Racial Segregation in Interwar United States: A Dynamic Segregation Approach

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Abstract

Between 1910 and 1950, more than 3.5 million African Americans migrated from the south, largely to northern, urban areas (Collins 1997). Yet when they arrived, they found themselves often limited in their choice of neighborhoods via racially restrictive covenants (Brooks 2011). This paper follows the dynamic segregation literature of Schelling (1971) and Card, Mas, Rothstein (2008) to explore whether neighborhoods in interwar cities in the United States demonstrated "tipping behavior" and how these tipping points evolved over time. Using census-tract data from both the 1930 and 1940 U.S. Census as well as the 1934 Real Property Inventory, our results suggest that tipping behaviour did occur although the tipping points are typically lower than those found in the modern era. Unlike the modern era in which the white population fell around 12 percent, our results suggest that growth of white households in neighborhoods substantially slowed compared to the growth of the white population in the cities as a whole.

Keywords: Racial Segregation, Tipping, Structural Break

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1. Introduction

Racial segregation across neighborhoods is a salient characteristic of urban areas in the United States. Despite the Constitution of the United States and federal law declaring that everyone was equal, racial segregation in housing has not only persisted in many cities, it has become more extreme over the last century (Massey, 2001). The interwar period in the United States is associated with a time when racial segregation takes on a different dynamic. Blacks in America who were living in difficult conditions during the Great Depression were drawn to the north, mainly motivated by the benefits of the New Deal programs that were distributed more effectively in the north than in the south. The availability of jobs with the arrival of World War II (1939-45) also provided an impetus that witnessed millions of blacks moved from the rural south to the north (Wright, 1986). It was estimated that between 1910 and 1950, more than 3.5 million African Americans migrated from the south, largely to northern, urban areas (Collins, 1997). Yet when they arrived, African Americans found themselves often limited in their choice of neighborhoods via racially restrictive covenants (Brooks, 2011).

This paper examines the dynamic of racial segregation that took place in seven cities of the United States between 1930 and 1940. There are two schools of thought that rationalize racial segregation. The classic models of Tiebout (1956) and Rosen (1974) attribute urban segregation to households' differences in incomes and preferences, which determine their willingness to pay for location characteristics. Given that black people had restricted access to education, this tended to lead black households towards those neighborhoods with lower levels of amenities commensurate with their income.

On the other hand, the models of Schelling (1971), Becker and Murphy (2000) attribute racial segregation to households' concerns about the demography of their neighbors. The seminal work of Card, Mas, and Rothstein (2008), which examines the process by which a neighborhood can polarize towards complete segregation – tipping, attributes segregation in urban neighborhoods to white's distaste for residing near minorities. Still there are models that attempt to explain racial segregation as an outcome of interaction between an exogenous location characteristic and demographic segregation (see Banzhaf and Walsh, 2010).[1] Yet this belief that racial segregation is born out of white's distaste for residing near minorities clearly need not hold only for the modern period. Motivated by the historical background of racial segregation in interwar period, we test whether tipping was present in some U.S. cities, and if so, did it differ from results based on modern period.

Following the regression discontinuity approach of Card, Mas, and Rothstein (2008), we test for tipping points in the United States for seven cities and estimate the "tipping point" at which previously predominantly white neighborhoods see large increases in the share of minority households. Our results are then compared to that of Card et al. (2008) to inform the change in the dynamic of tipping that occurred in the interwar years. Our preliminary results, using a structural-break procedure, suggest that tipping in cities occurred and appear to have been lower than those found in the modern era with the exception of Washington, DC and Chicago.

These preliminary results suggest that white households were more sensitive to the arrival of blacks than their modern counterparts. This is consistent with the notion that white households were actively trying to prevent blacks from moving into the neighborhoods via restrictive covenants. Yet our results suggest that once a neighborhood "tipped" whites did not leave the neighborhoods, but rather the growth in the neighborhood largely stopped. We speculate that this is largely due to the economic climate of the period combined with the lack of suitable white neighborhoods that became available through the combination of mortgage insurance and improved transport networks.

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[1] Banzhaf and Walsh (2010) develop a general equilibrium model that captures the behavior of households when choosing the neighborhood they want to live in based on its endogenous demographics and its exogenous public good. Several interesting findings emerge from the model. When sorting arises from tastes for the exogenous public good rather than demographic tastes, some racial segregation can occur with richer households benefiting from higher levels of the public good. However, when tastes for endogenous demographic composition are incorporated in the model, further segregation occurs consistent with the prediction of Schelling's "tipping model". More importantly, policy that improves the public good in a low-quality but high minority neighborhood may lead to an increase in group segregation, as richer minorities move into the neighborhood due to the improvement in the public good. In neighborhoods where differences in public goods are less important, sorting is dominated by tastes of demographic preferences over income-based sorting on the public good.

On the methodology front, we (will) contribute to the literature by employing a robust method of inference when dealing with the problem of small sample and inference for threshold models. We also test the robustness of our estimated tipping points using an alternative technique and analyze whether the estimated tipping points are associated with any discontinuities in housing prices and contract rents as suggested is apparent by Hoyt (1939).

2. Model for Tipping

Our model follows the methodology developed first by Schelling (1971) and fleshed out by in Becker and Murphy (2000) as well as Card, Mas, and Rothstein (2008). In these models, households segregate as their utility functions are directly dependent on the racial or ethnic composition of their neighborhood. Yet as discussed in Tiebout (1956) and Rosen (1974), segregation may merely be a product of sorting on exogenous location amenities, preferences for which are highly dependent on racial or ethnic characteristics.

