, police and fire protection, and public welfare). For these reasons, we view county as a meaningful unit for examining the relation between local governance and mortality disparities.

Given that political fragmentation data are available for 1972, 1977, 1982, and 1987, we restricted the mortality analysis to these years as well as the year immediately following the census (i.e., 1973, 1978, 1983, and 1988). We included the post-census year to increase stability of county-level mortality rates given that death counts for a narrow age range, in any particular county, may be low. We identified our outcome variable as the age-specific mortality rate in a county (i.e., deaths per population-at-risk). We estimated age-specific mortality for four distinct periods: 1972–73, 1977–1978, 1982–83, and 1987–88.

For each time period and race, we aggregated the number of deaths for the following age groups: less than 1, 1–4 years, 5–24 years, 25–44 years, 45–64 years, 65–74 years, and 75 years and higher. We used age-specific mortality to control for the strong confounding role of a county's age structure on the risk of mortality. We calculated age-specific mortality rates by dividing the death count for that age group by the population at risk in that age group, then multiplying it by 1000. We retrieved population-at-risk data from Surveillance, Epidemiology, and End Results program (SEER) for the years 1972–1987.

To minimize the possibility of artificially high outlying values in mortality rates due to underestimates of the population-at-risk, we used direct standardization in which we applied population age and race-specific estimates. First, we calculated the age distribution of each state's population for each year and each race by summing age and race-specific population of all counties in each state, and dividing it by total race-specific population of the state. Then, we multiplied that proportion by the total county population size provided by population-at risk data from SEER (see Appendix for details of direct standardization). We also eliminated observations if the estimates of age and race-specific population size for any particular county were smaller than 50. This process allowed us to detect outliers in mortality data in which, for instance, death counts are larger than or equal to race and age-specific population at risk. We found that such outliers resulted from underestimates of population at risk.

To control for other antecedents of disparities that may correlate with, but not be caused by, political fragmentation, we included the following county control variables: median household income, the proportion of population with some college degree, population density, and the proportion of population that identifies as African-American. These education and income metrics approximate a community's socioeconomic status (Adler & Newman, 2002; Braveman, Cubbin, Egerter, Williams, & Pamuk, 2010; Davey, Neaton, Wentworth, Stamier, & Stamier, 1998; Pappas, Queen, Hadden, & Fisher, 1993; Shavers, 2007). Income and education attainment data at the county level are available only for 1970, 1980, and 1990. We, therefore, used linear interpolation to impute median household income and the proportion of population with some college education. We focused on urban counties whose population size is larger than or equal to 200,000, because these areas in particular rapidly reformed local governments. This focus also aligns with prior empirical research on urban areas with a large population that promoted political fragmentation (Hutson et al., 2012; Pastor, Dreier, Grigsby, & Lopez-Garza, 2000). These urban counties are geographically distributed as follows: 62 counties (27.4%) in Northeast, 51 counties (22.6%) in Mid-West, 69 counties (30.5%) in South and 44 counties (19.5%) in West.

We conducted a variance component analysis (proc "varcomp" in SAS) that partitions across-county variation in political fragmentation from within-county variation over time. Results indicate that over 98% of the variance in political fragmentation occurs across counties. This circumstance also holds for other explanatory variables, including % African-American. We, therefore, reasoned that using county as a "fixed effect" variable may offer poor model fit and yield imprecise standard error estimates. As an alternative, we applied Generalized Estimating Equations (GEE) to analyze a relation between political fragmentation and age-specific mortality rates. Age-specific mortality rates in the same county, as well as mortality rates within a county but across years, are likely to be correlated due to shared county characteristics. This within-unit clustering violates the assumption of uncorrelated errors in linear regression (Hubbard, Ahern, Fleischerm, Van der Laan, Lippman, Jewell, Bruckner, 2010). GEE addresses this issue (Liang & Zeger, 1986) and is widely used to examine clustered data. The GEE approach provides a robust estimator of variance that takes into account the dependence, or clustering, of observations within higher-level (i.e., hierarchical) units such as counties. The GEE approach also has the advantage of not requiring additional distributional assumptions since the model estimation applies the observed data-generating distribution other than the joint distribution of observed data (Cui, 2007; Hubbard et al. 2010; Zeger & Liang, 1986).

