LS
7/26/06
11:30 AM

# Tipping and the Dynamics of Segregation in Neighborhoods and Schools

July 2006


| David Card | Alexandre Mas | Jesse Rothstein |
|---|---|---|
| UC Berkeley | UC Berkeley | Princeton University |


**ABSTRACT**


*Neighborhoods and schools in most U.S. cities are highly segregated by race and ethnicity. In a classic paper Schelling (1971) showed that segregation can arise from preferences of whites that generate a tipping point: once the minority share in a neighborhood exceeds a certain threshold, all the whites leave. We use regression discontinuity methods and Census tract data from the past four decades to study the mobility responses of whites to differences in minority shares. Nearly all cities exhibit tipping-like responses, with a range of tipping points centered around a 13% minority share. Similar patterns emerge from an analysis of school-level data over the 1990s. A variety of specifications rule out the possibility that the discontinuity in the initial minority share is driven by income stratification or other factors, pointing strongly to an important role for white preferences over neighbors' race and ethnicity in the dynamic process of racial segregation. Further evidence that tipping derives, at least in part, from whites' desire to avoid minorities comes from an analysis of survey data: White racial attitudes are robustly correlated with the location of the local tipping point.*



*We are grateful to David Walton for outstanding research assistance, and to Ted Miguel and participants in the Berkeley labor lunch for comments and suggestions. This research was funded in part by the Center for Labor Economics at UC Berkeley and by the Industrial Relations Section at Princeton University.

What explains the persistence of racial and ethnic segregation in the United States? In a classic paper Schelling (1971) proposed that segregation arises from strategic interactions among white families, each moderately tolerant in isolation.[1] Once the fraction of minorities in a neighborhood exceeds a certain tipping point, a cascade of white flight leads to complete racial isolation.[2]

Despite the appeal of Schelling's model, empirical evidence on the existence of tipping behavior is limited. A long literature (starting with Duncan and Duncan, 1957 and Taeuber and Taeuber, 1965) has documented the process of racial change in neighborhoods, though few studies have attempted to identify specific tipping points. Two comprehensive recent analyses have failed to find evidence of tipping at either the neighborhood or school level. Easterly (2004) argues that neighborhood racial composition exhibits mean regression rather than the divergent pattern predicted by a tipping model. Clotfelter (2001) concludes that the effect of increased exposure to minority schoolmates on the growth of white enrollment is essentially linear.[3]

In this paper we use tract level data from the 1970-2000 Censuses and school level data from the Common Core of Data to re-examine white mobility patterns and test for tipping behavior across major U.S. cities. To illustrate our approach, Figure 1 shows the estimated

---

[1] In the decades before Schelling's paper, sociologists had documented the evolution of neighborhood racial composition in response to the arrival of black families. Grodzins (1958) defined the "tip point" as the percent of black residents that "...exceeds the limits of the neighborhood's tolerance for inter-racial living."

[2] Granovetter (1978) proposed a tipping-style model to explain technology adoption voting, migration, and social conformity. See Heal and Kunreuther (2006) for a game-theoretic model of tipping that emphasizes strategic interdependence.

[3] Earlier studies of school district data, including studies of school desegregation programs, are reviewed in Section II, below.

relationship between the minority share of census tracts in Chicago in 1970 and the change in the tract-level white population (as a share of the initial population) from 1970 to 1980. Note that the fitted change in the white share from a local linear regression drops off sharply at approximately a 5 percent minority share, suggesting a "tipping point." Tracts with initial minority shares around 3% saw white population growth, while those with just slightly higher minority shares experienced white outflow rates that approached 20% of the 1970 population. The figure also shows the frequency distribution of minority shares across tracts in 1970. This exhibits a clear shoulder at about the same 5% point, as would be expected if neighborhoods with higher minority shares were dynamically unstable. Inspection of similar plots for other major cities confirms that the pattern in Chicago is the norm. In most cities there is a discernable point beyond which the white population growth rate falls substantially.

To formalize this analysis, we use flexible regression specifications to estimate metropolitan- and decade-specific relationships between the minority share in a neighborhood or school and the subsequent change in the white population share. We identify the location of a potential tipping point as the level of the initial minority share at which the predicted rate of change of the white share equals the city-wide average. For Chicago in 1970, this point is 5.6%, where the fitted regression line in Figure 1 intersects a horizontal line that represents the average tract's white population growth rate (again expressed as a share of the 1970 population) between 1970 and 1980. We are able to identify such a point for almost every city in each decade. The potential tipping points range from 1 to 41%, with a mean of about 12% in the 1970s, 13% in the 1980s, and 14% in the 1990s, and are highly correlated (around 0.6) over time.

Having identified city-specific potential tipping points we turn to an analysis of the dynamic behavior of white shares around these points. We pool the data from different cities by deviating the initial minority share for a given tract or school from the corresponding city-specific potential tipping point. Our analysis provides strong evidence of tipping behavior, in the form of large discontinuities in the white population growth rate at the potential tipping points. In the 1970s, a Census tract with a minority share just beyond the potential tipping point lost whites (relative to the metropolitan average) accounting for 20 percent of its 1970 population over the next decade, while a tract just below the tipping point saw relative growth in its white population equal to about 15 percent of its initial population. The discontinuity was smaller in the 1980s and 1990s, but still important. We also examine school-level tipping during the 1990s, and again find substantial discontinuities, comparable to those seen in neighborhoods, in elementary schools' white enrollment growth.

While Schelling's (1971) model treats racial composition as the source of externalities in location choices, an alternative is that people care about their neighbors' incomes (Schelling, 1978; Bond and Coulson, 1989). Since minorities have lower average incomes than whites, such preferences will lead to some degree of segregation. Moreover, even in the absence of concerns over neighbors' characteristics, standard models of spatial equilibrium imply that households will sort into neighborhoods with similar incomes, again leading to some segregation. To distinguish race-based tipping from these alternative explanations, we augment our models with tract-level controls for mean family income, average unemployment, and the characteristics of the housing stock. These models continue to show discontinuities in the rate of change in the white share at the tipping point, suggesting that preferences over the race and ethnicity of neighbors per se influence white families' location choices.

In addition to testing for tipping behavior in white population shares, we look at several other neighborhood outcomes, including average prices for owner-occupied homes and the number of housing units in a tract. Values of owner-occupied homes show a modest discontinuity at the tipping point that predates the population change. That is, tracts that tipped during the 1990s had lower values in 1980 than those that did not. There is also a discontinuity in the rate of new housing construction. Both effects are consistent with a modified Schelling model that incorporates heterogeneity in housing quality within and between neighborhoods.

Finally, we relate the location of the neighborhood tipping point in each city to the racial attitudes of white residents in the city. Specifically, we construct an index of racial attitudes based on responses to four questions in the General Social Survey regarding inter-racial marriage, school busing, and housing segregation. Controlling for other city characteristics (including racial composition, mean incomes of different racial/ethnic groups, and regional controls), we find a robust and quantitatively important link between racial attitudes and the location of the average tipping point. This lends some additional support to the view that the tipping behavior identified in our main analysis is driven by white preferences over minority contact.

## II. Previous Literature on the Dynamics of Racial Composition

### *a. Neighborhood Change*

In the 1920s the "Chicago school" sociologists (e.g., Park, Burgess, and McKenzie, 1925) developed a descriptive model of the neighborhood transition process set off by the arrival of new immigrant groups. This model formed the conceptual framework for Duncan and

Duncan's (1957) landmark study of racial transition in Chicago neighborhoods. The Duncans argued that after the initial arrival of black residents, neighborhoods passed through a series of stages that culminated with a 100% black population. They implicitly assumed that white flight was inevitable once more than a handful of blacks moved into a neighborhood.[4] Grodzins (1957, 1958) called the threshold level of black penetration the "tip point." Taeuber and Taeuber (1965) studied 10 other cities and concluded that the rate of change of the white population in a given neighborhood was also affected by city-wide trends in the relative population growth of whites and blacks, a feature that we use in our classification of tipping points, below. More recent sociological studies have confirmed that the effect of the minority share on white mobility varies across cities (e.g., Lee and Wood, 1991), and in some cases with the identity of the minority group (Massey and Denton, 1991).

Economists have focused on theoretical aspects of neighborhood transition. Building on the findings of Duncan and Duncan (1957) and Taeuber and Taeuber (1965), Schelling (1971) developed a simple choice model emphasizing the interdependence of whites' location decisions.[5] Miyao (1979) and Kanemoto (1980) derived tipping-like behavior in models with explicit land markets. In contrast to neighborhood externality models, standard "filtering" models (Muth, 1973; Brueckner, 1977) assume that neighborhood transitions arise from the depreciation of the housing stock. To the extent that minorities have lower incomes than whites,

---

[4] Rapkin and Grigsby (1960) present a detailed analysis of neighborhood change in Philadelphia, focusing on real estate transactions in one year (1955) in racially mixed areas. Their results are largely consistent with tipping at very low concentrations of black residents. Schwirian (1983) provides a somewhat dated review of the sociological literature on neighborhood change.

[5] An earlier model of the boundary externality between rich and poor neighborhoods was developed by Baily (1959). Schelling (1969) presents a model of a linear neighborhood that is closer to Baily's.

filtering implies a smooth (endogenous) rise in the fraction of minorities in a neighborhood, rather than a causal effect of racial composition. Bond and Coulson (1989) develop a model in which filtering combines with neighborhood externalities to generate tipping. We present a variant of this model in the next section.

In a recent study Easterly (2004) tests for nonlinear effects of initial minority share on white mobility, using the same national Census tract data that we use here. Easterly relates the change in the white share of tract residents between Censuses (or over multiple decades) to a fourth order polynomial in the initial white share. He finds strong evidence of "mean regressive" behavior, particularly for tracts with initial white shares under 20 percent, but little or no evidence of discontinuous tipping. We believe there are several features of Easterly's (2004) analysis that mask tipping behavior, including the assumption of similar dynamic behavior in all cities (or in large groups of cities) and the use of a smooth polynomial (rather than a potentially discontinuous model). Our analysis of the same data reaches a very different conclusion.

A few studies have examined how neighborhood composition affects individual mobility. South and Crowder (1998) and Quillian (2002) analyze the effects of black neighborhood shares on the outflow rates of white and black families, using longitudinal data from the PSID. Both studies find that white exit rates are higher from neighborhoods with more than 10% black residents, though neither attempts to identify a specific tipping point. Mare and Bruch (2003) study residential choices for a sample of Los Angeles residents. They find that whites' probabilities of moving to a new neighborhood are positively affected by the fraction of black residents in their original tracts, with some evidence of non-linearity. Unfortunately the sample sizes in these studies are modest and given the uneven distribution of minority shares they have limited power to detect nonlinear or discontinuous effects.

*b. Direct Survey Evidence on Preferences over Neighborhoods*

Another related group of studies asks people directly about their willingness to live in neighborhoods of different compositions. Farley et al. (1978) presented stylized neighborhood maps to Detroit residents and asked whether they would move out of homes with different fractions of black neighbors.[6] Whites apparently disliked black neighbors: 7% stated they would try to move out if 1 out of 15 nearby homes were occupied by a black family, 24% if the black share was 3/15, and 41% if the black share was 5/15. Black respondents showed some preference for neighborhoods where the black share was around 50%.[7] In the context of Schelling's (1971) model, the distribution of white preferences reported by Farley et al. (1978) imply that any non-zero minority presence is sufficient to tip a neighborhood. A follow-up study in 1992 (Farley et al., 1993) finds similar but less extreme race-sensitivity among whites.

A number of studies have tried to identify the reasons for the unwillingness of whites to move into (or remain in) neighborhoods with a high black share. Bobo and Zubrinsky (1996), for example, attempt to separate general attitudes against people in other racial/ethnic "out-groups," or in lower ranked social groups, from prejudicial attitudes against black neighbors. They conclude that prejudice is the underlying factor.

[6] Pettigrew (1973) reviews earlier studies of attitudes regarding race and housing.

[7] Some analysts have argued that current patterns of residential segregation are largely driven by black preferences against living in areas with less than a majority of black neighbors (e.g., Patterson, 1997; Thernstrom and Thernstrom, 1997). This hypothesis does not specify how the fraction of minorites affects *white* mobility, which is the object of interest here and in most of the tipping literature.

*c. School Desegregation and White Flight*

A third relevant literature studies the reactions of white students to increases in the fraction of minority schoolmates caused by desegregation programs. An appealing feature of these studies is that the changes in minority enrollment following implementation of a desegregation plan are potentially exogenous, unlike the residential flows that determine neighborhood minority shares. Coleman, Kelley, and Moore (1975) study changes in white enrollment shares at 67 larger districts that underwent desegregation in the late 1960s, and find a significant negative effect of increased exposure to black schoolmates.[8] Welch and Light (1987) likewise find significant declines in white enrollment following implementation of court-ordered desegregation plans. Reber (2005) extends the Welch and Light data and finds persistent losses in white enrollment following desegregation, with larger losses in cities with more suburban school districts.

A recent study by Clotfelter (2001) examines district-level data for larger metropolitan areas of the US over the 1987-96 period. Clotfelter relates the growth rate of white enrollment in a district to the exposure of whites to minorities in the base year and other controls (including a measure of the proximity to nearby low-exposure districts). Consistent with earlier studies, he finds a negative effect of exposure on the growth in the white share. Clotfelter uses flexible (polynomial) regression models to look for evidence "tipping like" behavior, but concludes that the response of the white share to increasing exposure is essentially linear over the relevant range. In the analysis below we extend Clotfelter's analysis by studying data at the school level, and by allowing city-specific functional forms for the relation between the change in white

[8] Clotfelter (1979) and Farley, Richards, and Wurdock (1980) reach similar conclusions.

enrollment and initial minority share. As in our neighborhood analyses, we find that this flexibility is important to the identification of strong tipping patterns in nearly every city.

## III. Theoretical Models of Tipping Behavior

### *a. A Basic Model*

It is useful to begin by restating Schelling's simplified (1971) model, which ignores both housing quality and prices. Schelling assumed that whites are indifferent between neighborhoods with minority shares less than $m^*$ but are unwilling to tolerate a minority share above $m^*$, while minorities do not care about neighborhood composition. Ignoring for the moment any heterogeneity in $m^*$, an equilibrium then has the property that families are distributed across two types of neighborhoods: integrated neighborhoods (with $m<m^*$) and all-minority neighborhoods. This equilibrium exhibits discontinuous "tipping" behavior. A neighborhood with a minority share $m<m^*$ is stable, but if for any reason an integrated neighborhood receives an injection of minorities that raises the minority share above $m^*$, all the whites will leave.

The assumption that all whites have the same threshold is highly restrictive. Following Schelling (1971), assume instead that $m^*$ is a random variable, distributed across the white population with distribution function F[ ]. The fraction of the white population willing to live in a neighborhood with minority share m (and white share 1-m) is 1-F[m]. Thus, if a neighborhood that initially contains a random sample of whites sees its minority share rise to the point that 1-F[m] is less than 1-m, there will no longer be enough whites willing to live in that neighborhood to occupy 1-m of the houses, and m will rise. Thus, for stability we need:

(2) $1 - F[m] \geq 1\text{-}m$ or $m \leq F[m]$.