Both Banzhaf and Walsh (2010) and Kasy (2015) indicate strategies to econometrically identify these separate effects, yet current data limitations for the interwar period for the United States does not allow us to make such a distinction. Instead, we rely on narrative evidence of the period to guide the modeling.

An economist for the Federal Housing Administration, Homer Hoyt wrote in 1939:

It is a mere truism to enunciate that colored people tend to live in segregated districts of American cities. As we have said [earlier], the reflection of adverse housing characteristics should tend to operate in the same manner in areas populated entirely by colored races as in areas populated only by whites. It is in the twilight zone, where members of different races live together that racial mixtures tend to have a depressing effect upon land values -- and therefore, upon rents.

Moreover, the existence of racial covenants in housing deeds restricting various racial, ethnic, and religious households from purchasing housing during the period is again suggestive that the white majority had a preference for living with other white households. The enforcement of these covenants would have thus been unnecessary in a model in which blacks self-segregated into neighborhoods with other blacks based on the existing amenities.

Figure 1 provides further evidence of the diffusion of blacks into the well-known black enclaves, Harlem and Bedford-Stuyvesant, between 1910 and 1940. This figure shows a marked spatial diffusion process in which black households are moving into neighborhoods in which either was largely black or in which the surrounding neighborhoods already were.

The historical narrative of Harlem suggests that it began life as a black enclave in the early 20th century after the advent of a local housing bubble left housing prices at below market. This allowed several black churches from lower Manhattan to begin to purchase properties and either sell or rent out to black households (Kollmann 2012). While this is suggestive that exogenous housing characteristics must be accounted for, the likelihood that the only neighborhoods experiencing intra-city falls in market prices being the only neighborhoods experiencing black in-migration is unlikely.

The premise is that whites have an aversion to residing with minorities and their bid-rent curve for housing in a neighborhood will begin to decline past a certain threshold. At the point where white's bid-rent curve falls below that of African Americans, the neighborhood effectively "tips" and thus should become predominately black. As suggested in Figure 1, Harlem and Bedford-Stuyvesant appear to be candidates that illustrate this tipping behavior. This *prima facie* evidence suggests that "tipping" could well exist in some cities, including that of New York, during the interwar period.

While a more comprehensive model will be developed in a later draft (suggestions welcome), we begin with a partial equilibrium model in which we consider two groups, whites

(w) and blacks (m).[2] We let $b^g(n^g, m)$ be the inverse demand function for group g in the neighborhood which as a share of blacks equaling m and n^g is the number of families in the group willing to pay at least b^g to live there. By definition, the bid-rent functions are constructed such that $\partial b^g/\partial n^g \leq 0$. Likewise, we assume $\partial b^w/\partial m$ represents the social interaction effects of the share of the minorities on the bid-rent function of whites. If the social interaction term becomes increasingly negative as the share of minorities increases, we will see the white bid-rent function initially increase, but eventually fall.

Figure 2, which is taken from Card, Mas, and Rothstein (2008), illustrates the idea of tipping behavior in the face of an increased demand for housing by blacks. As this demand increases, the stable equilibrium for the share of black in a neighborhood increases up to a certain threshold. After that point, we see the neighborhood should tip and become predominately black. The model is simple in that it assumes no change in demand for white households and neither does it suggest on how quickly this dynamic should shift. Yet as suggested in Figure 1, the dynamic process of a neighborhood tipping appears to take a decade or two to complete, likely a result of limitations in household mobility.

3. Empirical Methodology

4.1 The tipping model of Card, Mas and Rothstein (2008)

The tipping model of Card, Mas and Rothstein (2008) is developed for a local housing market in which white's willingness to pay for homes is a function of the minority share in the neighborhood. The minority share of the neighborhood will vary according to changes in the relative demand of whites and minorities but the variation is expected to be smooth as long as the minority share remains below a critical threshold level. However, when the minority share exceeds the threshold, all white households will leave. This abrupt change in the dynamic of white share beyond a certain threshold is a salient feature of tipping. The location of the

[2] While the data from the period does not always disaggregate between non-white minorities, non-whites in our selected cities are largely black during the time period.

threshold (i.e. tipping point) is determined by the degree of whites' tolerance concerning minority contact; the higher the tolerance the higher is the threshold.

To test for this tipping phenomenon, we use inter-period (i.e. 1930-1934, 1934-1940 and 1930-1940) changes in neighborhood racial composition. Given that the tipping point is unobserved, we use two methods to identify city-specific possible tipping points. The first method relies on structural break tests and chooses the break point associated with the best-fitting model for tract-level white population changes. The second method uses a nonparametric method that is considered to be more flexible and does not make specific assumptions about the functional form for tract-level changes in white population shares in each city. However, unlike the methods employed by Card, Mas and Rothstein (2008), we undertake further analysis by utilizing different weights and applying them to the data. The basis of this approach is that the seven cities we examined vary significantly in terms of the size of the population. To the extent that the dynamic of tipping may be contingent on the size of the population, we explore both the tipping point without weighting a census tract as well as controlling for the population in the base year to underweight tracts with small populations that may be driving the results.[3]