The GEE method requires choosing an acceptable working correlation structure within units. We assessed four potential correlation structures that include independent, exchangeable, autoregressive, and unstructured (no specification). Then, we selected the correlation structure with the smallest QIC statistic, which corresponds to the lowest AIC in a likelihood-based method (Cui, 2007). The exchangeable covariance structure provides the smallest QIC statistics for both African-Americans and whites (results not shown). Our estimation, performed separately for African-Americans and whites, used the following equation across three different levels that include age group i in county j at time period t.

$$M_{ijt} = \beta_0 + \beta_1 G_{jt} + \mathbf{Age}_{ijt} + \beta_2 \mathrm{Inc}_{jt} + \beta_3 \mathrm{Edu}_{jt} + \beta_4 \mathrm{PAA}_{jt} + \beta_5 \mathrm{PD}_{jt} + \mathbf{Yr}_t + \epsilon_{ijt}$$

where M_{ijt} is the age-specific mortality rate (death count per 1000 population) in county j at time period t (e.g., 1972–73); G_{jt} is the indicator of political fragmentation (the number of local government per 1000 residents) in county j at time period t; $\mathbf{Age}_{jt}$ is the age indicator in county j at time period t ($i = 1, \ldots, 7$); Inc_{jt} is median household income in county j at time period t; Edu_{jt} is the proportion of its population with some college degree in county j at time period t; PAA_{jt} is the proportion of African-American population in county j at time period t; PD_{jt} is population density in county j at time period t; $\mathbf{Yr}_t$ is a vector of binary year indicator variables for all time periods t minus the referent time period 1972–3; β_i ($i=$ 1, …,5) are the coefficients for each of the five variables; β_0 is a constant; and ε_{ijt} is the error term that accounts for clustering of age-specific mortality rates within a county (exchangeable correlation).

β_1 is the coefficient of interest and measures the extent to which changes in the number of local governments per 1000 persons vary with age-specific mortality, controlling for county characteristics. The panel nature of the data, as well as the inclusion of indicator variables for year, allows us to detect a relation between political fragmentation and mortality that is not biased by time effects that might explain, for example, a secular decrease in mortality rates.

4. Results

The number of local governments increases over the 1970s and the 1980s. Each type of local government, except for school districts, shows an upward trend overall. Fig. 1 displays local government statistics over time for counties with a population size greater than or equal to 200,000. The total number of local governments in 1972 is 16,380 and increases to 19,401 in 1987 by 18.4%. The total population of these counties grows by 22.03% over this period (approximately from 111 million to 136 million). Special district governments show a remarkable increase by 36.5% over the 15 year period, whereas municipal governments show a milder increase by 13.5%. However, the rate of increase for school districts is relatively small, at 10.9%.

From 1972 to 1987, the mean number of local governments per county is 71.35. Table 1 lists the government types for the 20 most populous countries. Cook County (IL), Harris County (TX), King County (WA) and Allegheny County (OR) have relatively more local government entities than other counties with similar population size. Their indicators of political fragmentation are larger than 0.1. For example, a 0.1 fragmentation score indicates that a county has 0.1 government per 1000 residents.

Table 2 compares age-specific mortality rates between whites and African-Americans, aggregated over the four time periods. Consistent with the literature, African-American mortality rates exceed those of whites in all age groups except for 75 years or higher. The gap between whites and African-Americans is the largest among infants, followed by 65–74 years. The oldest group (75+) shows the well-documented mortality "crossover" where white mortality exceeds that of African-Americans (Martin & Soldo, 1987).