Assuming a conventional shape for F, this leads to the situation shown in Figure 2. Any minority share less than the point $m^{**}$ – where the distribution function F cuts the 45 degree line from below– is a potential equilibrium. Any minority share above $m^{**}$ is unstable. A neighborhood that found itself with $m>m^{**}$ would experience cascading white flight, with the least-tolerant whites leaving first and thereby lowering the white share to the point that even the more tolerant whites leave. $m^{**}$ is a tipping point.[9]

Although Schelling's model ignores housing prices, it is straightforward to include land markets so long as white preferences—and their implied bid-rent functions—are discontinuous in m.[10] It is perhaps not surprising, however, that such preferences produce discontinuous responses. Moreover, it is unclear how (or why) the minority share would ever cross the tipping point $m^{**}$. To construct a more general model that exhibits tipping with continuous preferences, we present a modified version of the model developed by Bond and Coulson (1989).

*b. A Model with Quality Heterogeneity*

Consider a city with many neighborhoods and two race groups, whites (denoted by W) and minorities (denoted by M). Assume that a family in race group $r \in \{M,W\}$ has preferences represented by the function $U^r(X, q, m)$, where X is a numeraire consumption good, q is housing

---

[9] Though in our example $m^{**} = E[m^*]$, this is not generically true. In the simplified case of identical preferences among whites, F is just a step function at the common minority threshold $m^*$, leading to a tipping point at $m^*$.

[10] Alternatively, one could assume that whites have a bid-rent function that is smoothly downward-sloping in m and that cuts the bid-rent function for minorities from above at some minority share $m^*$. This would yield tipping, but in both directions: A neighborhood that had $m<m^*$ would tip toward m=0 just as one with $m>m^*$ would tip toward m=1. This does not seem consistent with evidence that neighborhoods with relatively low minority shares are stable.

quality (with $\partial U/\partial q>0$), and m is the minority share in the neighborhood (with $\partial U/\partial m<0$). To simplify the analysis, assume that all whites have the same income $Y^W$ and all minorities have the same income $Y^M<Y^W$.[11]

With free mobility, families of race group r will achieve the same utility level $U^r$ in any neighborhood that they choose to live in. This yields a pair of equilibrium bid-rent functions $b^r(q, m, U^r, Y^r)$, implicitly defined by the equality $U^r = U^r(Y^r \sim b^r(q, m, U^r, Y^r), q, m)$. We make two assumptions about these functions:

(3a) $\partial b^W(q, m, U^W, Y^W)/ \partial m < \partial b^M(q, m, U^M, Y^M)/ \partial m$ .

(3b) $\partial b^W(q, m, U^W, Y^W)/ \partial q > \partial b^M(q, m, U^M, Y^M)/ \partial q$.

These inequalities imply that white families have a higher willingness-to-pay than minorities for marginal increases in q or decreases in m. If whites and minorities have identical preferences but whites have higher incomes, (3a) and (3b) are ensured by a standard "single crossing" assumption on preferences.

Consider now the demand for houses in a neighborhood with minority share m. Define the quality threshold a(m) such that whites and minorities would both bid an equal amount for a house of quality q=a(m):

(4) $b^W(a(m), m, U^W, Y^W) = b^M(a(m), m, U^M, Y^M)$.

By assumption (3b), minority families outbid white families for any lower-quality houses (those with q<a(m)), while white families outbid minorities for higher-quality houses (q>a(m)). (3a) and (3b) ensure that a(m) is upward-sloping and invertible. Let H denote the distribution

[11] The results below also hold in a more complex model with overlapping income distributions where the white share is increasing in income. The assumption that minority families' utility is declining in m is also unnecessary but convenient.

function of quality q in a given neighborhood. If the neighborhood has minority share m, a fraction H(a(m)) of its houses will attract higher bids from minorities than from whites. Thus, for an integrated neighborhood's housing market to be in equilibrium, it must be that H(a(m))=m. If H(a(m))>m for all values of m, the only equilibrium has m=1; while if H(a(m))<m for all m, then m must equal 0 in equilibrium.

It is convenient to recast this in terms of q and $m=a^{-1}(q)$. An integrated equilibrium requires that there exists some q with $H(q)=a^{-1}(q)$; the neighborhood will then have minority share $m=a^{-1}(q)$. Figure 3 plots illustrative functions $a^{-1}(q)$ and H(q), assuming for simplicity that q is uniformly distributed between $q_0$ and $q_1$.[12] There are three equilibria, labeled A, B, and C in the figure. Point A represents a stable interior equilibrium: if there was a small increase in the fraction of minority residents starting from this point, then locally $H[q] < a^{-1}(q)$ and so demand from whites would push m down. Point B represents an unstable interior equilibrium: a small increase in the fraction of minority residents would lead to a situation where $H[q] > a^{-1}(q)$, so minorities would buy/rent more houses, eventually pushing the neighborhood to the all-minority corner equilibrium at C.

Assuming that quality declines as houses age, this simple model leads to an interesting set of predictions about neighborhood succession. Consider the sequence of H functions in Figure 4. $H^1$ represents an initial situation in which housing quality in a neighborhood is so high that only whites live there. As the housing stock ages, the H[q] function shifts left to $H^2$. Here, there is a stable integrated equilibrium (point B) in which minorities outbid whites for some of

---

[12] Thus $H[q] = 0$ for $q<q_0$, $H[q] = (q \sim q_0)/(q_1 \sim q_0)$ for $q_0 \le q \le q_1$, and $H[q] = 1$ for $q>q_1$. Note that the $a^{-1}(q)$ function need not be concave, as in Figure 3, though Bond and Coulson (1989) present an example with a Cobb-Douglas utility function in which it is.

the lowest-q homes. As houses age further and H shifts further left, the minority share in the stable equilibrium shifts upward, as does the quality of the threshold house. Eventually, we arrive at $H^3$, which is just tangent to the $a^{-1}(q)$ function at C. This is a tipping point: with any further aging of the housing stock, the H[q] and $a^{-1}[q]$ functions will no longer intersect, and the only equilibrium will be at point D, with m=1.

Bond and Coulson (1989) explore the implications of this model for hedonic-style efforts to estimate the utility function and for analyses of "filtering." For our purposes, it is useful to draw out just a few implications. First, if housing quality steadily declines, then a stable equilibrium's minority share will gradually increase over time until the tipping point is reached, at which point the neighborhood will jump to m=1. Second, average rents will gradually decline until the tipping point is reached, then jump downward. This is because prices are constant on the a(m) locus and increasing in the difference q-a(m), and aging of the housing stock only gradually lowers the q-a(m) distribution until m and a(m) jump upward when the neighborhood tips. Third, tipping is fully predictable. The downward jump in rents that accompanies tipping will thus be capitalized into sale prices well in advance. The price of a constant-q house will gradually trend downward (as the future stream of rents comes to include more post-tipping periods) until tipping occurs, then stabilize.

A final important implication of the model relates to new home construction. Assuming that newly built homes are at the top of the quality distribution, new construction pulls the H function rightward, offsetting the leftward drift due to depreciation and potentially ensuring that the neighborhood never tips. Construction in all-white or integrated neighborhoods is more profitable than in all-minority neighborhoods because a(m) is lower in a low minority area:

hence the gap q-a(m), which determines the price for which a new house can be rented, is higher. Thus, the model predicts that neighborhoods that are tipping or that are expected to tip in the future will have lower rates of new home construction than do neighborhoods that are not on a path toward tipping.[13]

## IV. Testing for Tipping Behavior

### *a. Regression Discontinuity Framework*

In this section we develop our empirical approach to measuring tipping behavior in neighborhoods and schools. Following most of the existing literature on neighborhood transition, we take as our dependent variable the change in the white (non-Hispanic) share of a neighborhood or school, and as the key explanatory variable the initial fraction of racial or ethnic minorities (i.e., Hispanics and nonwhites) in the neighborhood or school. Let $P_t$ represent the total population in a neighborhood (or school) in period t, let $W_t$ represent the number of white residents (or students) in period t, and define $m_t = (P_t - W_t)/P_t$ as the minority share in period t. The stylized tipping model of Schelling (1971) suggests that the expected change in the white share from some baseline period t-j to the end period t, conditional on the minority share $m_{t-j}$ in the baseline period, evolves as:

$$(5) \qquad E[\ \Delta(W_t/P_t) \mid m_{t-j}] \begin{cases} = a & m_{t-j} \leq m^*, \\ = b & m_{t-j} > m^*, \end{cases}$$

[13] The potential for new construction to prevent tipping can create self-fulfilling prophecies: A neighborhood that is expected not to tip might attract sufficiently rapid construction and renovation to prevent tipping, while expectations of tipping could lead to lower construction rates that bring this about. Note also that "un-tipping" a neighborhood most likely requires its wholesale redevelopment: As long as there are some homes in the neighborhood with q<a(1), m=1 is a stable equilibrium, so un-tipping without demolishing all low-quality homes can only be accomplished if the neighborhood can jump from one stable equilibrium to another.

where $m^*$ is the tipping point, and b<a. This has the form of a "regression discontinuity" model for the growth rate of the white share, with a discontinuity at $m^*$. Inspection of (5), however, suggests two problems. First, $m^*$ is unknown. Thus, conventional regression discontinuity methods cannot be applied directly. Second, if b<0 equation (5) cannot literally hold for high-minority neighborhoods (i.e., those with $m_{t-j} > 1-b$) since the white share cannot fall below 0.

Our approach to the second problem is to specify a and b as "smooth" functions of the initial minority share, with a potential discontinuity at a city-specific tipping point.[14] Thus, conditional on $m^*$, we assume

(6) $$E[\ \Delta(W_t/P_t) \mid m_{t-j}] = a(m_{t-j}) * 1[m_{t-j} \leq m^*] + b(m_{t-j}) * 1[m_{t-j} > m^*],$$

where 1[ ] is an indicator function, and the functions a(m) and b(m) are low order polynomials. The extent of tipping is the gap $a(m^*) - b(m^*)$.

The first problem is the more serious. There is as yet no consensus method for estimating regression discontinuity models with an unknown point of discontinuity. One way to proceed is to adopt the methods developed for the estimation of structural breaks in time series (as reviewed in Hansen, 2001). Since we have to identify over 100 city-specific potential tipping points, however, we adopt an alternative approach based on the structure of our problem. The existing neighborhood transition literature, and our own analysis reported below, suggests that in most cities $E[\ \Delta(W_t/P_t) \mid m_{t-j}]$ is a downward-sloping function over most of the range of the initial minority share (e.g., between 0 and 70%). Assuming that $a(0) > E[\ \Delta(W_t/P_t)\ ] > b(\overline{m})$ for some $\overline{m} > m^*$ and that the a() and b() functions are relatively flat, it will be the case that

(7) $$\lim_{m \uparrow m^*} E[\ \Delta(W_t/P_t) \mid m\ ] > E[\ \Delta(W_t/P_t)\ ] > \lim_{m \downarrow m^*} E[\ \Delta(W_t/P_t) \mid m\ ],$$

[14] As will be seen in the next section, the shape of the b function at higher initial minority shares reflects mean-regressive behavior.

where the notation "$\lim_{m \downarrow m^*}$" means the limit as m approaches $m^*$ from above and "$\lim_{m \uparrow m^*}$" means the limit from below. Equation (7) expresses the simple observation that if there is a tipping point, then the white share in a neighborhood will grow faster than the average in the city as a whole to the left of the discontinuity, and slower than the city-wide average for neighborhoods to the right.

To use equation (7), we first smooth the data to obtain a well-behaved approximation, R(m), to $E[\ \Delta(W_t/P_t)\ |\ m\ ]$. We then choose as a *potential* tipping point the level of minority share m′ such that $R(m') = E[\ \Delta(W_t/P_t)\ ]$ and $R'(m')<0$. In cities where $R(m) - E[\ \Delta(W_t/P_t)\ ]$ has multiple roots, we choose the root for which R' is the most negative.[15] We are able to identify a potential tipping point by this procedure in nearly all cities. Note that this approach does not condition on there being a sharp discontinuity at the potential tipping point, but rather leaves the magnitude of the discontinuity as something to be estimated. Evidence of a discontinuity at the identified points then suggests that they are indeed tipping points, on average.

*b. Other Specification Issues*

Other features of a neighborhood may also affect the location choices of white families. Consideration of these confounding influences suggests a number of modifications to the basic

[15] We have found that a two-stage approach to estimating R(m) local to $m^*$ improves precision. First, we fit a city-specific regression of $\Delta(W_t/P_t)$ on a fourth-order polynomial in $m_{t-j}$, using tracts with initial minority shares below 60%. We identify the point at which the fitted regression curve passes through the city-level mean, when there are several candidates selecting the one through which the slope is most negative. Then we estimate a second fourth-order polynomial regression on a sample restricted to tracts with $m_{t-j}$ within ten percentage points of the identified point, and use the fitted values from this regression to obtain a more refined estimate of the fixed point. A more detailed description of our procedure is in the Data Appendix, and programs for its implementation are available from the authors.

model given by equation (6). A first issue is the dependent variable. The white population share can fall because whites move out or because nonwhites move in. To abstract from the latter, we define the change in the white share using a constant denominator:

$$\Delta(W_t/P_t) \ = \ (W_t - W_{t\text{-}j})/ P_{t\text{-}j} ,$$

i.e., the change in the number of white residents (or students) divided by the initial population of the neighborhood (or school). We also present analyses of the corresponding changes in the minority population, $(M_t - M_{t\text{-}j})/ P_{t\text{-}j}$ and the total population, $(P_t - P_{t\text{-}j})/P_{t\text{-}j}$.

A second issue is the definition of "minorities". For our main analysis, we define all non-whites as "minorities". Much of the older literature on neighborhood transition focuses on white reactions to blacks. Moreover, several recent studies have suggested that white mobility is particularly sensitive to the presence of blacks (e.g., Massey and Denton, 1991). We therefore present a supplementary analysis in which we redefine the minority fraction to include only African Americans or only African Americans and Hispanics (i.e. to excluding Asian Americans).

A final issue is the influence of other factors such as the average income of neighbors or schoolmates that may be correlated with the initial minority share. These can be readily accommodated. A generalized version of equation (6) is

$$(8) \qquad E[\ \Delta(W_t/P_t) \mid \ d] \ = \ f(d) + \beta * 1[d > 0] + \gamma * X,$$

where $d = m_{t\text{-}j} - m^*$, $f(d)$ is a smooth function (implemented as a polynomial), $\beta = b(m^*) - a(m^*)$, and X is a vector of other predetermined variables. In our analysis of neighborhood transitions we present models with controls for average household income, the local unemployment and poverty rates, and characteristics of the local housing market. We also explore specifications that include polynomial terms in these characteristics and in the distance to the nearest "minority" neighborhood. Fewer control variables are available for our analysis of school transitions, but we present a model that includes the fraction of students who receive a free or reduced price lunch.

An alternative approach to potential omitted variables bias in our regression discontinuity specification exploits geographic information, comparing a tract with $m_{t-j}$ just above $m^*$ with another nearby tract for which $m_{t-j}$ is just below $m^*$. This "within-neighbors" analysis is immune to bias from omitted variables that are smoothly distributed across space. Implementation is made somewhat tricky, however, by the spatial nature of the data. In contrast to, for example, between-twins comparisons for estimation of the return to education (Ashenfelter and Rouse, 1998), the "nearest neighbor" relationship is not transitive. As an alternative to fixed effect estimation, we include in (8) averages of the independent variables across the five closest neighboring tracts (measured from centroid to centroid and capping the distance at four miles). Denoting the average of 1[d>0] across i's neighbors (including tract i in the average) as $\bar{t}_{n(i)}$ and the average of the terms in the polynomial f(d) as $\overline{f(d)}_{n(i)}$, we estimate

(9) $$E[\ \Delta(W_t/P_t)\ |\ d_i, dbar_i] = f(d_i) + \overline{f(d)}_{n(i)} + \beta * 1[d_i>0] + \beta' * \bar{t}_{n(i)}$$

In this specification, β is the "within-neighbors" estimate of the tipping discontinuity. An estimate of $\beta \neq 0$ indicates that even differences within narrowly-defined groups of neighboring tracts in the beyond-$m^*$ indicator have impacts on the growth in the white population share.