3.2 Empirical model

Given the period examined is associated with a large influx of black moving from the rural south to the north, we accommodate changes in the population of a neighborhood by expressing changes in the numbers of white and black residents as a fraction of the base year population. The base year is either 1930 or 1934. Define $W_{ic,t}$ as the number of whites, $M_{ic,t}$ the number of minorities and $P_{ic,t}$ ($=W_{ic,t}+M_{ic,t}$) the number of whites, minorities, and total residents of neighborhood *i* in city *c* in year *t* (=1930, 1934, 1940). The dependent variable is the change in the neighborhood's white population, taken as a share of the initial population in the base year, $Dw_{ic,t} = (W_{ic,t} - W_{ic,\text{base year}})/P_{ic,\text{base year}}$. To establish the dynamic of tipping, the dependent variable is specified as

[3] An alternative approach to weight the census tracts by the inverse z-score of the population was also employed to underweight both the smallest and largest census tracts. However, these results reflect very similar results to when the census tract are unweighted and are thus omitted from further discussion.

$$Dw_{ic,t} = p(\delta_{ic,base\ year}) + d\mathbf{1}[\delta_{ic,base\ year} > 0] + \tau_c + \boldsymbol{X}_{ic,base\ year}\,\beta + \varepsilon_{ic,t}, \qquad (1)$$

where $\delta_{ic,base\ year} = m_{ic,base\ year} - m^*_{ic,base\ year}$, such that $m_{ic,base\ year} = M_{ic,base\ year}/P_{ic,base\ year}$. Here, $m^*_{ic,base\ year}$ is the tipping point or threshold. Note that τ_c is a city fixed effect and $\boldsymbol{X}_{ic,base\ year}$ is a vector of neighborhood control variables. Depending on the availability of data for the neighborhood control variables in different periods, we include the homeownership rate in base year, share of multiple dwellings in base year and the median rent in base year in the specification. $p(\delta_{ic,base\ year})$ is a smooth control fourth-order polynomial function. This specification is estimated for the following sample period: 1930-34, 1934-1940 and 1930-1940.

Before estimating equation (1), it is necessary to estimate $m^*_{ic,base\ year}$ from the data. This requires the assumption that there exists a tipping point for which $d \neq 0$. In establishing the location of the tipping point, we employ the structural break test which involves searching over the time series data of $m_{ic,t}$ for a break point satisfying certain condition. Using a simplified version of equation (1) which ignores the covariates and replacing the polynomial function p() with a constant, we estimate

$$Dw_{ic,t} = \alpha_c + d_c\mathbf{1}[m_{ic,base\ year} > m^*_{ic,base\ year}] + \varepsilon_{ic,t}, \qquad (2)$$

for $0 \leq m_{ic,base\ year} \leq M$ where M is set to 60% and the value of $m^*_{ic,base\ year}$ is determined in the [0,50%] interval based on the condition that the R^2 of (2) is maximized for each city and each period. A consistent estimate of the threshold can be obtained following this procedure as long as equation (2) is correctly specified (Hansen, 2000).

The structural break method for determining tipping point can be unreliable when applied to small cities and it can be heavily influenced by outliers. The second method utilises the approximation of the smoothed polynomial function for $E(Dw_{ic,t}| c, m_{ic,base\ year})$ for many different cities. To ensure that our results are not biased by city-wide trends of rising minority shares, we subtract $E(Dw_{ic,t}| c)$ from $E(Dw_{ic,t}| c, m_{ic,base\ year})$. Tipping in the second approach is defined by:

$$E(Dw_{ic,t}| c, m_{ic,base\ year}=m^*-\varepsilon)>E(Dw_{ic,t}| c)> E(Dw_{ic,t}| c, m_{ic,base\ year}=m^*+\varepsilon) \text{ for } \varepsilon>0. \qquad (3)$$

Here, there is a "fixed point" for the city-specific tipping point, that is at m*. This is also the level of minority share at which the neighborhood white population grows at the average rate for the city. In identifying this fixed point, we need to smooth the data to obtain a continuous approximation, $R(m_{base\ year})$, to $E(Dw_{ic,t}| c, m_{ic,base\ year})- E(Dw_{ic,t}| c)$. We then choose the root to this function. The steps involve first fitting $Dw_{ic,t} - E(Dw_{ic,t}| c)$ to a quartic polynomial in $m_{ic,base\ year}$ for neighborhoods with $m_{ic,base\ year}<60\%$ which yields $R(m_{base\ year})$. Using a root of this polynomial, m', we exclude all neighborhoods with $abs(m_{ic,base\ year}-m')>10$ before fitting a second quartic polynomial to the remaining sample. The "fixed point", m*, is the root of this second polynomial. For the purpose of consistency with the first method, we only consider minority shares below 50% as fixed points. When multiple roots are present in this range, we choose the one at which the slope of R(m) is most negative.

Our results suggest that the use of weights in determining the threshold is important for some cities. There is significant variation in the threshold with and without the use of weights. Weighting the data by population size gives rise to an increase (decrease) in the threshold such as Washington DC (Boston) and New York (Louisville).

We undertake further robustness analysis by focusing on tract samples which have the share of non-white lesser than 60%. The intuition is that city specific analysis involving tracts with a high proportion of non-white is prone to overestimate the threshold. Our results indicate that, by and large, the tipping point is robust to the inclusion of tracts that contain a high share of non-white (i.e. greater than 60%). These results are not reported here for brevity, but are available from the authors upon request.