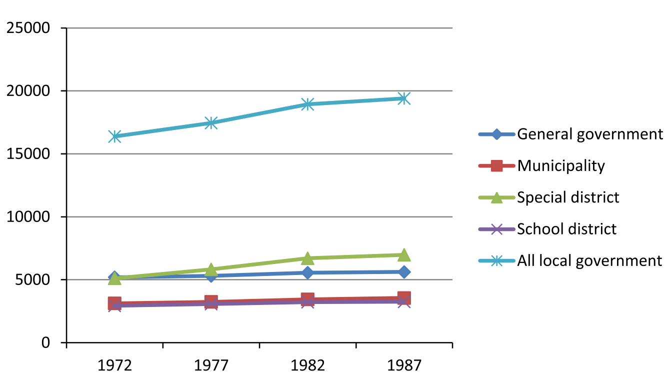

Fig. 1. Number of local governments, by type, in the U.S. from 1972 to 1987, among counties $\geq$200,000. *Note*: The counts of all local governments are the sum of county, municipality, town (or township), special district, school district governments.

Table 3 provides results of the two GEE analyses – one for each race/ethnicity. We analyzed 5768 age-specific mortality rates for whites, and 5628 rates for African-Americans across 226 counties in the U.S. (189 counties for 1972, 197 for 1977, 215 for 1982, and 223 for 1987). We excluded counties if population size did not reach 200,000 for a specific year. Thus, the number of observations varies with the time period. The coefficients of fragmentation are positive and statistically significant for both whites (coef: 2.60, Standard Error [SE]: 0.60, $p<0.001$) and African-Americans (coef: 5.31, SE: 1.56, $p<0.001$). The positive coefficient for African-Americans is more than twice the magnitude of that for whites.

To give the reader a sense of the magnitude of the results, we calculated the health disparities in mortality rate statistically attributable to political fragmentation. We applied the estimated coefficients to Cook County (IL), which shows relatively high fragmentation among the 20 largest counties. Applying the coefficients in Table 3 to the population size of Cook County shows a rise of 3.55 deaths per 100,000 for whites and 7.25 deaths per 100,000 for African-Americans with one standard deviation (SD) increase in political fragmentation. This result implies an excess of 3.70 African-American deaths per 100,000 persons statistically attributable to a one SD increase in political fragmentation. The magnitude of this excess death rate is similar to the current death rate due to cervical cancer in the U.S. (Murphy, Xu, & Kochanek, 2013).

Median household income has a negative but weak association with mortality only for African-Americans (coef: -0.094, $p=0.0001$). Consistent with previous research (Elo & Preston, 1996; Lleras-Muney, 2005; Pappas et al., 1993), the proportion of some college graduates varies inversely with mortality for both races (coef: -0.08, $p<0.0001$ for whites, coef: -0.19, $p<0.0001$ for African-Americans). The proportion of African-American population varies positively with mortality rates among both whites (coef: 4.68, $p<0.0001$) and African-Americans (coef: 2.88, $p=0.04$). The coefficient is higher for whites in higher percentage African-American areas. Given the debate about whether to include proportion African-American in ecological studies of mortality disparities, we assessed potential multicollinearity between this variable and political fragmentation (Ash & Robinson, 2009; Deaton & Lubotsky, 2003, 2009). We also re-ran the GEE model after removing proportion African-American to assess whether the political fragmentation coefficient changed. We find no evidence of multicollinearity; inference in the revised GEE model also remained similar to the original test (results available upon request).