The β' estimate is of independent interest: If tracts do not correspond perfectly to the "neighborhoods" that enter into residents' preferences, this would produce apparent spillovers in a tract-based analysis.[16] An indication that $\beta' \neq 0$ would suggest that these spillovers are potentially important. We find strong effects of having neighbors beyond the tipping point— indicating spillovers— but also that there is a discontinuous effect of the minority share in a given neighborhood after controlling for the composition of nearby tracts.

As a final test for the possibility of specification error, we conduct a parallel analysis of tipping in neighborhood poverty rates. Tract-level minority share is highly correlated with the poverty rate, raising the possibility that what appears to be race-based tipping might instead be capturing differences among neighborhoods of differing economic status. While the data offer some support for poverty-based tipping, this is less robust than the race-based version, and our estimates of race-based tipping are little affected by inclusion of poverty rate controls.

## V. Data and Potential Tipping Points

We discuss here our analyses of neighborhood transition, returning to our school-level data below in Section IX. We use Census tract data for the 1970-2000 Censuses from the Urban Institute's Neighborhood Change Database (NCDB).[17] For "cities" we use metropolitan statistical areas (MSAs) and primary metropolitan statistical areas (PMSAs) as defined in 1999. We require reasonably large samples within cities to estimate $E[\ \Delta(W_t/P_t) \mid m\ ]$, so our analysis is limited to MSAs with at least 100 tracts that can be merged between consecutive Censuses. In general, we use all the available tracts in the NCDB that lie within the 1999 boundaries of the included MSAs, although we exclude a few tracts with highly anomalous population changes

[16] This might be a particular issue in the later years of our tract-based analysis. Although tracts were initially drawn with the goal of corresponding to socially-understood neighborhoods, their boundaries may not have changed to keep up with shifts in true neighborhood boundaries. Thus, by 1990 there could be substantial slippage between the units used in our analysis and those that enter into residents' preferences.

[17] Census tracts are geographic units of approximately 4,000 people.

(see Data Appendix). In 1970, only the central areas of many MSAs were assigned to Census tracts, so our analysis for the 1970-80 period is largely based on tracts in central city areas. Additional outlying areas were assigned to tracts in 1980 and 1990, allowing us to include more suburban tracts in the analysis of later decades.

Tract boundaries are generally stable between census years but are sometimes revised, especially in fast-growing areas. Although the NCDB attempts to hold individual Census tracts constant at their 2000 boundaries, this is not always possible and there is little information about the quality of the matches. To test whether our results are sensitive to mis-matched tracts, we used a block-level crosswalk to construct our own panel of tracts between 1990 and 2000. Our results were robust to the use of this alternative panel, to the use of 1990 rather than 2000 tract boundaries as the basis for the analysis, and to a sample restriction that eliminated any tracts for which the 1990-2000 mapping was non-trivially imperfect.[18]

Table 1 presents some of the main characteristics of our primary sample of tracts from larger MSA's. The first row of the table shows the total number of tracts in the U.S. The number of defined tracts has risen steadily, from about 46,000 in 1970—nearly all metropolitan—to about 65,000—one fifth non-metropolitan—in 1990. Rows 2 and 3 report the number of MSAs in our sample and the number of tracts in those MSAs. There are 104 MSAs for which we can match at least 100 tracts between 1970 and 1980, 113 for 1980-1990, and 114 for 1990-2000. Rows 4-7 show average (unweighted) demographic characteristics of tracts in our samples.

[18] We use the criteria that the maximum error in the 2000 white share of the 1990 tract that could be induced by overlapping blocks was 2.5%. We also experimented with alternative treatments of multi-race individuals in the 2000 Census (which, unlike earlier censuses, allowed respondents to report more than one race). While the NCDB counts as "white" anyone who reports at least one of their races as white, our results are very similar when we limit the white population to those reporting only a single race.

The remainder of the table shows statistics for four subgroups of tracts, based on the fraction of minority residents in the base year: 0-5%, 5-20%, 20-70%, and 70% or higher. In 1970, just under half of all tracts, and a similar share of our larger-MSA subsample, had minority shares below 5%. By 1990, only a quarter of tracts in large MSAs had minority shares below 5%. The decline in "nearly all white" tracts was offset by growth in the other categories, particularly the 20-70% minority group (which rose from 15% to 27%) and the 70% or higher minority group (which rose from 7% to 16%).

Comparisons across the four groups of tracts show substantial growth in the white population in the two lower minority share subgroups but small or negative growth in the two higher minority share groups. The difference in 1970-1980 growth rates of the white share between tracts with 5-20% minority share in 1970 and those with 20-70% is about 49 percentage points, falling to 24 points in 1980-1990 and 22 in 1990-2000. This is consistent with potential tipping behavior around a 20% minority share. However, the table also shows that higher minority share tracts have lower family incomes, higher unemployment, and a higher fraction of multi-unit housing – factors that may confound the effects of racial composition.

The first step in our procedure for evaluating the importance of tipping behavior is to identify city-specific potential tipping points using the algorithm described in Section IV (a). Our visual inspection of graphs like Figure 1 suggests three basic patterns. Most cities share the features shown for Chicago in Figure 1: The fitted change in the white share drops sharply as it passes through the city-level average, and the frequency distribution of initial minority shares exhibits a prominent shoulder at about the same point. We interpret both patterns as strong evidence of tipping behavior.

In a second, much smaller group of cities the predicted relationship between the initial minority share and the change in the white share is more linear. Figure 5 shows data for one such city, Houston in the 1990-2000 period. The potential tipping point identified by our algorithm is 30%, but there is no obvious non-linearity in $E[\Delta(W_{2000}/P_{2000}) \mid m_{1990}]$. Patterns like this seem less consistent with tipping models.

Finally, there is a tiny minority of cities (4 between 1970 and 1980, 9 in 1980-1990, and 4 in 1990-2000) with non-monotonic, or even upward-sloping, relationships between the initial minority share and the change in the white share. Most such cities have anomalous demographic characteristics. An example is El Paso, Texas (Figure 6), which was 80% Hispanic in 1990 and had no tracts with minority shares below 25%. Tracts with higher 1990 minority shares exhibited *faster* growth in their white populations between 1990 and 2000. Our algorithm fails to identify a potential tipping point in such cases.[19]

Table 2 presents an overview of the potential tipping points. It shows the mean potential tipping points in each decade and the correlations among points in the same city in different decades. The average potential tipping point increased slightly between 1970-1980 (11.9) and 1990-2000 (13.8). Within a city, the potential tipping points appear to be quite stable over time, with correlation coefficients of 0.5 – 0.6 between decades.

## VI. Pooled Analysis

Having identified period-specific potential tipping points for most cities, we turn in this section to the question of whether the growth rate in the white population share changes

[19] We also exclude a very few cities where the algorithm selects a potential tipping point, but where visual inspection shows a highly non-monotonic relationship. See the Data Appendix for details.

discontinuously around the potential tipping point. Section VII explores the robustness of these results to a variety of alternative specifications, Section VIII extends the regression discontinuity analysis to examine the housing market correlates of tipping, and Section IX presents a parallel analysis of school-level tipping.

We begin with a graphical analysis that pools data from all of the cities in our sample. We deviate each tract's initial minority share in year t-10 from the potential tipping point for its city in that year. We also deviate the change in the white share in the tract between t-10 and t from the corresponding metropolitan-wide mean. Finally, we fit a local linear regression to the re-centered data.

Figures 7a-7c show the estimated relationships between the initial minority share in a Census tract (relative to the city-specific potential tipping point) and the change in the white share in the tract (relative to the city-wide average change) for 1970-1980, 1980-1990, and 1990-2000. The figures also show histograms of initial minority shares, which demonstrate that there is substantial density around the potential discontinuity point.[20] In all three decades the pooled data show two striking features. First, in each decade tracts with initial minority shares *below* the potential tipping point exhibit faster growth in their white population shares than the average for their cities. Second, tracts with initial minority shares *above* the potential tipping point exhibit a sharp relative decline in the white share over the next decade. These figures suggest that on average the white population share changes *discontinuously* at the potential tipping point—by about -30% in the 1970s, -25% in the 1980s and -20% in the 1990s—as predicted by a tipping model.

[20] Caution is required in interpreting the histograms at negative values of $m\text{-}m^*$, as the minimum possible value for a tract in a city with potential tipping point $m^*$ is $-m^*$.

Table 3 presents regression versions of the graphical analyses in Figures 7a-c, using the specification in equation (6). We explore two alternative functional forms for the $f(m_{t-10}-m^*)$ function. First, in column 1, we model f() as a global fourth order polynomial. Second, in column 2, we model f() as a pair of quadratic functions, one defined over positive values (i.e. over tracts with $m_{t-10}>m^*$) and the other over negative values. This approach allows the first and second derivatives of f() to vary discontinuously around $m_{t-10} - m^* = 0$, and estimates the discontinuity as the difference in intercepts between the two quadratics. In each case we include a full set of MSA fixed effects to capture differences across cities in white population growth rates, and we cluster standard errors on the MSA. Column 3 adds to the specification five tract level controls measured in the base year: the unemployment rate in the tract, the log of mean family income, and the (not mutually exclusive) fractions of single-unit, vacant, and renter-occupied housing units.

The estimates in Table 3 confirm that the growth rate of the white population share is discontinuous in the initial minority share around the potential tipping points. The models with fully-interacted quadratics for a(m) and b(m), in Column 2, yield precisely estimated discontinuities of -21% , -20%, and -11% for the 1970-1980, 1980-1990, and 1990-2000 periods, respectively. Estimates from the fourth order polynomial models are comparable. Adding the control variables shrinks the estimated discontinuities by perhaps a quarter, though they remain large and highly statistically significant.

The dependent variable for the models in Table 3 is the change in the number of white residents in a tract, divided by the initial population of the tract. Thus, the estimates show a discontinuity in the net growth in the white population of a tract when the minority share exceeds the tipping point. Table 4 presents parallel sets of estimates using as dependent variables the change in the number of non-white residents and the change in the total population, each measured as a share of the initial tract population. In each decade, tracts just beyond the tipping point saw only small increases in their non-white populations, and total populations thus shrunk (again, relative to the metropolitan area average) by nearly as much as did white populations. Tipping appears not to involve replacement of white with non-white populations but rather overall neighborhood decline.

## VII. Robustness Checks

We discuss in this section several alternative analyses that demonstrate the robustness of the results in Table 3. We begin with several approaches to evaluating bias from omitted variables that may covary with our tipping indicator, then conclude with estimates that use the tract fraction black—rather than fraction non-white—as the relevant composition variable

### *a. Means of Other Neighborhood Characteristics around the Tipping Point*

In a classic regression discontinuity (RD) design, the precise location of the "running variable" relative to the potential discontinuity is assumed to be as good as random within some appropriately narrow bandwidth. This assumption implies that other pre-determined variables should be continuously distributed around the discontinuity, just as baseline characteristics should be uncorrelated with treatment status in true random-assignment experiments. Evidence of discontinuities in these variables is interpreted as an indication that the identifying assumption fails (Lee, forthcoming).

Although tipping behavior leads to a model for neighborhood dynamics that shares some features with an RD design, it is important to note that neighborhood composition is a dynamic process rather than a single "experiment." In any period, the racial composition and other characteristics of neighborhoods are endogenous outcomes of interactions in previous periods. It should thus be expected that a variety of characteristics are discontinuous around the tipping point. To help clarify the expected patterns of neighborhood characteristics implied by a dynamic tipping model, imagine that residents of all neighborhoods have identical distributions of some characteristic X with mean $\mu_1$. In-movers to neighborhoods below the tipping point have the same mean, while minority in-movers to neighborhoods that have passed the tipping point have a lower mean $\mu_2$. In this simple case,

$$
\begin{aligned}
(7) \qquad E[X \mid m_t] &= \mu_1, && m_t \leq m^*, \\
&= (1 - (m_t - m^*))\,\mu_1 + (m_t - m^*)\,\mu_2, && m_t \geq m^*.
\end{aligned}
$$

Note that $E[X \mid m_t]$ is continuous around the tipping point, but its derivative is discontinuous. In other words, if we assume that the post-tipping in-movers are different from the people who would live in a neighborhood prior to tipping, then we should expect a "kink" in the relationship between the mean of X and the minority share, with the kink point at the tipping point.

Indeed, we see trend breaks in a variety of baseline characteristics around the potential tipping points. Figure 8 shows the relationships between baseline mean family incomes, poverty rates, unemployment rates, renter shares, and distance to the nearest "minority" tract—defined as a tract with $m \geq m^* + 10\%$—and the tract minority share relative to the city-specific tipping point in 1990. To place the variables on the same scale, we subtract the metropolitan mean from each variable and divide each by the global standard deviation before estimating local linear regressions for each. The estimated functions show clear evidence of trend breaks, with worsening measures to the right than to the left of the tipping point, as would be expected to occur as a result of in-movers with lower SES characteristics in tracts transitioning past the tipping point. Appendix Table 1 presents regression discontinuity-style models for each of these dependent variables.

These results suggest the possibility that the discontinuities estimated in Table 3 reflect not race-based tipping but the influence of omitted tract characteristics that themselves vary discontinuously around $m^*$. To examine this, Table 5 presents models for the growth in the white population share that include controls for $4^{th}$-order polynomials in the variables in Figure 8. The earlier results are entirely robust to the inclusion of the first four measures. When we control for the distance to the nearest minority tract, however, the coefficient on the tract racial composition falls by about a third, suggesting that some portion of the earlier tipping results reflected differential proximity to historically minority areas. Even when we control flexibly for this proximity, however, the tipping estimate is large and significant.

*b. Neighboring Tracts*

Our second approach to controlling for potential omitted variables bias in our main specification takes advantage of the fact that many potential omitted variables are likely to be smoothly distributed across space. If so, tracts located near each other will tend to have similar values of these variables, and comparisons between neighboring tracts will be relatively free of bias. As described in Section IV, we group each tract with up to five neighboring tracts, and augment the specifications in Table 3 with group averages of the minority share polynomial and tipping point indicator. When the specification is so augmented, the coefficient on the individual tract tipping indicator amounts to a within-group estimator of the tipping discontinuity. Table 6 summarizes the results of this exercise. Although inclusion of the neighborhood minority share variables somewhat reduces the magnitude of the "own-tipping" coefficients, they remain large and statistically significant. Over the 1980s, for example, when neighbors' minority shares are held constant, moving a tract beyond the tipping point causes it to lose 15.7 (= 11.1 + 27.5/6) percentage points in white share. The group average of the tipping indicator also has a strong negative effect. The estimates imply that the loss in white population share is 22.9 (=27.5 * 5/6, from column 3) percentage points larger in a tract with all 5 neighboring tracts beyond the tipping point than in tracts with no neighboring tracts beyond the tipping point, holding the tract's own status constant. This result is consistent with a simple measurement story in which the relevant neighborhood for a given household is some average of the immediately surrounding tract, and other nearby tracts.

*c. Is there Tipping in Poverty Rate?*

In order to determine whether there is evidence of "tipping patterns" in other outcomes, we performed a parallel analysis using the poverty rate in t-10 as the "running" variable. We were able to identify potential poverty rate tipping points in 91 of 104 cities in 1970-1980, 101 of 113 in 1980-1990, and 103 of 114 in 1990-2000. We augment the models in Table 3 with controls for polynomials in the initial poverty rate relative to the potential tipping point in that variable and an indicator for whether the tract's poverty rate exceeds the tipping point.