3.3 Statistical Inference

Inference under the null of no discontinuity (i.e. d=0) is not straight forward in the threshold regressions (1) and (2). At the point of structural break, the estimate of d has a non-standard distribution. This arises from the specification search bias (Leamer, 1978), that is conventional test statistics have a tendency to reject the null hypothesis of d=0 given that the same data are used both for the identification of the location of a structural break and for estimating the magnitude of the break. Inference in threshold models usually relies on simulating the distribution of the $\hat{d}$ estimate under the null and the test statistic is compared to the simulated critical values at the appropriate significance level to determine whether the null fails to be rejected (Andrews, 1993 and Hansen, 2000). This method is cumbersome. Card et al. (2008) use a different approach for statistical inference of the null of no discontinuity following the method of Angrist, Imbens and Krueger (1999). They use a randomly selected subset of their sample for the search of a structural break point followed by the use of the remaining subsample for other analyses. Given that the two subsamples are independent, estimates of $\hat{d}$ from the second sample have a standard distribution under the null hypothesis that permits the use of conventional tests. Our smaller sample does not permit the adoption of Card, Mas and Rothstein (2008) split-sample procedure for inference.

For the purpose of inference we employ the method proposed by Gonzalo and Pitarakis (2002) which does not require any simulations but view the problem as one of model selection. The problem of detecting the presence of threshold effects is perceived as a model selection problem among two competing models given by the linear specification

$$Dw_{ic,t} = \tau_c + \boldsymbol{X}_{ic,\text{base year}}\beta + \varepsilon_{ic,t,} \qquad (4)$$

versus its threshold counterpart given by equation (1). The decision rule is based on the theoretic criterion $IC_T(\gamma) = lnS_T(\gamma) + \frac{\lambda_T}{T}(m)$ where λ_T is a deterministic function of the sample size or a constant independent of the sample size that is in turn multiplied by the number of free parameters. Here, S_T is the residual sum of squares. Intuitively, an increase in m arising from the threshold nonlinearity will lead to a reduction in $S_T(\gamma)$ but this reduction is penalized by the $\frac{\lambda_T}{T}(m)$ term due to the resulting increase in the number of estimated parameters. The optimal model is then selected as the model that leads to the smallest value of the IC criterion. In other words, the linear specification (4) is preferred if $IC_T < \min_{\gamma \in T} IC_T(\gamma)$, and opt for the threshold model otherwise. In their Monte Carlo experiments, Gonzalo and Pitarakis (2002) demonstrate that amongst the three types of widely used model selection criteria, namely AIC, HQ and the BIC, the best performance is displayed by the BIC in that it does not lead to spurious parsimonious choices even for finite sample sizes. For purpose of inference, we use the BIC to determine the adequacy of the threshold model.

4. Data and Descriptive Statistics

The primary data comes from various 1934 Real Property Inventories for seven cities: Boston, Chicago, Cleveland, Louisville, New York, Philadelphia, and Washington, DC.[4] This dataset contains census-tract level information detailing the condition of residential structures, racial and demographic distribution of the population, contract rents, and property values for a subset of cities. This data is then matched to the census-tract level data from both the 1930 and 1940 United States Census available from the National Historical Geographic Information System (Minnesota Population Center, 2011).[5] Descriptions of the variables available are found in Table 1.

[4] Data from Chicago was obtained via the NHGIS as it was part of a special Census taken in 1934.
[5] The current draft includes cities in 1930 that were found in 1934. Future drafts will explore cities in the 1930 U.S. Census that did not have an equivalent Real Property Inventory aggregated at the tract-level in 1934.

Table 2 describes the racial composition of black and white households in 1934 by city. To minimize the distortion of the small tracts on the outskirts of the city, we have weighted the means by the tract-level population in 1934.

We see substantial differences in the racial composition across cities, Louisville and Washington, DC having the largest average shares of blacks at 15.85 and 25.07 percent in a census tract, respectively. Interestingly, these two cities had the largest share of growth in white households as a share of the total population in 1934. The growth in Washington, DC is obvious due to the substantial expansion of the federal government via the creation of various New Deal programs during the Great Depression.

Blacks were also moving into the cities in our sample, but much more modestly than whites at this time. New York is an interesting case. It had established black enclaves during the period, yet the overall black population was small. However, Table 2 suggests strong growth, tracts on average saw 2.5 percent growth of blacks from 1934-40 as a share of the total 1934 population. Washington, DC saw the strongest growth among the sample of cities. Similar to the growth of the white population in DC, it is likely a reflection of the strong employment growth of the government sector around the nation's capital.

Yet before we explore the dynamic tipping model discussed above, it is a useful exercise to explore whether we see evidence of tipping in the summary statistics. Table 3 breaks down the growth in the white population from 1934-40 as a share of the total population in 1934 by the distribution of share of blacks in a tract in 1934. The summary statistics are again weighted by 1934 population in order to minimize the effects of small tracts distorting the means.

What we see from the summary statistics is evidence that as the share of blacks in a census tract increases, there is a clear trend towards a reduction in the growth rate of whites from 1934 to 1940. Moreover, the growth rate becomes negative once a census tract has become at least 20 percent black. While the results are not as stark as those in the 1970s through 1990s presented in Card, Mas, and Rothstein (2008), they reflect evidence that there are disincentives for whites living in increasingly black neighborhoods.