We conducted four additional analyses to assess the robustness of our results. First, we re-estimated the GEE model with one equation that included both races and a race indicator variable (African-American$=1$; white$=0$) and its interaction term with political fragmentation, controlling for other covariates. Consistent with the initial test, we find that the African-American-fragmentation interaction term is positive and significant (coef: 3.28, $p=0.02$ for interaction). Second, we added the numeric year variable, year squared variable, and the interaction term of year and state indicator to determine whether state-specific time trends in mortality drove results. The control of time trends also addresses the possibility that mortality rates show an autocorrelated pattern not captured by use of binary indicator variables for year. The test shows similar results to the original model, indicating an increase in health disparities (coef: 1.53, SE: 0.47, $p=0.001$ for whites; coef: 4.89, SE: 1.47, $p=0.007$ for African-Americans). Third, consistent with Hutson and colleagues (Hutson et al., 2012), we re-ran the GEE analysis at the level of MSA to examine whether choice of geographic unit affected results. Whereas the magnitude of the findings differed from the original test, we found a larger association between political fragmentation and mortality among African-Americans relative to that of whites

Table 1
Number of local governments, by type, among the 20 most populous U.S. counties, 1987.

Counties	Population	# of general governments	# of special districts	# of school districts	# of All local Governments	Total governments per 1000 population	Population per sq. mile
LOS ANGELES, CA	8,553,846	85	142	95	322	0.04	2101.63
COOK, IL	5,172,402	151	213	152	516	0.10	5400.46
HARRIS, TX	2,745,989	29	432	24	485	0.18	1583.34
ORANGE, CA	2,281,483	27	73	32	132	0.06	2859.57
SAN DIEGO, CA	2,275,305	19	116	48	183	0.08	540.25
WAYNE, MI	2,161,100	45	10	36	91	0.04	3513.41
MARICOPA, AZ	1,991,399	22	59	56	137	0.07	218.19
MIAMI-DADE, FL	1,831,368	27	4	2	33	0.02	936.64
DALLAS, TX	1,816,646	27	25	16	68	0.04	2064.49
PHILADELPHIA CITY, PA	1,639,165	1	7	2	10	0.01	12,052.68
NEW YORK, NY	1,481,531	22	7	5	34	0.02	66,795.81
SANTA CLARA, CA	1,447,592	16	32	37	85	0.06	1119.57
CUYAHOGA, OH	1,438,906	61	8	32	101	0.07	3136.92
KING, WA	1,406,366	29	101	19	149	0.11	660.89
MIDDLESEX, MA	1,392,804	55	55	12	122	0.09	1695.34
ALLEGHENY, PA	1,355,064	129	126	43	298	0.22	1864.14
SUFFOLK, NY	1,310,590	23	10	7	40	0.03	1437.92
NASSAU, NY	1,304,805	31	27	17	75	0.06	451.59
ALAMEDA, CA	1,229,320	15	43	22	80	0.07	1669.46
SAN BERNARDINO, CA	1,192,192	18	92	37	147	0.12	59.42

Note: General governments include counties, municipalities, and town (township) governments. The number of all local governments is the sum of all local governments (general governments, special districts and school districts; Fragmentation indicators indicates the number of local government per 1000 resident; New York in this table refers to Now York County, which is one of five FIPS counties (Bronx, Kings, New York, Queens and Richmond) in New York City.

Table 2
Age-specific deaths per 1000 population for whites and African-Americans spanning 1972 to 1988[*].

Age group (yr)	White		African-Americans	
	Mean	Std dev.	Mean	Std dev.
0–1	11.85	7.26	24.33	13.73
1–4	0.63	0.83	1.06	2.23
5–24	0.82	0.51	0.85	1.33
25–44	1.65	0.81	3.42	3.24
45–64	9.12	2.41	15.41	8.29
65–74	29.17	6.29	39.26	13.48
75+	89.06	14.52	84.92	22.78

[*] We obtained age-specific mortality rates using death counts at the county level; the data years include 1972, 1973, 1977, 1978, 1982, 1983, 1987, and 1988.