Table 7 reports the results of this exercise. The inclusion of a potential tipping point in the poverty rate does not diminish the estimated discontinuity around the tipping point in minority share. However, there also appears to be tipping in the poverty rate. For example, in the 1980s, we estimate a discontinuity in the poverty rate of -7.9 (column 4), which is statistically distinguishable from zero. The estimated discontinuity in the poverty rate appears generally smaller than the discontinuity in minority share. It is also more sensitive to the particular specification. In columns 3, 6, and 9, we present a model that includes cubic functions of (m-m*) and (p-p*)—where p is the poverty rate in the tract—on either side of m* and p*, respectively. This has little effect on the minority share discontinuity, but at least in 1970 it shrinks the poverty rate discontinuity and causes it to become insignificant. Thus, while the exercise provides intriguing evidence of potential tipping in non-race characteristics, it offers no indication that our results derive from specifications that privilege race over other neighborhood characteristics.

*d. Alternative Measures of Racial Composition*

Our analysis has assumed that tipping depends on the share non-white in the Census tract. This contrasts with much of the sociological literature—much of which pre-dates the recent wave of Hispanic immigration—which focuses on the fraction black. As a final specification test, we estimate parallel versions of our tipping models that use alternative measures of the neighborhood racial composition. We consider both the fraction black and the fraction black or Hispanic in the tract. In each case, we compute new potential tipping points for each MSA, then re-estimate our regression discontinuity model using these new points. In the fraction black analysis, the 1990-2000 potential tipping points correlate 0.43 with those from our main specification (though this rises to 0.74 when we exclude cities—primarily in the Southwest—where blacks are less than half of the non-white population). Potential tipping points for the combined black and Hispanic share are more highly correlated (around 0.8) with our original minority share points.

Table 8 presents the results of these analyses. When we consider the various potential tipping points separately, we find significant evidence of tipping around each in each decade, with one exception.[21] Columns 4, 8, and 12 present analyses that allow for all three tipping points simultaneously, along with paired quadratics in the deviation of each measure from its potential tipping point. In each case, our minority share tipping estimate is robust, though we find evidence for additional tipping in the black share and, in the 1990s, in the black or Hispanic share.

## VIII. Housing Markets

The above analyses have demonstrated that "white flight" accelerates discontinuously when a neighborhood's minority share exceeds a city-specific tipping point. Moreover, white exodus is not offset by growth in the minority population of the tract, so the total population grows less quickly than the MSA average in tracts beyond the tipping point. In this section, we examine the housing market correlates of tipping. Building on the model in Section III, we focus on home prices and on new construction, both of which are measured in our Census data. Our model suggested that the implications for each depend on whether tracts to the left of the tipping

[21] The exception, tipping in the combined black and Hispanic share in 1970, appears to reflect data problems. The 1970 Census data do not separately identify black non-Hispanics, so we must impute the fraction of the tract population that is either black or Hispanic. (We use a similar imputation—described in the Appendix—to identify the white non-Hispanic share for the earlier analyses, though this seems to work better.)

point will inevitably cross it as their housing stocks age, or whether new construction and replacement of older homes can keep a tract's minority share below the tipping point.

Figure 9 presents a pooled analysis of housing construction during the 1960s, 1970s, 1980s, and 1990s as a function of a tract's distance in 1990 from its city's 1990-2000 tipping point. We measure construction in each decade as the number of housing units built during that decade divided by the total number of units in the tract at the end of the decade. That is, the 1970 series depicts the share of 1970 homes built between 1960 and 1970.[22] We again deviate each measure from the metropolitan average. Tracts that were to the left of the 1990 tipping point see growth in their housing stock relative to their metropolitan areas of 7 percentage points during the 1990s. That is, if new construction in the metropolitan area between 1990 and 2000 amounts to 5% of the housing units that existed in 1990, a tract with $m<m^*$ will average 12% new construction. Tracts to the right of the tipping point, by contrast, grow at rates comparable to the metropolitan average. Interestingly, we see similar discontinuities in earlier decades, indicating that tracts to the right of the 1990 tipping point have had slower growth rates for at least 30 years than those to the left, though each of these discontinuities is smaller than that during the 1990s, and each seems to occur a bit to the right of the 1990 tipping point. All of these patterns are consistent with a model of anticipated tipping, in which developers are hesitant to invest in neighborhoods in which tipping seems likely.

Figure 10 presents a similar analyses of log mean values of owner-occupied homes in 1970, 1980, 1990, and 2000 as a function of a tract's distance from the 1990-2000 (left panel) or the 1970-1980 (right panel) tipping point. The former reveals price "effects" prior to tipping and

[22] We have also examined housing de-accession, measured, e.g., as the change between the 1990 and 2000 censuses in the number of housing units built before 1990. We found no evidence of tipping effects on de-accession.

the latter shows the long-run consequences of tipping, though of course it must be kept in mind that our earlier analyses indicate that the tipping phenomenon may have been different in the 1970s than the 1990s.[23] Focusing first on the left panel, there seems to have been a small difference in housing prices between tracts to the left and the right of the 1990 tipping point as far back as 1970, with prices already lower for the latter. The discontinuity grows as we move closer to 1990, with perhaps a small additional increment during the 1990s. This is again entirely consistent with the model of fully-anticipated tipping developed in Section III. The right panel of Figure 10 turns to tracts surrounding the 1970-1980 tipping point. It shows a large discontinuity in housing prices—approaching 20% —in 1970. Over the following 30 years, this difference in housing values persists but does not grow, suggesting again that tipping was anticipated by the housing market.

Table 9 presents regression analogues of the series in Figure 10. The first row examines discontinuities around the 1990-2000 tipping point. As in the graphical presentation, we see small discontinuities in 1970 housing values that grow in each successive decade. The final three columns show estimates for the change in mean housing values. Essentially all of the tipping discontinuity in the 1970-2000 change in prices seems to have occurred before 1990, as would be expected if tipping is anticipated by 1990. The second row shows similar analyses for the 1970-1980 tipping point. It again echoes the graphical presentation, with large discontinuities in 1970 values, slight growth between 1970 and 1980, and if anything a closing of the discontinuity between 1980 and 2000.

[23] For some purposes, one might want to exclude from these long-run analyses tracts that were redeveloped or abandoned at some point between 1970 and 2000. When we exclude tracts that saw 10-year changes in the occupied housing stock greater than 50% (in either direction) at any point in the period, we see smaller discontinuities around the 1990-2000 tipping point but obtain similar results for the 1970-1980 discontinuity.

## IX. Schools

Thus far, we have focused on residential tipping, and the evidence has been clear that the racial composition of a neighborhood plays an important, discontinuous role in determining white population growth. One plausible determinant of white preferences over neighborhood composition is that most schools draw from the neighborhood population, so a concern over race-driven peer effects in educational production might lead whites to prefer neighborhoods with lower minority shares. In this section, we use the Common Core of Data, an annual census of public schools with information on school racial composition since the 1987-88 school year, to analyze tipping points in elementary school composition. An important advantage of the school-level analysis is that, while census tracts may or may not correspond with subjective "neighborhoods," the school is a natural unit over which parents' concerns about student composition may extend.

The analysis parallels that of neighborhoods, though we restrict attention to 81 metropolitan areas with at least 100 elementary schools. Figure 11 displays the estimated relationship between the 1990 minority share in an elementary school (relative to the city-specific potential tipping point) and the change in the white population in the school between 1990 and 2000, expressed as a share of the 1990 enrollment and deviated from the MSA-level average.[24] As with neighborhoods, there is a clear discontinuity, in excess of ten percentage points, in white enrollment growth at the potential tipping point.

Table 10 reports regression discontinuity models for school-level tipping. When we

[24] As before, we exclude a small number (4) of cities for which we are unable to identify a potential tipping point.

control for a fourth order polynomial in initial enrollment share relative to the potential tipping point, the discontinuity is estimated as -9.4. This falls slightly when we instead control for separate quadratics on either side of the tipping point (Column 2), but is unaffected when we also control for the fraction of students in the school that receive free lunches (Column 3). Columns 4-6 present analogous models for the school's total enrollment, and again show large discontinuities: Schools with minority shares in excess of the city-level tipping point appear to lose white pupils accounting for about 9% of their 1990 enrollment, a deficit that is not made up for by minority inflows.

The data indicate clear tipping points in schools as well as neighborhoods. However, the school-based potential tipping points are only weakly correlated (r=0.34) with those for neighborhoods. One reason may be that, unlike neighborhoods, schools may be affected by intentional court- or district-managed desegregation plans (e.g. busing, rezoning, etc.) that serve to minimize tipping. In future work, we hope to exploit changes in regulatory regimes coming from court rulings that release districts from desegregation orders (Lutz, 2005) to investigate the role of desegregation efforts in school-level tipping.

## VI. Attitudes of Whites and the Location of the Tipping Point

We have argued that a highly non-linear "inverse S" relationship between the initial minority share in a Census tract or school and the subsequent change in the white population share provides evidence of a tipping point, driven by strategic interactions between the choices of white families. To provide more direct evidence on the link between tipping and white families' preferences, we use information on the racial attitudes of white residents in different cities from the General Social Survey (GSS). Simple theoretical models imply that the tipping point will shift to the right if whites are more tolerant of integrated neighborhoods. We therefore test whether our estimated city-specific tipping points are higher in cities with more racially tolerant whites, controlling for other characteristics of the city.

The annual GSS samples are small, and the survey instrument changes substantially from year to year. To develop a reasonably reliable index of white attitudes, we pooled GSS data from 1975 to 1998 and selected white respondents who could be matched to MSAs.[25] We used four questions that elicit direct information on preferences regarding contact between races and that have been asked relatively frequently, coding each into a binary measure of prejudice:

I: Do you think there should be laws against marriages between blacks and whites?

II: In general, do you favor or oppose the busing of black and white school children from one school district to another?

III. How strongly do you agree or disagree with the statement: "White people have a right to keep blacks out of their neighborhoods if they want to, and blacks should respect that right"?

IV. Suppose there is a community wide vote on the general housing issue. Which (of the following two) laws would you vote for:

A. One law says that a homeowner can decide for himself whom to sell his house to, even if he prefers not to sell to blacks.

B. The second law says that a homeowner cannot refuse to sell to someone because of their race or color

[25] The GSS uses a geographically stratified sample, with changes in the sampling frame in 1983 and 1993. The mapping from Primary Sampling Units to MSAs is necessarily approximate. Note also that in many cases only a subset of an MSA is in the GSS sample.

For each question, we estimated a linear regression of the responses on year dummies, MSA dummies, and a set of controls for the characteristics of the respondent (age, gender, and education). We then standardized the estimated MSA effects to have mean 0 and standard deviation 1. As reported in Appendix Table 2, the MSA-average responses to the four questions are reasonably highly correlated (with correlations between 0.26 and 0.52). We formed a racial attitudes index as the average of the standardized MSA effects from the four questions.[26] This index has standard deviation 0.72.

We were able to construct a value of the index for 66 MSAs in our tipping sample; in these MSAs, we had an average of approximately 38 GSS responses per question on questions II-IV, and 50 responses on question I. City-specific values of the index are reported in Appendix Table 3. The cities with highest values of the index (indicating more strongly held views *against* racial contact) are Memphis (value=1.65), Knoxville (1.53), Birmingham (1.43), and New Orleans (1.32). The cities with lowest values of the index are Rochester (-1.08) San Diego (-1.07) and Worcester (-1.11).

Table 11 reports a series of models that relate the location of the tipping point for a city (the average of our 1980-90 and 1990-2000 points) to city characteristics.[27] For reference, the first column shows the mean of each of the independent variables. The first regression model includes only the attitudes index, and the second adds controls for the fractions of blacks and Hispanics in the city. The latter variables have strong effects, with coefficients around 0.4, suggesting that tipping points are much higher (but not proportionately so) in cities with higher

[26] We also explored using the principal component of the four sets of MSA effects. This put approximately equal weight on each factor.

[27] We exclude 1970-80 from this average to maximize the available sample, as some cities had too few tracts in 1970 to be used for our tipping analyses. Inclusion of the earlier data does not affect the qualitative results.

minority shares.  Controlling for minority composition, there is a negative but statistically insignificant relationship between the location of a city's tipping point and the index of attitudes. Adding region controls (column 3) leads to a more negative and more precise significant coefficient on the attitudes index.  Column 4 adds the log mean incomes of blacks, Hispanics, and whites in the city.  Higher white incomes are associated with lower tipping points, and the inclusion of income controls considerably strengthens the attitude effect.

Given the small set of cities for which the attitudes index is available, we explore the additional control variables in several sets.  Column 5 adds the fractions of black and white adults without high school diplomas.  Column 6 adds controls for two "structural" characteristics of the local school system that may influence residential location decisions:  the fraction of 5-12 year olds in private school and a Herfindahl index measuring the concentration of students across school districts (Hoxby 2000; Rothstein, forthcoming).  Finally, column 7 explores two controls for structural characteristics of the housing stock in an MSA: The population density and the rate of new construction (measured as houses built between 1985 and 1990 divided by total houses in 1990).  In each case, the additional variables have small, insignificant effects on the location of the tipping point, and the attitudes coefficient is largely unaffected by their inclusion.

To understand the magnitude of the effect implied by the models in Table 11, consider the difference between a city in which whites have strong views against inter-racial contact (e.g. Memphis) and one where whites are relatively tolerant (e.g., San Diego).  The difference in the attitudes index between these cities is 2.7.  Multiplying this by a coefficient of -5 implies that the tipping point is shifted to the right by about 13.5 percentage points.  Compared to a mean tipping point (averaged between 1980-1990 and 1990-2000) of 12.1% and a standard deviation of 7.2%, this is a large effect.  Assuming the same -5 coefficient, a standard deviation change in the value of the attitudes index implies a 3.6 percentage point rise in the tipping point, or a 0.5 "effect size".

## VII. Conclusion

One longstanding explanation for the prevalence and persistence of racial segregation in is that white families are unwilling to live in neighborhoods, or send their children to schools, with large minority shares.  Schelling (1971) demonstrated that these sorts of preferences could give rise to "tipping points," such that neighborhoods whose minority shares exceed these points experience sharp declines in their white populations.  Modern regression discontinuity techniques are well suited for estimation of tipping behavior.  Applying them, we find strong evidence for well-defined tipping points for both neighborhoods and schools.  Although the extent of tipping declined between the 1970s and 1990s, it remains statistically and practically significant.

Several alternative specifications indicate that tipping behavior reflects the influence of neighborhood racial composition per se rather than that of omitted neighborhood-level variables.  The location of the city-specific tipping point is moreover robustly correlated with survey-based estimates of white attitudes about integration, reinforcing the inference that tipping reflects white families' preferences over the racial composition of their neighbors. [28]  Finally, an analysis of

[28] It is important to mention an alternative potential explanation, however.  Lower tipping points mean that white families in the MSA have less exposure to minority neighbors.  If unfamiliarity breeds contempt, it may be that causation runs from the tipping point to attitudes rather than the reverse.  Given the small changes in tipping points over time and our limited ability to measure changes in a city's attitudes index, we are unable to test this alternative.

home values suggests that tipping is anticipated and capitalized at least a decade before it actually occurs.

## References

Ashenfelter, Orley and Cecilia E. Rouse (1998). "Income, Schooling, and Ability: Evidence from a New Sample of Identical Twins." *Quarterly Journal of Economics*, 113 (February), pp. 253-284.