5. Empirical Results

5.1 Tipping Point Estimates

The preliminary results for the estimates of the tipping points for each city using the structural break method are found in Figure 3.[6] This figure plots several scatter plots where we plot the share of the black population from 1934 on the x-axis with the percent change in the white population from 1934 through 1940 on the y-axis. The red line indicates the estimated tipping point constructed from the full sample of census tracts as found in Table 4. As we can see from the results, the tipping points for Louisville, Boston, New York, and Philadelphia are under four percent. In particular, the estimates for Philadelphia at 0.1 percent suggest that the preferences for segregation are extreme as compared to the other cities in the sample.

On the other end of the spectrum we have Washington, DC and Chicago with predicted tipping points of 10.1 percent and 23.3 percent respectively. In the case of Washington, DC, this should be expected given that the physical composition of the city during the period resulted in many African Americans living in the alleyways behind housing that was reserved for white families. However, the results for Chicago are striking given that Chicago more recently has had severe issues of segregation in the southern portion of the city.

Yet these estimates of tipping points appear to be sensitive to the specification for Boston and Chicago. If we again estimate the tipping points, but restrict the sample to only tracts with a share of less than 60 percent black, the estimated tipping point for Boston in the 1934-40 period (Table 4) rises from 0.6 percent to 50 percent. Chicago estimates fall from 23.9 percent to 0.2 percent. Yet the estimates for the other cities as well as the other time frames found in Table 3 and Table 5 remain largely consistent.

The remaining estimates for the cities have been estimated to be quite small and in the case of restricted specification, under one percent for Chicago and Philadelphia. The results from the tipping points however are smaller than those that are found in Card, Mas, and

[6] Estimates using the fixed point methodology are currently not included in this draft of the paper.

Rothstein (2008) which typically found estimates of tipping points to be between 5 and 20 percent. While it is plausible and likely that racial attitudes have improved from the 1930s through to the 1990s, the methodology currently employed in this paper as well as CMR suggest that households in a census tract are making decisions to remain or move into a neighborhood completely independent of what is occurring in the surrounding census tracts. Given that a census tract is at best an imperfect definition of a neighborhood, it is plausible that white households begin moving out of a tract once surrounding neighborhoods reach a particular threshold. This could lead to an underestimate of the true tipping point. This will be discussed briefly in the next section which will describe our intended approach to this potential issue.

5.2 Model Estimates

Table 6 presents the estimation results of Equation 1 in which we regress the change in the share of the white population as a share of the base year population on an indicator variable describing whether a tract is above the estimated tipping point, a quartic polynomial of the difference between the share of blacks in a tract and the tipping point, a set of neighborhood control variables and city fixed effects.

Our main coefficient of interest is “Beyond Tipping Point”. As we can see across the two different weighting schemes, the models suggest substantially different effects of a tract being beyond the estimated tipping point. For example, from 1930-34, the unweighted model suggests that tracts above the tipping point experience a fall of 38.6 percent of the white population as a share of the 1930 total population. Yet if we weight by the 1930 population, this decrease is estimated to be only 6.92 percent. It becomes quickly evident when viewing the unweighted summary statistics that these results are being driven by small tracts on the outskirts of the cities in the sample, primarily New York. Yet it would be incorrect to omit these tracts as they suggest an interesting dynamic between tracts on the outskirts of the city and those within the heart. These results suggest that white households living in underdeveloped portions of the city are much more likely to leave a tract in the presence of blacks than those elsewhere.

Figure 4 plots the coefficient estimates from the "beyond the candidate tipping point" as well as the quartic polynomial coefficients in 1930-40 specification found in Table 6. The x-axis is the share of blacks in a tract while the y-axis plots the percent change in the white population from 1930 to 1940. The figure is shown assuming a hypothetical tipping point of 10 percent.

The results are again interesting as they show the substantially different response found in the unweighted and weighted coefficient estimates. We see in the unweighted results, a large increase in the white population before the tipping point, yet a large decline after the threshold. This is contrasted to the results where we weight for the population. This suggests that while the white population growth shows a decline after the candidate tipping point, whites are still moving into these neighborhoods, albeit at a reduced rate.

6. Discussion

6.1 Inclusion of Spatial Dimension

As discussed above, the framework employed currently both in this paper as well as Card, Mas, and Rothstein (2008) fails to incorporate that households make choose to move in or out of a census tract based not only on the composition of the tract, but those surrounding it. It is clear from Figure 1 that there is an element of spatial diffusion occurring as we can visually see blacks moving into census tracts near those that are already primarily black. Failure to capture this effect is likely underestimating the tipping points. While this is not suggesting that we suspect white households to be more tolerant, it is more of a reflection that a census tract is not fully capturing the definition of a neighborhood.

Our proposal is to re-estimate the tipping points of census tracts based on a variation of the structural break methodology using a Spatial Durbin model:

$$D_{w_{ic,t}} = \alpha_c + W_c D_{w_{c,t}} + d_c 1\left[m_{ic,base} > m^*_{ic,base}\right] + \delta_c W_c 1\left[m_{c,base} > m^*_{c,base}\right] + \varepsilon_{ic,t} \quad (5)$$

Where W_c is a $n \times n$ row-normalized spatial weighting matrix for city c. The choice of the weighting matrix is typically a data driven process in which we choose a specification that maximizes a set of Moran's I test statistics. The procedure has not yet been implemented, but should take into account both the simultaneous decisions that household's will move both on surrounding characteristics as well as consistently estimating the tipping point threshold for the city.