(coef: 4.54, SE: 1.95, $p=0.02$ for African-Americans; coef: 3.01, SE: 0.62, $p<0.001$ for whites). Fourth, we partitioned the variance of political fragmentation to determine whether there exists substantial within-county variation over time to examine our hypothesis in a fixed effects framework. A fixed effect framework exploits only changes across time, within county, by holding county "fixed" in the regression equation. We discovered relatively little within-county variation over time (results available upon request), which precluded a fixed-effects estimation.

5. Conclusion

This study examines whether broad changes in urban governance correspond with a rise in mortality rate disparities between African-Americans and whites. Specifically, we test whether a county's political fragmentation from 1972 to 1987 varies positively with age-specific mortality rates more for African-Americans than for whites. We find that African-American mortality rises more than does white mortality as urban counties show greater political fragmentation. Results indicate that, in urban areas in the 1970s and 1980s, decentralization of political authority may have diverted county resources away from African-Americans more than from whites in ways that increased racial disparities.

Our findings bolster a prior study that reports a positive association between political fragmentation and health disparities (Hutson et al., 2012). We show that Hutson and colleagues' results have external validity to the 1970s and 1980s, a period in which important, and rapid, changes in urban restructuring occurred. In addition, our serial cross-section study design, which controls for clustering of place-level characteristics across time that jointly affect political fragmentation and mortality, yields more precise and policy-relevant estimates than do studies using one point in time. Our use of several time periods, moreover, allows us to detect secular patterns in the magnitude of the difference between white and African-American mortality. Furthermore, unlike earlier work, our analysis provides race-specific estimates between relation between fragmentation and mortality. This race-specific inquiry uncovered a positive relation between fragmentation and all-cause mortality in both whites and African-Americans – although of different magnitudes.

Strengths of our analysis include examination of political fragmentation over a 15 year period. Unlike earlier studies of aggregated annual data, our study minimizes confounding by strong secular patterns in mortality and minimizes bias due to unmeasured county factors. We further controlled for county characteristics such as median household income, proportion of some college degree, population density, and proportion of African-American population. Whereas we acknowledge the debate regarding the value of the "proportion African-American" variable in studies of mortality, we concur with the logic of Deaton and Lubotsky (2009) to include it as a control variable since it may capture structural differences across counties in terms of political or health opportunities. We note, moreover, that inference regarding our hypothesis remains robust to its omission.

Limitations include that the coverage of fragmentation data for 1972, 1977, 1982 and 1987 precluded analysis of mortality data in most intermediate years. Thus, our analysis could not capture more nuanced trends of mortality or any sudden transformation of government structure that occurred between these four time points. For this reason, we caution against using our results to infer associations for intermediate years. In addition, population counts between decennial Census years contain measurement error that the SEER population-based estimates may not correct.

Table 3
Generalized estimating equations of county-level, age-specific mortality in 226 urban counties (population $\geq$ 200,000) as a function of political fragmentation, income, education, proportion of African-American, and population density in the county. Whites and African-Americans analyzed separately.

Parameter	White				African-American			
	Estimate	95% C.I.		S.E.	Estimate	95% C.I.		S.E.
Intercept	87.50***	85.49	89.51	1.03	91.68***	87.63	95.72	2.06
Fragmentation	2.60***	1.42	3.77	0.60	5.31***	2.26	8.36	1.56
Year								
1972	2.80***	2.13	3.47	0.34	0.79	−0.48	2.05	0.65
1977	0.76**	0.30	1.21	0.23	−0.97*	−1.91	−0.03	0.48
1982	−0.23	−0.59	0.13	0.18	−2.08***	−3.12	−1.04	0.53
1987 (ref.)	–	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Age group (in years)								
0–1	−77.29***	−78.14	−76.44	0.43	−62.90***	−64.73	−61.06	0.94
1–4	−87.68***	−88.58	−86.79	0.45	−84.97***	−86.73	−83.22	0.90
5–24	−87.58***	−88.47	−86.69	0.45	−85.11***	−86.87	−83.35	0.90
25–44	−86.76***	−87.64	−85.87	0.45	−82.78***	−84.54	−81.01	0.90
45–64	−79.33***	−80.15	−78.50	0.42	−71.73***	−73.49	−69.96	0.90
65–74	−58.60***	−59.18	−58.01	0.30	−46.79***	−48.54	−45.05	0.89
75+ (ref.)	–	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Median household income (1000 USD)	0.009	−0.01	−0.03	0.01	−0.094***	−0.14	−0.05	0.02
% Some college	−0.08***	−0.11	−0.04	0.02	−0.19***	−0.28	−0.11	0.04
% African-Americans	4.68***	3.06	6.31	0.83	2.88*	−0.09	5.66	1.42
Population density	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
N	5768				5628			
# of counties	226				226			
QIC	5789.17				5624.37			