Baily, Martin J. (1959). "Note on the Economics of Residential Zoning and Urban Renewal." *Land Economics* 35 (August), pp. 288-292.

Bobo, Lawrence and Camille L. Zubrinsky (1996). "Attitudes on Residential Integration: Perceived Status Differences, Mere In-Group Preference, or Racial Prejudice?" *Social Forces* 74 (March), pp 883-909.

Bond, Eric W. and N. Edward Coulson (1989). "Externalities, Filtering, and Neighborhood Change." *Journal of Urban Economics* 26 (September), pp. 231-249.

Brueckner, Jan. (1977). "The Determinants of Residential Succession." *Journal of Urban Economics* 4 (January), pp. 45-59.

Clotfelter, Charles T. (2001). "Are Whites Still Fleeing? Racial Patterns and Enrollment Shifts in Urban Public Schools, 1987-1996." *Journal of Policy Analysis and Management* 20 (Spring), pp. 199-221.

Coleman, James S., S. Kelley, and J. Moore (1975). "Trends in School Segregation, 1968-73." Urban Institute Paper No. 772-03-91, August.

Denton, Nancy A. and Douglas S. Massey (1991). "Patterns of Neighborhood Transition in a Multiethnic World: U.S. Metropolitan Areas, 1970-80." *Demography* 28 (February), pp. 41-63.

Duncan, Otis Dudley and Beverly Duncan (1957). *The Negro Population of Chicago: A Study of Residential Succession*. Chicago: University of Chicago Press.

Easterly, William (2005). "Empirics of Strategic Interdependence: The Case of the Racial Tipping Point." New York University DRI Working Paper No. 5, October.

Farley, Reynolds, Howard Schuman, Suzanne Bianchi, Diane Colasanto and Shirley Hatchett (1978). "Chocolate City, Vanilla Suburbs: Will the Trend Toward Racially Separate Communities Continue?" *Social Science Research* 7 (December), pp. 319-344.

Farley Reynolds, Charlotte Steeh, Tara Jackson, Maria Krysan, and Keith Reeves (1993). "Continued Racial Residential Segregation in Detriot: 'Chocolate City, Vanilla Suburbs Revisited'." *Journal of Housing Research* 4 (1), pp 1-38.

Granovetter, Mark (1978). "Threshold Models of Collective Action." *American Journal of*

*Sociology* 83 (May), pp. 1420-1443.

Grodzins, Morton (1957). "Metropolitan Segregation." *Scientific American* 197 (4), pp. 33-41.

Grodzins, Morton (1958). *The Metropolitan Area as a Racial Problem.* Pittsburgh: University of Pittsburgh Press.

Hansen, Bruce E. (2001) "The New Econometrics of Structural Change: Dating Breaks in U.S. Labor Productivity." *Journal of Economic Perspectives* 15 (Fall), pp. 117-128.

Heal, Geoffrey and Howard Kunreuther (2006). "Supermodularity and Tipping." National Bureau of Economic Research Working Paper #12281 (June).

Hoxby, Caroline M. (2000). "Does Competition Among Schools Benefit Students and Taxpayers?" *American Economic Review* 90 (December), pp. 1209-1238.

Kanemoto, Yoshitsugu (1980). "Externality, Migration, and Urban Crises." *Journal of Urban Economics* 8 (September), pp. 150-164.

Lee, David S. (forthcoming) Randomized Experiments from Non-random Selection in U.S. House Elections." *Journal of Econometrics*.

Lee, Barrett A. and Peter B. Wood (1991). "Is Neighborhood Racial Succession Place-Specific?" *Demography* 28 (February), pp. 21-40.

Lutz, Byron F. (2005). "Post Brown vs. the Board of Education: The Effects of the End of Court-Ordered Desegregation." Working paper, Finance and Economics Discussion Series 2005-64. Washington: Board of Governors of the Federal Reserve System.

Mare, Robert D. and Elizabeth E. Bruch (2003). "Spatial Inequality, Neighborhood Mobility, and Residential Segregation." California Center for Population Research, University of California, Los Angeles, Paper CCPR-002-03. Available at http://repositories.cdlib.org/ccrp/olwp/CCPR-002-03.

Miyao, Takahiro (1979). "Dynamic Stability of an Open City with Many Household Classes." *Journal of Urban Economics* 6 (July), pp. 292-298.

Muth, Richard (1973). "A Vintage Model of the Housing Stock." *Regional Science Association Papers and Proceedings* 30, pp. 141-156.

Park, Robert E., Ernest W. Burgess, and Roderick D. McKenzie (1925). *The City.* Chicago: University of Chicago Press.

Patterson, Orlando (1997). *The Ordeal of Integration: Progress and Resentment in America's Racial Crisis*. Washington, DC: Civitas/Counterpoint.

Quillian, Lincoln (2002). "Why Is Black-White Residential Segregation So Persistent? Evidence on Three Theories from Migration Data." *Social Science Research* 31 (June), pp. 197-229.

Rapkin, C. and W. Grigsby (1960). *The Demand for Housing in Racially Mixed Areas*. Berkeley: University of California Press.

Reber, Sarah J. (2005). "Court-Ordered Desegregation: Successes and Failures Integrating American Schools Since *Brown versus Board of Education*." *Journal of Human Resources* 40 (November), pp. 559-590.

Rothstein, Jesse (forthcoming). "Does Competition Among Public Schools Benefit Students and Taxpayers? A Comment." *American Economic Review*.

Schelling, Thomas C. (1969). "Models of Segregation." *American Economic Review* 59 (May), pp. 488-493.

Schelling, Thomas C. (1971). "Dynamic Models of Segregation." *Journal of Mathematical Sociology* 1 (July), pp. 143-186.

Schelling, Thomas C. (1978). *Micromotives and Macrobehavior*. New York: Norton.

Schwirian, Kent P. (1983). "Models of Neighborhood Change." *Annual Review of Sociology* 9, pp. 83-102.

South, Scott J. and Kyle D. Crowder (1998). "Leaving the 'Hood: Residential Mobility between Black, White, and Integrated Neighborhoods." *American Sociological Review* 63 (February), pp. 17-26.

Taeuber, Karl E. and A. R. Taeuber (1965). *Negros in Cities*. Chicago: Aldine Press.

Thernstrom, Stephan and Abigail Thernstrom (1997). *America in Black and White: One Nation, Indivisible*. New York: Simon and Schuster.

Welch, Finis and Audrey Light (1987). *New Evidence on School Desegregation*. Washington, DC: U.S. Civil Rights Commission.

## Data Appendix

*Sample for identification of candidate tipping points*

The sample that is used to identify the candidate tipping points for neighborhoods is from the Urban Institute's Neighborhood Change Database (NCDB). We assign each tract to the 1999 MSA in which it lies. We exclude tracts where the population growth rate exceeds five standard deviations from the MSA mean growth rate, tracts with fewer than 200 residents in the base year, and tracts where the ten-year growth in the white population exceeds 500% of the base-year total population. We focus on MSAs for which we still have 100 matched tracts after these exclusions. For each of these MSAs, we use the procedure identified in the text to identify a candidate tipping point.

We define the white population as the number of non-Hispanic whites, and minorities as all other residents. Because the 1970 data do not separately identify white Hispanics and non-Hispanics, we impute the number of white/non-Hispanics in each tract using information on the share of black, white and Hispanic household heads. Specifically, we use 1980 data to estimate a regression of white/non-Hispanic share in a tract on the black share, white share, and Hispanic share. The R-squared of this regression is 0.99. Using the coefficient estimates from this regression and 1970 data on the tract's white share, black share, and Hispanic share in 1970, we predict the 1970 non-Hispanic white share, censoring predicted values at 0 and 1. When we compute changes in the non-Hispanic white population between 1970 and 1980, we use fitted values in both years. We use a similar imputation procedure to identify the number of non-Hispanic blacks in each tract in 1970 for our analysis of alternative tipping points in Table 8.

*Analysis sample*

Once candidate tipping points are identified, we use slightly less restrictive samples for the remaining analysis, adding back tracts with base-year populations between 100 and 200 and tracts with changes in white population shares between 500 and 800%. We exclude MSAs for which we are unable to identify a candidate tipping point under 50%. We also exclude a few cities for which our algorithm identified a candidate point but a visual inspection reveals the R(m) function to be highly non-monotonic. These are El Paso, Honolulu, Miami, Nassau-Suffolk (NY), and Vallejo-Fairfield-Napa (CA) for the 1980-1990 analysis, and Jersey City for the 1990-2000 analysis.