6.2 Concluding Remarks

The preliminary results suggest that while there is evidence of tipping occurring during the interwar period in the United States, the dynamic is different to what had typically occurred post-war. With the exception of smaller census tracts in which whites appear to have left as black households entered, the estimates suggest that whites enter a neighborhood only at a reduced rate from the city at-large.

While it is difficult to quantitatively identify the cause from the data, it is not particularly difficult to envisage scenarios given the earlier quote by Homer Hoyt in which the Federal Housing Administration may have crafted policies such that promoted a segregationist housing policy. Thus it should be of no surprise as Jackson (1985, p. 207) notes that "the [FHA] allowed personal and agency bias in favor of all-white subdivisions in the suburbs to affect the kinds of loans it guaranteed." Yet these policies, along with the construction of the federal highway system largely occurred after World War 2.

We are thus left to tentatively conclude that white households continued to reside in mixed racial neighborhoods out of economic necessity. The opportunities for white households to find suitable housing in predominately or exclusively white neighborhoods was potentially a phenomenon that became possible after the development of mortgage insurance via the FHA and the development of transport networks that could allow households to travel between suburbs and the urban centers of employment. This development would ultimately lead to a change in the bid-rent curve of white households, leading a shift in the dynamic of urban segregation throughout the 20th century.

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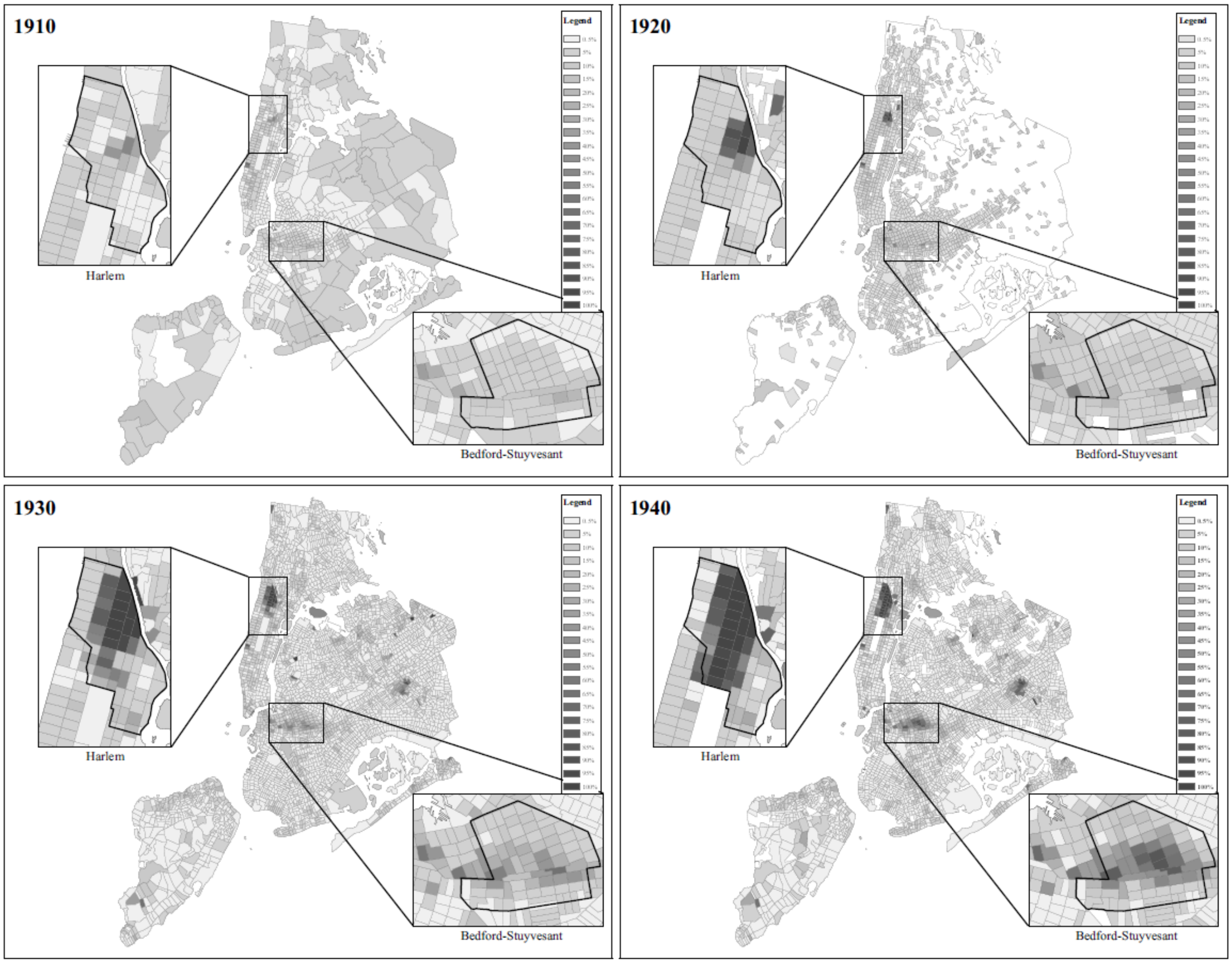
Figure 1 Share of Non-White Population in New York by Census Tract



Note: The darker shaded areas refer to the share black in a census tract.
Source: NHGIS