Note: The regression model estimates a change in mortality rates with an increase in political fragmentation across 226 counties over the four government census years; Year variables detect a change in mortality over time; QIC is information criteria in GEE model.; all tests are two sided.

*** $p < 0.001$.
** $p < 0.01$.
* $p < 0.5$.

Such error may yield imprecise estimates of our age-specific death rates.

We analyze all counties with a population size at or above 200,000. Whereas previous literature led us to focus on these urban areas, our results may not generalize to suburban or rural areas. Suburban and rural populations in the 1970s and 1980s show distinct patterns of growth, in terms of racial/ethnic and socioeconomic composition, relative to cities (Beale, 1977; Morrill, 1993; Probst, Moore, Glover, & Samuels, 2004). We know of no literature in political science or urban planning that develops predictions about the health consequences of political fragmentation in suburban areas. Given the documented "white flight" from cities to suburbs in the mid-20th century, we encourage future theoretical and empirical development regarding the relations among such migration, political fragmentation, and racial/ethnic health disparities.

We caution against using our results to make causal predictions about changes in one county's governance structure on the racial disparity in age-specific mortality. We discovered relatively little within-county variation in political fragmentation over time, which ruled out a county fixed effects estimation approach to minimize confounding by county-level factors. For this reason, we cannot rule out unmeasured confounding by factors at the county level that vary with, but are not caused by, political fragmentation and predict African-American more than white mortality.

In addition, we do not identify structural, economic, or social pathways that account for how political fragmentation moves with population health. Based on previous studies (Acevedo-Garcia & Lochner, 2003; Baron & Kenny, 1986; Jargowsky, 1997), we suspect that racial segregation plays a key role that accelerates poverty concentration or income inequality, and in turn increases race-based health disparities. However, the absence of validated variables on segregation or income inequality over this time period did not allow for testing these pathways. We recommend replication of our work using contemporary data with indicators of racial segregation, poverty, and income inequality. Current U.S. Census surveys on these variables include sub-county identifiers (e.g., city) which may allow refined examination of the partitioning of local governmental entities and resources across cities but within a county boundary. Increasing availability of expenditure data on public health services, moreover, may identify a fiscal pathway by which adjacent cities allocate county resources inequitably in ways that may increase race-based health disparities. These extensions to our work may elucidate whether, and to what extent, political fragmentation precedes an increase in race-based health disparities via segregation, poverty, and/or concentration of fiscal resources.

One somewhat counterintuitive finding involves the positive relation between political fragmentation and white mortality rates. Whereas we focus on the race/ethnic disparity in mortality rates, the observation that mortality rates for both whites and African-Americans vary positively with political fragmentation warrants more scrutiny. We speculate post hoc that migration of upwardly mobile whites to the suburbs in the 1970s and 1980s may have induced negative selection such that relatively less advantaged, and less healthy, whites may have remained in urban areas. Careful examination of these race-specific migration patterns would assist in providing support for, or rejection of, our speculative claims.

Our investigation includes all the counties whose population size is larger than or equal to 200,000 in the U.S. However, the process or the extent of government transformation in the 1970s and 1980s varied across U.S. regions. The Northeast region shows the highest political fragmentation among urban counties (0.22) whereas the South shows the lowest (0.07). We expect that future focus on a specific region or state would clarify the relation between fragmented governance and widening health disparities between whites and African-Americans.