**Figure 1. Neighborhood change in Chicago, 1970-1980, and the 1970 distribution of neighborhood minority shares**

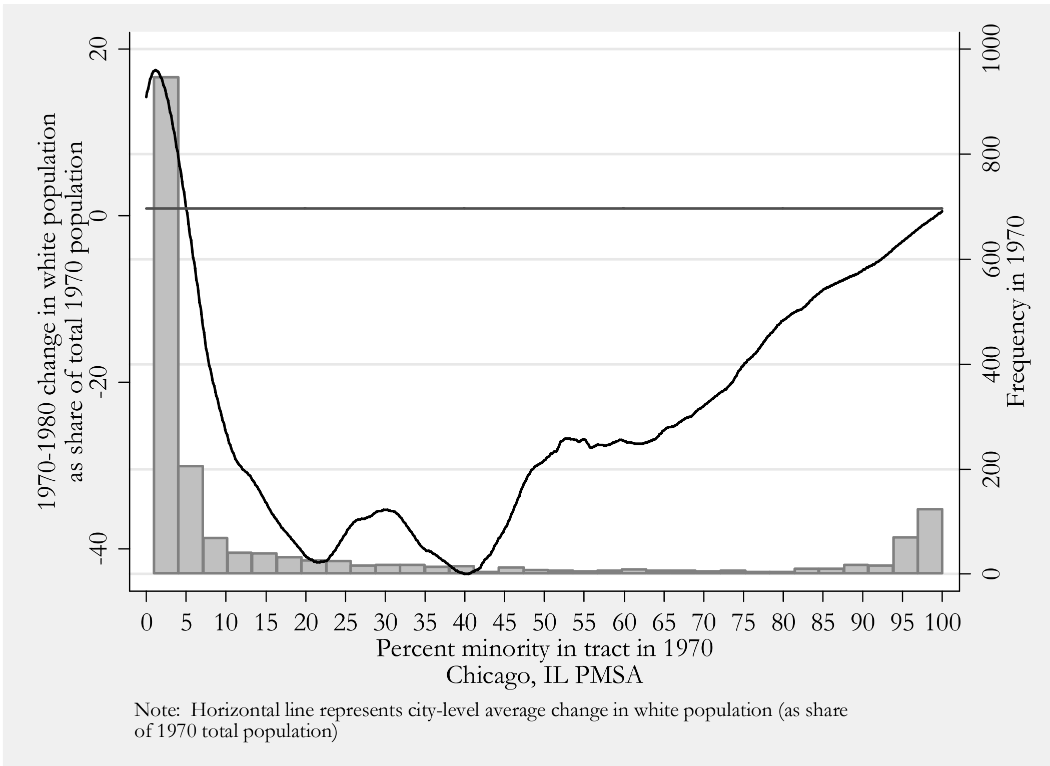



**Figure 2. A Tipping Point in the Schelling Model**

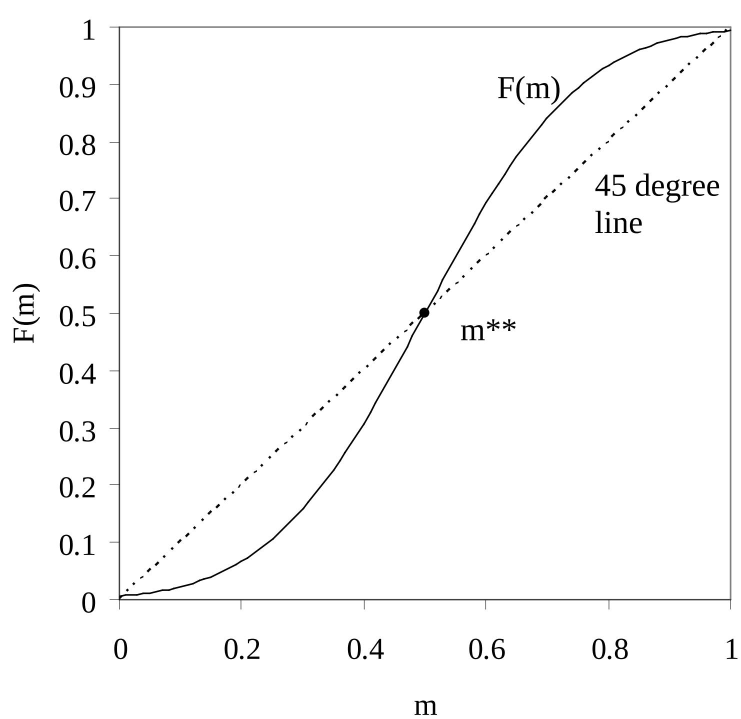


**Figure 3: Three equilibria in the Bond & Coulson tipping model**

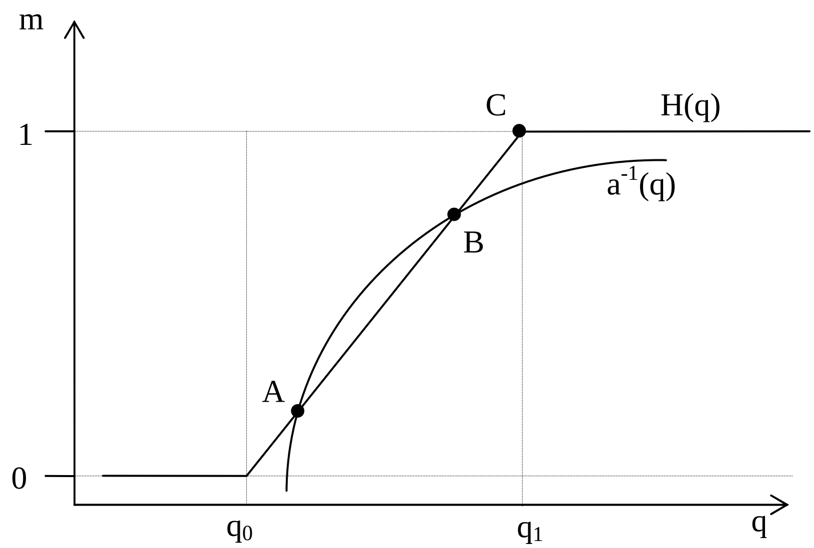



**Figure 4: Filtering in the Bond/Coulson Model**

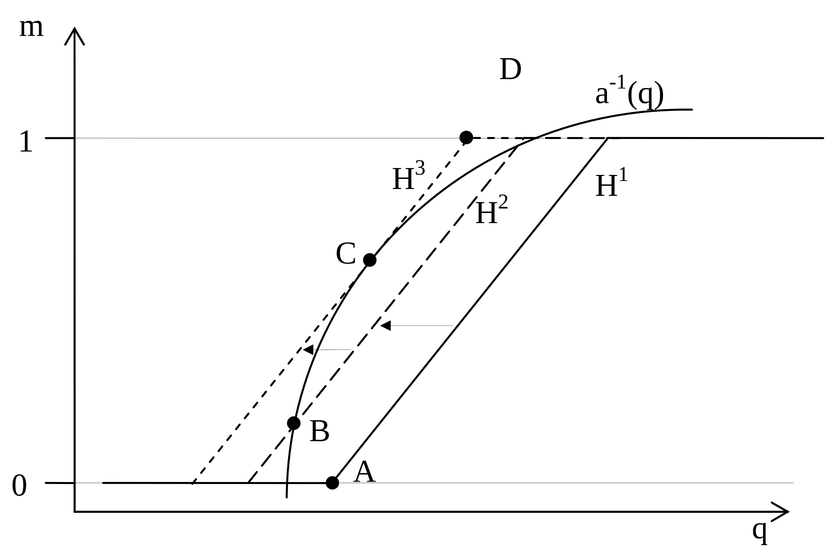


**Figure 5: Neighborhood change in Houston, 1990-2000, and the 1990 distribution of neighborhood minority shares**

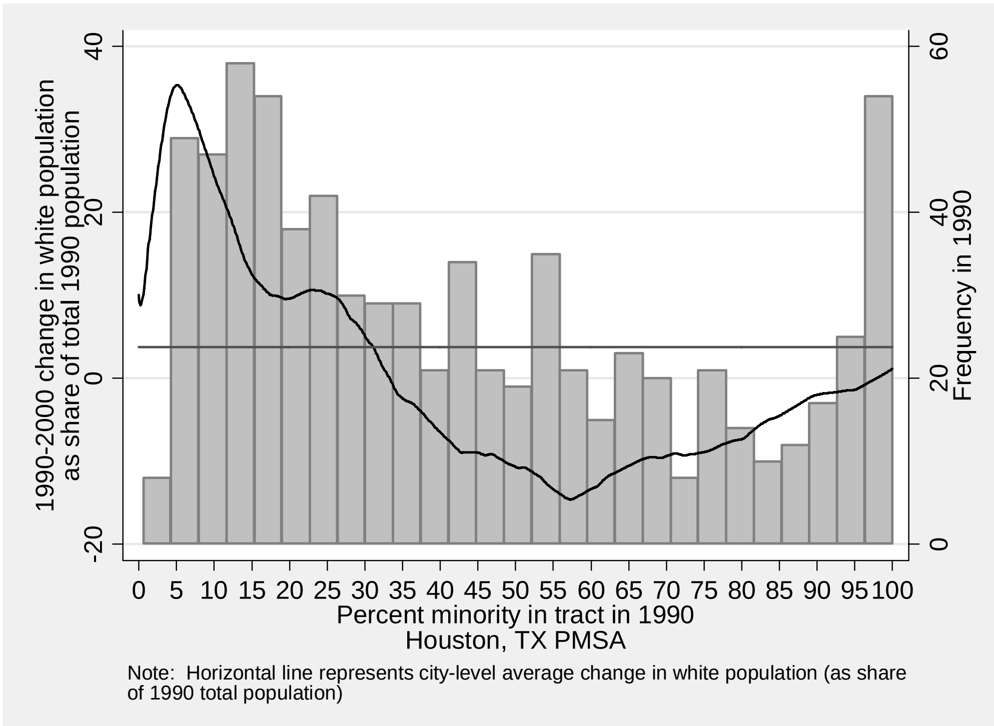



**Figure 6: Neighborhood change in El Paso, 1990-2000, and the 1990 distribution of neighborhood minority shares**

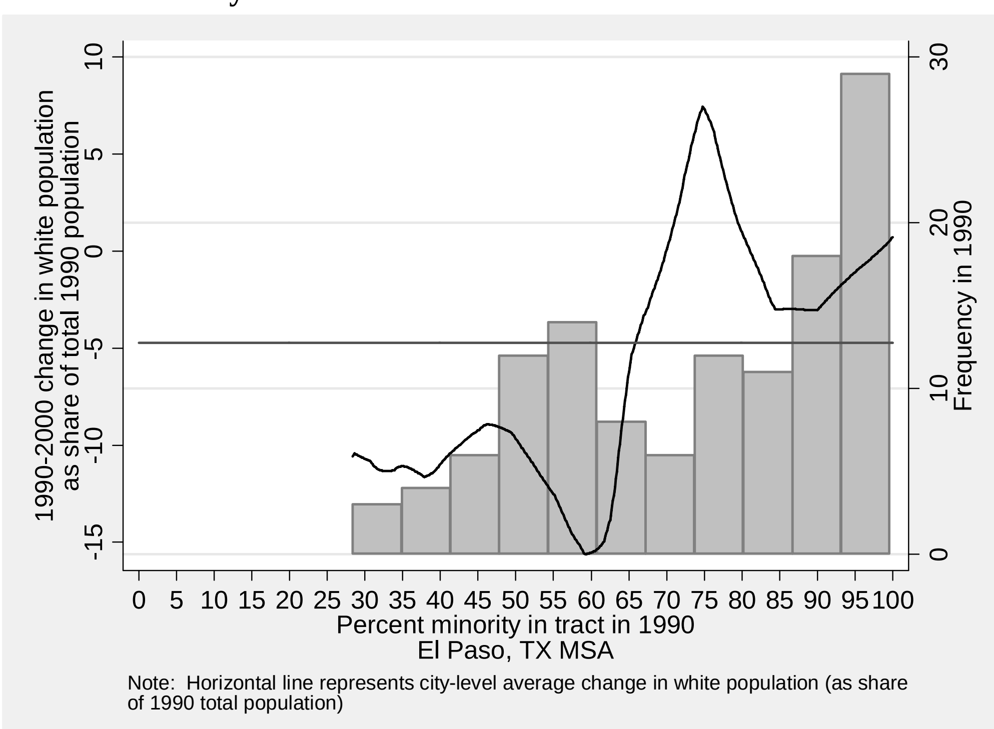
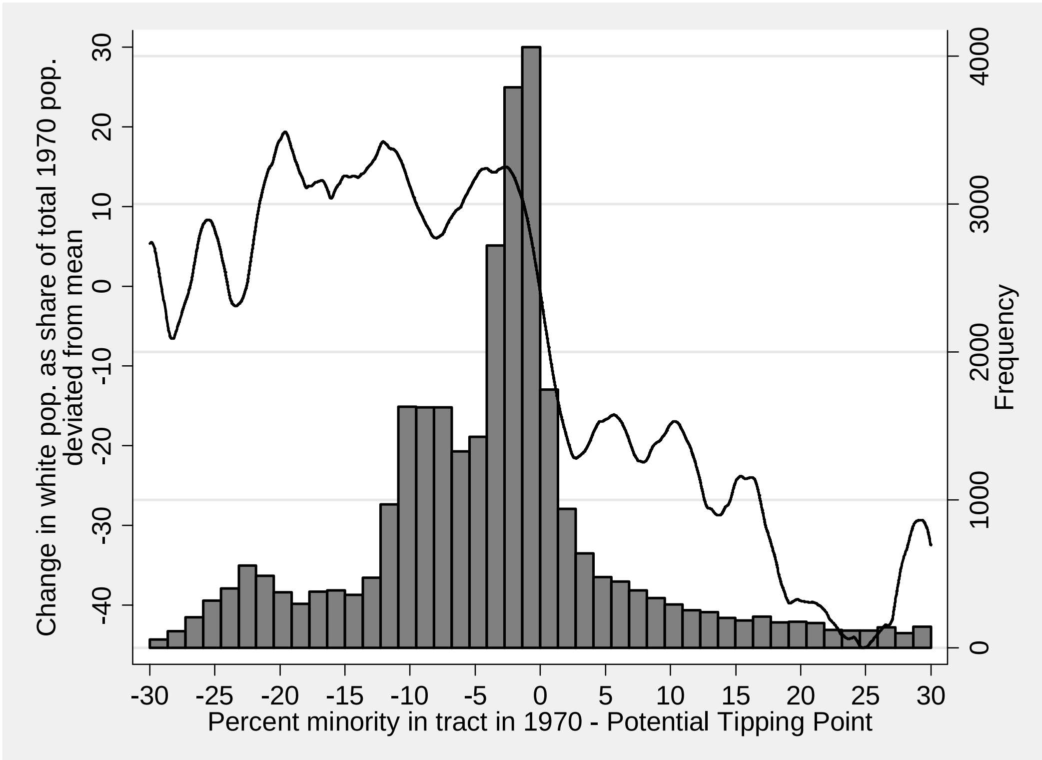
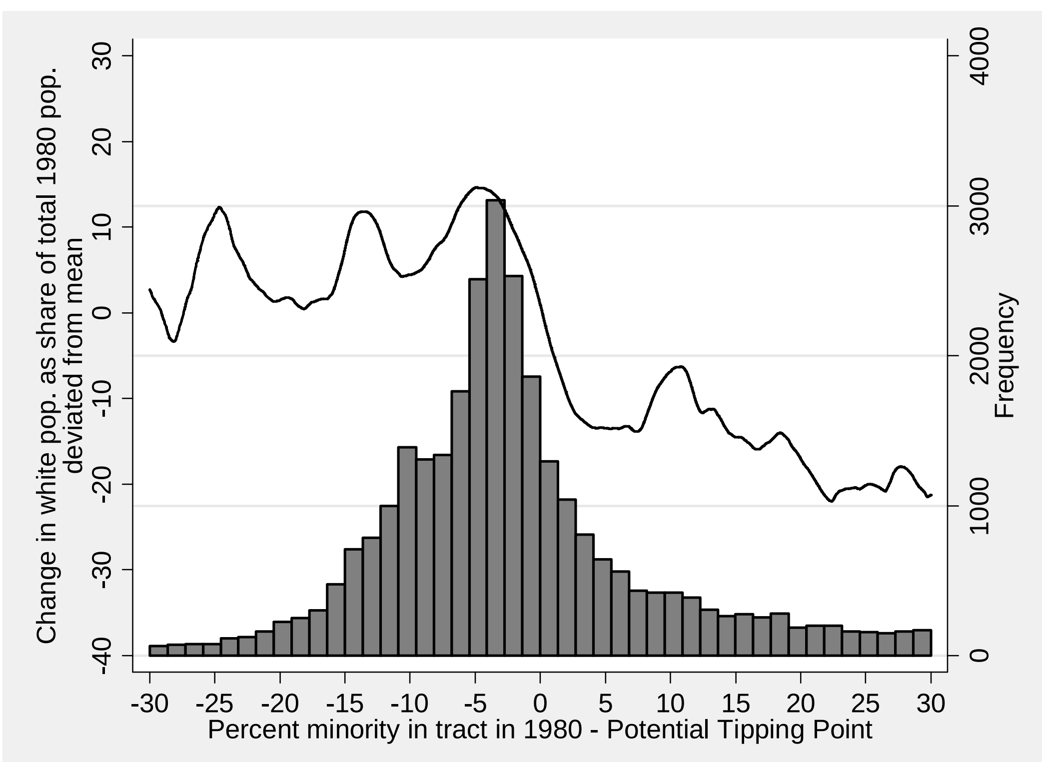


**Figure 7: Neighborhood change in a pooled sample of metropolitan tracts, by relationship to candidate tipping point**

**7A. 1970-1980**




**7B. 1980-1990**



**7C. 1990-2000**

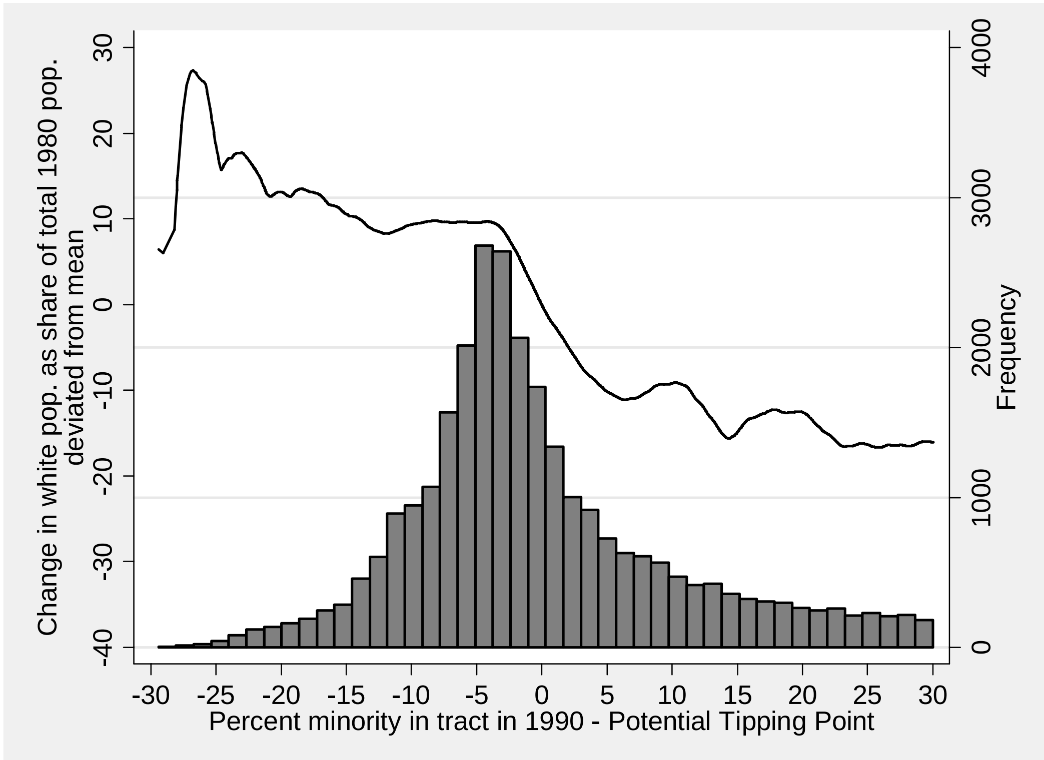



**Figure 8. 1990 neighborhood characteristics, by relationship to 1990-2000 candidate tipping points**

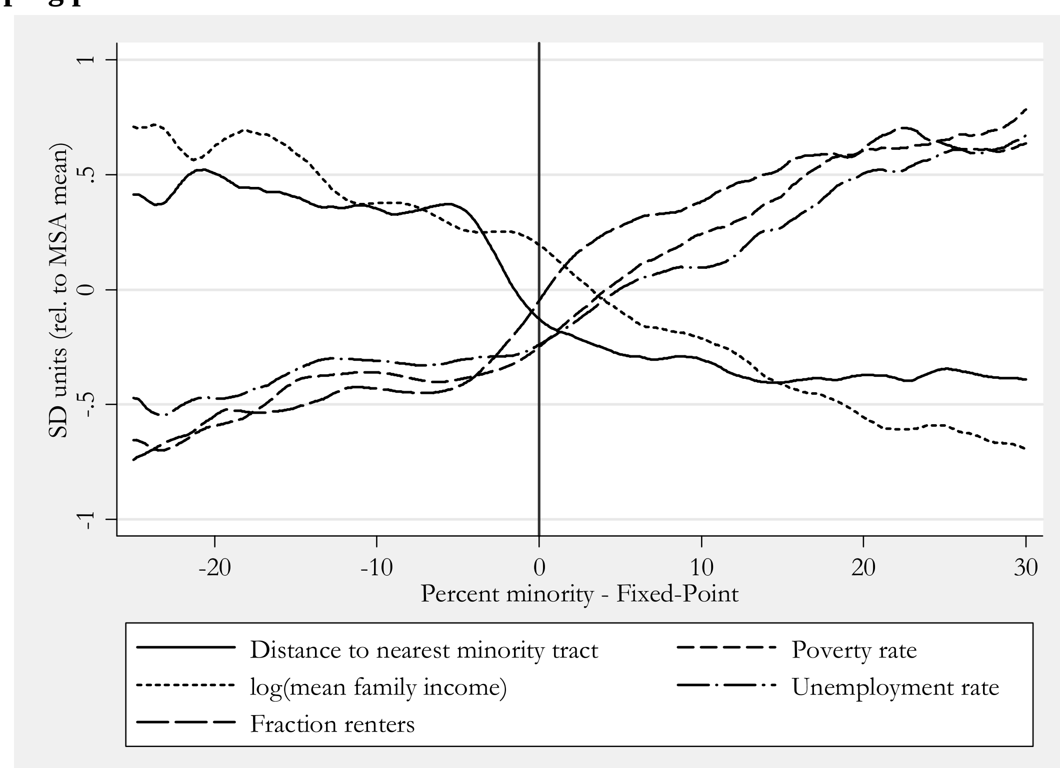


**Figure 9. Housing construction in the 1960s, 1970s, 1980s, and 1990s, by 1990 minority share relative to candidate tipping point**