Figure 2 Rising Minority Demand Leads to a Tipping Point

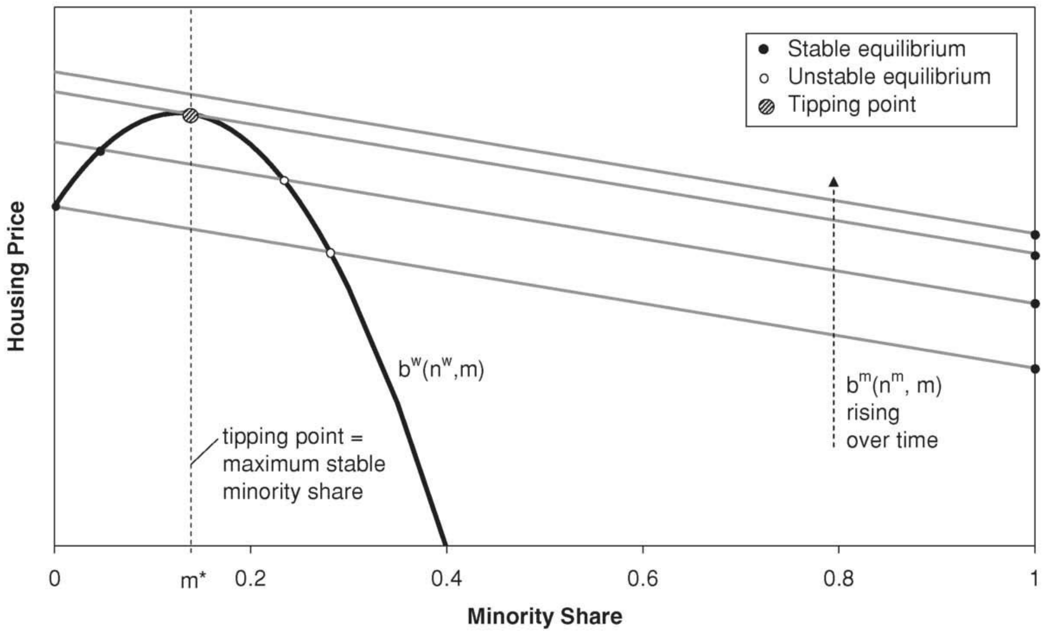


FIGURE III
Rising Minority Demand Leads to a Tipping Point

Source: Card, Mas and Rothstein (2008)

Figure 3 Scatterplot of black population versus the percent change in white population from 1934-40 by city.

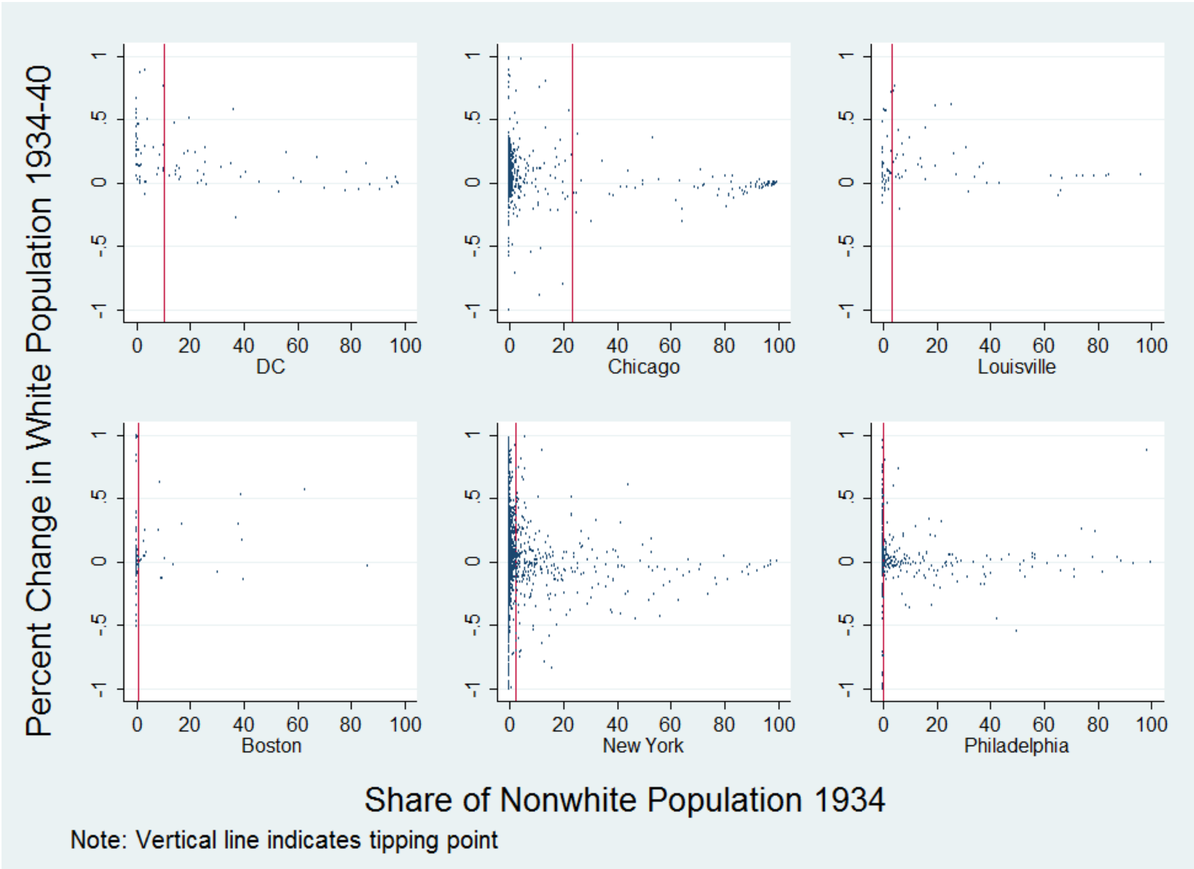


Figure 4

Note: Graph constructed from coefficient estimates found in Table 6 from the 1930-40 specifications. This graph assumes a hypothetical tipping point of 10 percent.