Use of different types of government metrics may yield important findings. Bischoff (2008) reported that the proliferation of school districts increase multiracial segregation between districts. In the historical data of government census, school districts showed a different pattern of growth from other types of local governments. These alternative metrics merit additional consideration.

The multiple cause-of-death data from NCHS contains a single underlying cause of death based on death certificate. Future studies may use this information to detect cause-specific mortality that responds sensitively to the political and governance changes of urban areas. For instance, literature reports that changes of local governance affect the local economy (Foster, 1997; Grassmueck & Shields, 2010; Stansel, 2005; Swanstrom, 2001), which in turn perturbs rates of suicide, infant mortality, and motor vehicle fatality (Bruckner & Catalano, 2006, Bruckner, 2008; Catalano et al., 2011; Ruhm, 2000)

Empirical research concerned with the health sequelae of political structures and organizations have been controversial (Catalano, 2014; Rodriguez, Bound & Geronimus, 2013; Geronimus, 2000). Regarding the topic of fragmentation, advocates argue that local governments more likely meet the needs of residents and promote technical and allocative efficiency by competition under fragmented governance settings (Akai & Sakata, 2002; Brennan & Buchanan, 1980; Grassmueck & Shields, 2010). By contrast, opponents stress that economic efficiency arises from consolidated government entities and expansion of a public sector (Foster, 1997; Nelson & Foster, 1999). Similarly, studies focusing on whether political fragmentation causes racial segregation, economic inequality and reduced opportunity do not converge (Bischoff, 2008; Hill, 1974; Lewis & Hamilton, 2011; Morgan & Mareschal, 1999). Our research, however, on political fragmentation and health disparities supports and extends earlier research (Hutson et al., 2012). Our results imply that heightened political fragmentation in large urban areas in the 1970s and 1980s in the U.S. corresponds with widening health disparities between whites and African-Americans. We anticipate that subsequent research in this area will assist with community health planning and policy efforts to balance economic efficiency with social equity.

Appendix. Direct standardization of mortality rates

The National Vital Statistics System routinely computes age-specific death rates using the direct method of standardization (Anderson & Rosenberg, 1998). Consistent with this method, we used the direct method to estimate county- and age-specific death rates, which we cannot directly observe given uncertainty in the county- and age-specific counts of the population at risk of death in any year.

Let

D_{ij} = the number of deaths in age interval i in county j.
P_{ij} = the population in age interval i in county j.
P_{ik} = the population in age interval i in state k.

We aim to estimate the county- and age-specific death rate (R_{ij})

$$R_{ij} = \frac{D_{ij}}{P_{ij}} \tag{1}$$

Which we express as deaths per 1000 population. Given uncertainty in the estimates of P_{ij} due to small counts, we approximated P_{ij} in two steps. First, we estimated the age-specific population counts for the entire state by summing the age-specific counts from all counties within that state

$$P_{ik} = \sum_{1}^{n} P_{ij} \tag{2}$$

where n represents the number of counties j in that state. Second, we divided the state's age-specific counts by the state's overall population count to yield standard weights of the age distribution of the population. The standard weights w_{sik} are then given by

$$w_{sik} = \frac{P_{ik}}{\sum_i P_{ik}} \tag{3}$$

where $0 < w_{sik} < 1$ and the w_{sik} sum to 1. We calculate the county- and age- specific death rate (CASDR) as follows:

$$CASDR_{ij} = \frac{D_{ij}}{w_{sik}{}^{*}P_j} \tag{4}$$

The SEER database provides the county's estimated population size P_j in that year. We applied all the steps above, separately for each of the age, race/ethnicity, state, and year combinations for which we had data (i.e., 0–1 year, African-American, 1972, Georgia, etc.).

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