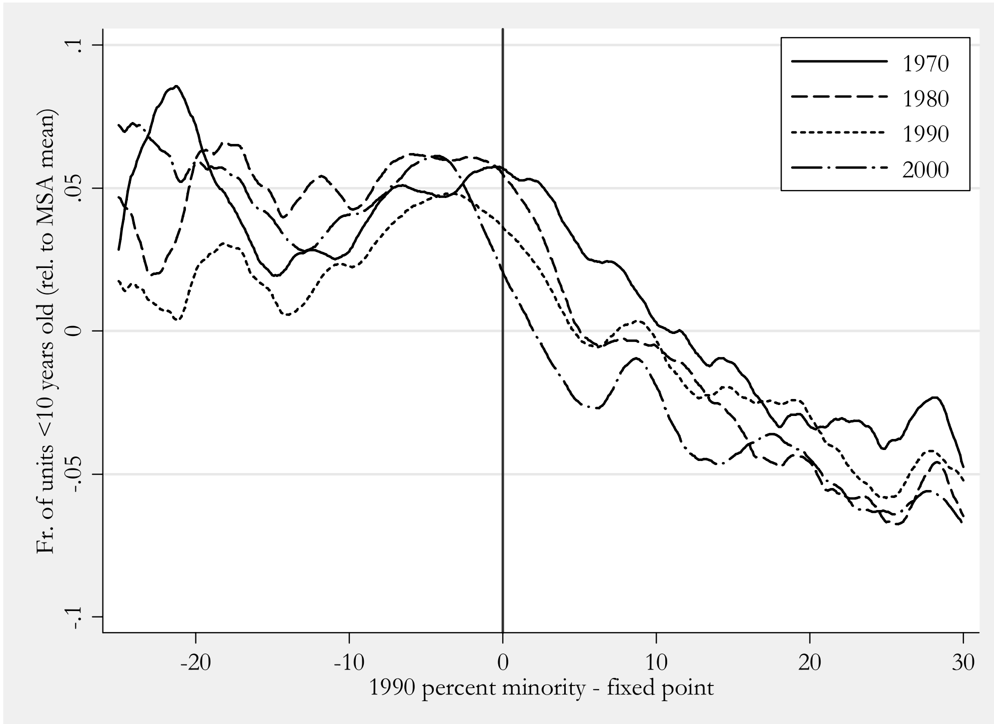



**Figure 10: Log mean housing prices in 1970, 1980, 1990, 2000, by minority share in 1990 and 1970**

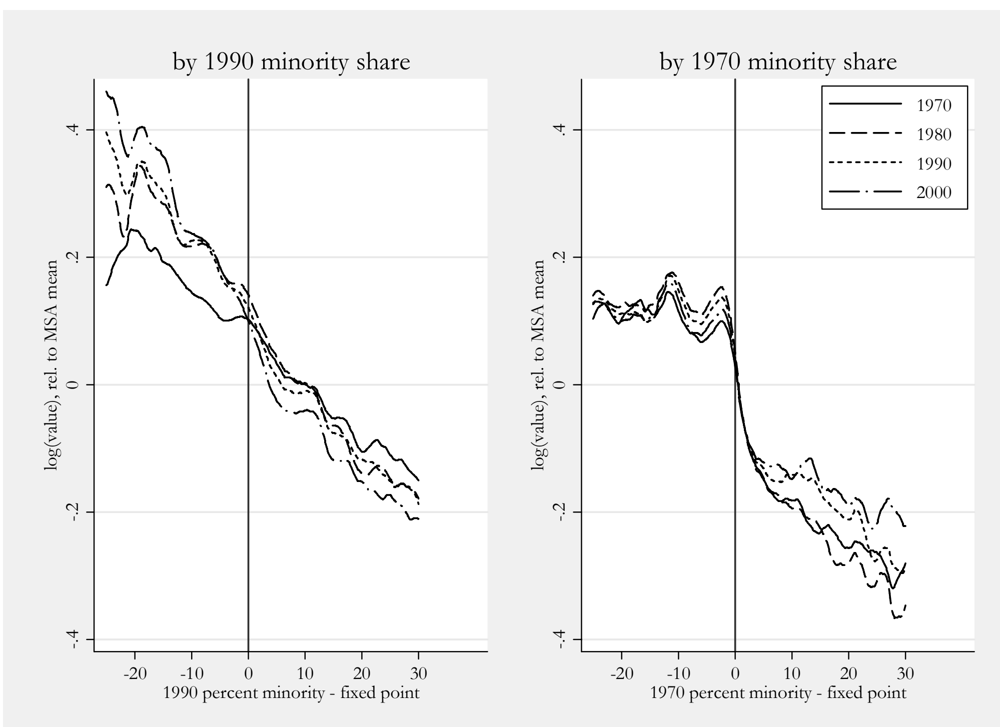


**Figure 11: White enrollment change in elementary schools, 1990-2000**

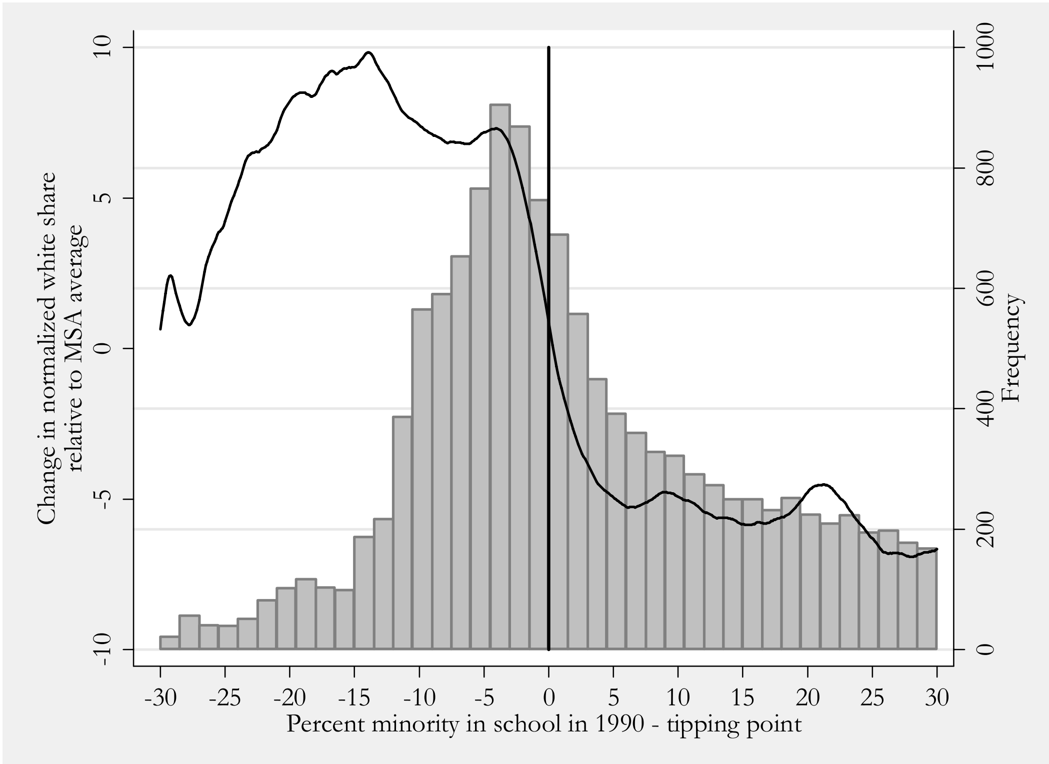


**Table 1: Summary statistics for metropolitan census tracts**

| | 1970 | 1980 | 1990 |
|---|---|---|---|
| # of tracts in country | 46,334 | 51,857 | 64,891 |
| # of tracts in 1999 MSAs | 45,636 | 49,896 | 51,037 |
| | | | |
| # of tracts in sample for 10-year comparisons | 36,684 | 39,663 | 40,570 |
| # of MSAs in sample | 104 | 113 | 114 |
| Mean % minority | 16.3 | 23.4 | 28.9 |
| Mean family income (nominal) | $12,148 | $24,666 | $47,224 |
| Unemployment rate | 4.3 | 6.6 | 6.8 |
| % single family homes | 68.2 | 64.9 | 64.0 |
| | | | |
| *0-5% minority in base year:* | 22,899 | 18,665 | 20,470 |
| # in MSAs | 22,415 | 17,896 | 17,439 |
| # in sample | 17,529 | 13,139 | 9,494 |
| Mean family income | $13,460 | $27,634 | $52,290 |
| Unemployment rate | 3.5 | 5.5 | 4.6 |
| % single family homes | 75.1 | 75.8 | 77.7 |
| Growth in total population, t to t+10 (%) | 46.6 | 25.4 | 23.1 |
| Growth in white population (as % of base year population), t to t+10 | 39.2 | 20.9 | 17.9 |
| | | | |
| *5-20% minority:* | 13,524 | 17,258 | 20,755 |
| # in MSAs | 13,481 | 16,755 | 17,439 |
| # in sample | 11,176 | 13,462 | 13,839 |
| Mean family income | $12,300 | $27,169 | $55,920 |
| Unemployment rate | 4.3 | 5.1 | 4.6 |
| % single family homes | 67.1 | 66.7 | 67.8 |
| Growth in total population | 74.3 | 42.0 | 36.4 |
| Growth in white population | 53.7 | 27.6 | 20.6 |
| | | | |
| *20-70% minority:* | 7,051 | 10,360 | 16,024 |
| # in MSAs | 6,887 | 9,966 | 13,236 |
| # in sample | 5,366 | 8,158 | 10,824 |
| Mean family income | $9,689 | $21,180 | $42,509 |
| Unemployment rate | 5.6 | 7.2 | 7.1 |
| % single family homes | 58.6 | 55.1 | 55.4 |
| Growth in total population | 34.8 | 28.6 | 25.0 |
| Growth in white population | 4.8 | 3.2 | -1.2 |
| | | | |
| *70-100% minority:* | 2,860 | 5,574 | 7,642 |
| # in MSAs | 2,853 | 5,492 | 7,155 |
| # in sample | 2,613 | 4,904 | 6,413 |
| Mean family income | $7,748 | $15,642 | $28,914 |
| Unemployment rate | 7.2 | 12.4 | 14.6 |
| % single family homes | 46.0 | 47.1 | 49.8 |
| Growth in total population | -17.4 | -0.6 | 2.9 |
| Growth in white population | -5.7 | -2.5 | -1.9 |

**Table 2: Overview of candidate tipping points**

| | 1970-1980 | 1980-1990 | 1990-2000 |
|---|---|---|---|
| # of MSAs in sample | 104 | 113 | 114 |
| # with identified candidate tipping points | 100 | 104 | 110 |
| | | | |
| Candidate tipping points: | | | |
| Mean | 11.90 | 12.88 | 13.84 |
| SD | 9.47 | 9.04 | 8.68 |
| Correlations | | | |
| 1970-1980 | 1 | | |
| 1980-1990 | 0.52 | 1 | |
| 1990-2000 | 0.60 | 0.63 | 1 |

Note: Tipping points describe the minority share in the census tract, measured in percentage points. Summary statistics are unweighted.

**Table 3. Basic regression discontinuity models for the change in white share around the candidate tipping point**

| | (1) | (2) | (3) |
|---|---|---|---|
| 1970 – 1980 | | | |
| Beyond fixed-point (Yes = 1) | -21.6 | -21.0 | -15.6 |
| | (2.8) | (3.1) | (3.0) |
| Global 4-th order polynomial in deviation from fixed point | X | | |
| Two quadratics in deviation from fixed point (one on each side) | | X | X |
| Baseline demographic / housing characteristic controls | | | X |
| Observations | 35,348 | 35,348 | 35,276 |
| $R^2$ | 0.20 | 0.20 | 0.25 |
| | | | |
| 1980 – 1990 | | | |
| Beyond fixed-point (Yes = 1) | -21.0 | -19.8 | -14.9 |
| | (2.8) | (2.9) | (2.9) |
| Observations | 35,027 | 35,027 | 34,965 |
| $R^2$ | 0.21 | 0.21 | 0.31 |
| | | | |
| 1990 – 2000 | | | |
| Beyond fixed-point (Yes = 1) | -11.0 | -11.2 | -9.9 |
| | (1.5) | (1.5) | (1.4) |
| Observations | 39,593 | 39,593 | 39,511 |
| $R^2$ | 0.14 | 0.14 | 0.16 |

Notes: Dependent variable is the change in white population in the tract over 10 years, expressed as a percentage (0 – 100) of the base-year total tract population. Demographic and housing characteristic controls are the base-year unemployment rate, log(mean family income), housing vacancy rate, renter share, and fraction of homes in single-unit buildings. All specifications include MSA fixed effects, and standard errors are clustered on the MSA.

**Table 4: Basic regression discontinuity models for the change in minority and total population around the candidate tipping point**

| | Change in minority share | | | Change in total population | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| 1970 – 1980 | | | | | | |
| Beyond fixed-point (Yes = 1) | 2.5 | 3.7 | 3.4 | -19.2 | -17.3 | -12.2 |
| | (1.1) | (1.1) | (1.1) | (3.1) | (3.5) | (3.5) |
| $R^2$ | 0.24 | 0.23 | 0.24 | 0.20 | 0.20 | 0.25 |
| Global 4-th order polynomial | X | | | X | | |
| Two quadratics | | X | X | | X | X |
| Demographic/housing characteristics | | | X | | | X |
| 1980 – 1990 | | | | | | |
| Beyond fixed-point (Yes = 1) | -0.6 | -0.3 | -0.2 | -21.6 | -20.1 | -15.1 |
| | (1.0) | (1.0) | (0.9) | (3.5) | (3.4) | (3.5) |
| $R^2$ | 0.28 | 0.28 | 0.30 | 0.23 | 0.23 | 0.30 |
| 1990 – 2000 | | | | | | |
| Beyond fixed-point (Yes = 1) | 1.8 | 1.5 | 2.0 | -9.2 | -9.7 | -7.8 |
| | (0.6) | (0.6) | (0.6) | (1.7) | (1.8) | (1.7) |
| $R^2$ | 0.22 | 0.22 | 0.23 | 0.13 | 0.13 | 0.16 |

Notes: Dependent variable in columns 1-3 is the change in the non-white population in the tract over 10 years, expressed as a percentage (0 – 100) of the base-year total tract population. In columns 4-6, dependent variable is the percentage change in the total tract population over 10 years. All specifications include MSA fixed effects, and standard errors are clustered on the MSA. See notes to Table 3 for "demographic/housing characteristics."

**Table 5. Sensitivity of 1990-2000 regression discontinuity results to additional baseline controls**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| Beyond fixed point (1=yes) | -9.9 | -9.4 | -9.6 | -9.9 | -9.2 | -5.9 | -5.5 |
| | (1.4) | (1.3) | (1.4) | (1.4) | (1.4) | (1.3) | (1.3) |
| Two quadratics in deviation from fixed point (one on each side) | X | X | X | X | X | X | X |
| Tract characteristics from Table 3, column 3 | X | X | X | X | X | X | X |
| 4th order polynomial in: | | | | | | | |
| Poverty rate | | X | | | | | X |
| log(mean family income) | | | X | | | | X |
| Unemployment rate | | | | X | | | X |
| Renter share | | | | | X | | X |
| Distance to nearest "minority" tract | | | | | | X | X |
| $R^2$ | 0.16 | 0.17 | 0.17 | 0.17 | 0.18 | 0.22 | 0.23 |

Notes: Dependent variable is the change in white population in the tract over 10 years, expressed as a percentage (0 – 100) of the base-year total tract population. All specifications include MSA fixed effects, and standard errors are clustered on the MSA.

**Table 6: Models with average minority share in neighboring tracts**

| | 1970-1980 | | 1980-1990 | | 1990-2000 | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| This tract is beyond the fixed point | -10.4 | -6.8 | -11.1 | -8.5 | -7.5 | -6.3 |
| | (1.8) | (1.8) | (2.0) | (2.0) | (1.3) | (1.2) |
| Fraction of neighbor group | -26.9 | -20.7 | -27.5 | -20.0 | -13.7 | -12.0 |
| beyond the fixed-point | (2.7) | (2.6) | (4.1) | (3.6) | (2.5) | (2.3) |
| $4^{th}$-order polynomial in this tract's deviation from fixed-point | X | X | X | X | X | X |
| Average across neighbor group of $4^{th}$-order polynomial in deviation from fixed-point | X | X | X | X | X | X |
| Tract characteristics | | X | | X | | X |
| Observations | 33,841 | 33,772 | 33,017 | 32,955 | 37,513 | 37,420 |
| $R^2$ | 0.21 | 0.26 | 0.23 | 0.33 | 0.10 | 0.12 |

Notes: Dependent variable is the change in the white population in a single tract, expressed as a percentage (0 – 100) of the base-year total tract population. The "neighbor group" is the five closest tracts within four miles, measuring distances from tract centroids. All specifications include MSA fixed effects, and standard errors are clustered on the MSA.

**Table 7: Tipping in racial composition and poverty rate**

| | 1970-1980 | | | 1980-1990 | | | 1990-2000 | | |
|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| Beyond fixed-point for minority share | -15.2 | -13.9 | -13.7 | -14.2 | -14.5 | -13.4 | -7.7 | -7.8 | -5.9 |
| | (2.6) | (3.1) | (3.7) | (2.8) | (3.2) | (3.8) | (1.0) | (1.1) | (1.1) |
| Beyond fixed-point for poverty | -10.8 | -13.4 | -6.7 | -7.9 | -6.5 | -6.3 | -7.5 | -4.9 | -3.5 |
| | (3.7) | (4.1) | (4.0) | (1.6) | (1.9) | (2.7) | (1.0) | (0.8) | (1.0) |
| Global 4-th order polynomials in deviations from fixed-points | X | | | X | | | X | | |
| Quadratics in deviations from fixed-points (one on each side of each fixed point) | | X | | | X | | | X | |
| Cubics in deviations from fixed-points (one on each side of each fixed point) | | | X | | | X | | | X |
| $R^2$ | 0.27 | 0.27 | 0.27 | 0.31 | 0.29 | 0.31 | 0.17 | 0.17 | 0.18 |

Notes: Dependent variable is the change in the white population in a single tract, expressed as a percentage (0 – 100) of the base-year total tract population. All specifications include MSA fixed effects, and standard errors are clustered on the MSA.

**Table 8: Tipping in minority share, black share, and black/Hispanic share**

| | 1970-1980 | | | | 1980-1990 | | | | 1990-2000 | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) | (12) |
| Beyond minority share fixed point | -15.6 | | | -9.5 | -14.9 | | | -9.7 | -9.9 | | | -5.5 |
| | (3.0) | | | (3.2) | (2.9) | | | (3.2) | (1.4) | | | (1.3) |
| Beyond black share fixed point | | -21.7 | | -15.0 | | -14.9 | | -11.4 | | -9.1 | | -3.9 |
| | | (3.0) | | (3.5) | | (3.6) | | (5.0) | | (1.3) | | (1.6) |
| Beyond black/Hispanic share fixed point | | | -9.2 | -18.7 | | | -12.1 | -0.6 | | | -10.3 | -5.2 |
| | | | (30.6) | (29.0) | | | (3.1) | (3.8) | | | (1.3) | (1.2) |
| Two quadratics in minority share deviation from fixed point | X | | | X | X | | | X | X | | | X |
| Two quadratics in black share deviation from fixed point | | X | | X | | X | | X | | X | | X |
| Two quadratics in black/Hisp share dev. from fixed point | | | X | X | | | X | X | | | X | X |
| Tract chars from Table 3, column 3 | X | X | X | X | X | X | X | X | X | X | X | X |
| R2 | 0.25 | 0.25 | 0.24 | 0.26 | 0.31 | 0.30 | 0.31 | 0.31 | 0.16 | 0.16 | 0.16 | 0.17 |

Note: Black, non-Hispanics are not separately identified in 1970 data, so must be imputed. This imputation affects the black/Hispanic share fixed points for 1970-1980. All specifications include MSA fixed effects, and standard errors are clustered on the MSA.

**Table 9: Regression discontinuity models for housing prices before and after tipping**

| | Log(mean value) | | | | Change in log(mean value) | | |
|---|---|---|---|---|---|---|---|
| | 1970 | 1980 | 1990 | 2000 | 1970-90 | 1990-2000 | 1970-2000 |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
| Beyond fixed point in 1990 | -0.02 | -0.06 | -0.07 | -0.08 | -0.04 | -0.01 | -0.05 |
| | (0.01) | (0.02) | (0.02) | (0.02) | (0.01) | (0.01) | (0.01) |
| | Log(mean value) | | | | Change in log(mean value) | | |
| | 1970 | 1980 | 1990 | 2000 | 1970-80 | 1980-2000 | 1970-2000 |
| Beyond fixed point in 1970 | -0.13 | -0.18 | -0.16 | -0.14 | -0.04 | 0.02 | -0.02 |
| | (0.02) | (0.02) | (0.03) | (0.03) | (0.02) | (0.01) | (0.02) |

Note: All specifications include MSA fixed effects, a quadratic in the minority share minus the fixed point (in 1990 in the first row and in 1970 in the second row), and the interaction of this quadratic with the "beyond fixed point" indicator. Standard errors are clustered on the MSA.

**Table 10. School-level tipping between 1990 and 2000**

| | Change in white enrollment, 1990-2000 | | | Change in total enrollment, 1990-2000 | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Beyond fixed point (1=yes) | -9.4 | -8.6 | -8.7 | -8.6 | -8.2 | -9.0 |
| | (0.9) | (0.9) | (0.9) | (1.0) | (1.1) | (1.1) |
| 4th order polynomial in deviation from fixed point | X | | | X | | |
| Two quadratics in deviation from fixed point | | X | X | | X | X |
| Fraction free lunch | | | X | | | X |
| N | 16,046 | 16,046 | 14,130 | 16,046 | 16,046 | 14,130 |
| R2 | 0.09 | 0.09 | 0.11 | 0.06 | 0.06 | 0.06 |

Note: Free lunch variable is unavailable for most schools in 1990. 1995 or 2000 values are assigned when the 1990 value is missing. All specifications include MSA fixed effects, and standard errors are clustered on the MSA.

**Table 11. Relation of residential tipping points with attitudes**
Dependent variable is the average of the 1980-1990 and 1990-2000 tipping points (in percentage points)

| | Sample means | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|---|
| Race Attitudes index | 0.05 | -0.85 | -1.62 | -2.62 | -5.91 | -4.70 | -6.12 | -5.31 |
| | [0.72] | (1.45) | (1.20) | (1.14) | (1.46) | (1.58) | (1.51) | (1.50) |
| % Black | 13.6 | | 0.43 | 0.37 | 0.57 | 0.61 | 0.58 | 0.58 |
| | [9.65] | | (0.09) | (0.08) | (0.09) | (0.09) | (0.10) | (0.10) |
| % Hispanic | 6.3 | | 0.43 | 0.30 | 0.42 | 0.58 | 0.43 | 0.46 |
| | [6.29] | | (0.06) | (0.07) | (0.10) | (0.12) | (0.10) | (0.10) |
| Log(per capita iIncome), whites | 10.0<br>[0.14] | | | | -15.58<br>(6.52) | -22.51<br>(7.99) | -12.38<br>(7.34) | -13.55<br>(6.62) |
| Log(per capita income), blacks | 9.45<br>[0.14] | | | | -6.38<br>(7.83) | -8.46<br>(9.30) | -10.40<br>(8.93) | -6.24<br>(7.82) |
| Log(per capita income), Hispanics | 9.54<br>[0.17] | | | | 1.81<br>(4.94) | 5.73<br>(5.08) | 2.72<br>(5.07) | 2.38<br>(4.91) |
| % w/ < HS, blacks | 33.0 | | | | | -0.07 | | |
| | [7.2] | | | | | (0.16) | | |
| % w/ < HS, whites | 19.5 | | | | | -0.38 | | |
| | [4.9] | | | | | (0.22) | | |
| % 5-12 year olds in private school | 13.7<br>[4.9] | | | | | | -0.13<br>(0.14) | |
| Herfindahl index over school districts | 0.21<br>[0.21] | | | | | | -0.82<br>(3.78) | |
| Population density | 114 | | | | | | | -0.005 |
| | [163] | | | | | | | (0.005) |
| % 1990 houses built in the 1980's | 21.1<br>[9.2] | | | | | | | 0.10<br>(0.09) |
| Region dummies (4) | | | | X | X | X | X | X |
| $R^2$ | | 0.006 | 0.54 | 0.69 | 0.76 | 0.79 | 0.77 | 0.78 |

Note: N=62. Sample is weighted by MSA population in 1990. All explanatory variables are measured at the MSA level in 1990. The racism index is derived from responses to the General Social Survey; see text for details.

**Appendix Table 1. Discontinuities in 1990 neighborhood demographic characteristics at 1990-2000 potential tipping point**

| | Poverty rate | log(mean family income) | Unemployment rate | Renter share | Distance to nearest "minority" tract (miles) |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| Beyond fixed point (1=yes) | 0.029<br>(0.004) | -0.063<br>(0.015) | 0.011<br>(0.002) | 0.082<br>(0.008) | -1.32<br>(0.24) |
| N | 39,557 | 39,511 | 39,560 | 39,534 | 39,593 |
| R2 | 0.54 | 0.49 | 0.54 | 0.37 | 0.26 |

Notes: Each specification includes MSA fixed effects, a quadratic in the deviation of the tract minority share from the city-level fixed point, and an interaction of each of the terms of this quadratic with an indicator for whether the tract is beyond the fixed point. Only the main effect on this indicator is reported. Standard errors are clustered on the MSA.

**Appendix Table 2: Correlations of MSA average responses across GSS questions about racial attitudes**

| | Individual mean | MSA-level correlation with question: I. | II. | III. | IV. |
|---|---|---|---|---|---|
| I. | 0.14 | 1.00 | | | |
| II. | 0.71 | 0.26 | 1.00 | | |
| III. | 0.17 | 0.52 | 0.36 | 1.00 | |
| IV. | 0.37 | 0.51 | 0.28 | 0.46 | 1.00 |
| Avg. | n/a | 0.77 | 0.64 | 0.79 | 0.76 |

Note: Questions are:

I: Do you think there should be laws against marriages between blacks and whites?
*[Coded as “Yes”=1, “No”=0, “Don’t know”=missing.]*
II: In general, do you favor or oppose the busing of black and white school children from one school district to another?
*[“Oppose”=1, “Favor”=0.]*
III. How strongly do you agree or disagree with the statement: “White people have a right to keep blacks out of their neighborhoods if they want to, and blacks should respect that right”? (1 = “agree strongly” or “agree slightly”)
*[“Agree strongly” and “agree slightly” = 1, “disagree strongly and “disagree slightly” = 0.]*
IV. Suppose there is a community wide vote on the general housing issue. Which (of the following two) laws would you vote for:
A. One law says that a homeowner can decide for himself whom to sell his house to, even if he prefers not to sell to blacks.
B. The second law says that a homeowner cannot refuse to sell to someone because of their race or color.
*[A = 1, B=0.]*

## Appendix Table 3. City values of the racism index

| MSA | Racism index value | MSA | Racism index value |
|---|---|---|---|
| Worcester, MA-CT | -1.11 | Norfolk-Virginia Beach-Newport News, VA-NC | 0.05 |
| Rochester, NY | -1.08 | Columbus, OH | 0.10 |
| San Diego, CA | -1.07 | Richmond-Petersburg, VA | 0.11 |
| Tucson, AZ | -0.98 | Lansing-East Lansing, MI | 0.11 |
| San Francisco, CA | -0.87 | Fort Lauderdale, FL | 0.15 |
| Minneapolis-St., Paul, MN-WI | -0.83 | Baltimore, MD | 0.19 |
| Phoenix-Mesa, AZ | -0.80 | Cincinnati, OH-KY-IN | 0.22 |
| Denver, CO | -0.71 | Kansas City, MO-KS | 0.23 |
| Washington, DC-MD-VA-WV | -0.69 | St. Louis, MO-IL | 0.23 |
| Allentown-Bethlehem-Easton, PA | -0.68 | Pittsburgh, PA | 0.30 |
| Philadelphia, PA-NJ | -0.68 | Chicago, IL | 0.36 |
| Seattle-Bellevue-Everett, WA | -0.67 | Tampa-St. Petersburg-Clearwater, FL | 0.38 |
| Portland-Vancouver, OR-WA | -0.64 | Dallas, TX | 0.44 |
| Grand Rapids-Muskegon-Holland, MI | -0.61 | Buffalo-Niagara Falls, NY | 0.45 |
| Boston, MA-NH | -0.60 | Detroit, MI | 0.50 |
| Des Moines, IA | -0.58 | Dayton-Springfield, OH | 0.50 |
| Riverside-San Bernardino, CA | -0.56 | Indianapolis, IN | 0.59 |
| Syracuse, NY | -0.53 | Atlanta, GA | 0.65 |
| Tacoma, WA | -0.50 | Oklahoma City, OK | 0.66 |
| Saginaw-Bay City-Midland, MI | -0.45 | Nashville, TN | 0.80 |
| New Haven-Meriden, CT | -0.42 | Lakeland-Winter Haven, FL | 0.84 |
| Providence-Fall River-Warwick, RI-MA | -0.41 | West Palm Beach-Boca Raton, FL | 0.89 |
| Fresno, CA | -0.41 | Jackson, MS | 0.89 |
| Newark, NJ | -0.34 | Jacksonville, FL | 0.95 |
| Los Angeles-Long Beach, CA | -0.26 | Charleston-North Charleston, SC | 0.99 |
| Houston, TX | -0.19 | Charlotte-Gastonia-Rock Hill, NC-SC | 1.25 |
| Harrisburg-Lebanon-Carlisle, PA | -0.14 | New Orleans, LA | 1.32 |
| Fort Wayne, IN | -0.14 | Johnson City-Kingsport-Bristol, TN-VA | 1.38 |
| Cleveland-Lorain-Elyria, OH | -0.11 | Birmingham, AL | 1.43 |
| Milwaukee-Waukesha, WI | -0.02 | Knoxville, TN | 1.53 |
| Austin-San Marcos, TX | 0.00 | Memphis, TN-AR-MS | 1.65 |