Table 1 Variable Definitions

Variable	**Definition**
Share Black	In 1934, Share of families that are non-white in a census tract. Family generally follows the 1930 US Census definition: "a group of persons, related by either blood or by marriage or adoption, who live together as one household." The 1930 and 1940 definition is the share of non-white people residing in a census tract.
Median Contract Rent	Median contract rents per month for rental-occupied dwellings within a census tract
Homeownership Rate	Ratio of owner-occupied to occupied dwellings in a census tract.
Share of Multifamily Units	Share of occupied dwellings in structures exceeding three dwellings per structure or two dwellings per structure in Louisville or New York City
Population Density	Total population in thousands residing in a census tract per square mile.

Table 2 Weighted Summary Statistics by City

Means	Share Black 34	$\Delta White_{34-40}$	$\Delta \text{Black}_{34-40}$	N
Boston	2.75	0.051	0.0073	127
Chicago	7.78	0.034	0.0088	916
Louisville	15.85	0.12	0.0079	89
New York	4.44	0.057	0.025	2862
Philadelphia	11.46	0.023	0.022	396
Washington DC	25.07	0.21	0.13	95

Note: $\Delta White_{34-40}$ indicates the change in the white population from 1934-40 as a share of the total population in 1934. Summary statistics weighted by 1934 tract-level population.

Table 3 Percent change in white population 1934-40 from total population in 1934 over the distribution of share black in 1934.

Share Black in 1934	Mean (Pct Change White Pop 34-40)	SD	N
0-1	0.069	0.76	3462
1-5	0.044	0.31	433
5-10	0.032	0.34	146
10-20	0.020	0.39	134
20-30	-0.011	0.18	70
30-40	-0.025	0.21	47
40-50	-0.087	0.19	28
50-100	-0.018	0.24	165

Note: Summary statistics weighted by 1934 population.

Table 3 Estimates of Tipping Points (1930-34)

	Full Sample		Share Black < 60	
	Unweighted	**Population**	**Unweighted**	**Population**
Boston	0.5	0.5	0.5	0.5
Chicago	0.1	0.4	2.7	0.4
Louisville	---	---	---	---
New York	0.1	0.9	0.1	0.9
Philadelphia	---	---	---	---
Washington DC	4.3	4.3	4.3	4.3

Note: Estimation of Tipping Point constructed using the Structural Break Method

Table 4 Estimates of Tipping Points (1934-40)

	Full Sample		Share Black < 60	
	Unweighted	**Population**	**Unweighted**	**Population**
Boston	5.8	0.6	5.8	50
Chicago	23.9	23.9	13	0.2
Louisville	43.5	3.1	3.1	3.1
New York	0.1	2.4	0.1	2.4
Philadelphia	0.1	0.1	0.1	0.1
Washington DC	7.9	10.1	7.9	10.1

Note: Estimation of Tipping Point constructed using the Structural Break Method

Table 5 Estimates of Tipping Points (1930-40)

	Full Sample		Share Black < 60	
	Unweighted	**Population**	**Unweighted**	**Population**
Boston	0.6	0.6	0.6	0.6
Chicago	1.1	2.7	1.1	2.7
Louisville	---	---	---	---
New York	0.1	1.5	0.1	1.5
Philadelphia	---	---	---	---
Washington DC	15.4	15.4	15.4	15.4

Note: Estimation of Tipping Point constructed using the Structural Break Method

Table 6 Estimation Table: Dependent Variable: Change in Share of White Population as a share of base year total population

	1930-34		1934-40		1930-40	
	Unweight	Pop Weight	Unweight	Pop Weight	Unweight	Pop Weight
	b/(se)	b/(se)	b/(se)	b/(se)	b/(se)	b/(se)
Beyond Tipping Point	-0.386***	-0.0692***	-0.888**	-0.00909	-1.619	-0.0786
	(0.0794)	(0.00979)	(0.356)	(0.0272)	(0.429)	(0.0304
Difference in Share	0.0106	-0.00659***	-0.00252	-0.00366	0.0429	-0.00778
Nonwhite and Tipping Point	(0.0206)	(0.00256)	(0.0402)	(0.0024)	(0.0799)	(0.00365)
Difference Squared	-0.000391	0.000448***	-0.000284	-0.0000339	-0.0018	0.000285
	(0.00129)	(0.000148)	(0.00218)	(0.0000392)	(0.00494)	(0.000204)
Difference Cubed	0.00000469	-0.00000836***	0.00000494	0.00000155*	0.0000292	-0.00000498
	(0.0000241)	(0.00000265)	(0.0000497)	(0.000000915)	(0.0000971)	(0.00000408)
Difference 4th	-1.20E-08	4.76e-08***	-9.00E-09	-8.66E-09	-0.000000139	3.14E-08
	(0.000000136)	(1.45E-08)	(0.00000034)	(8.55E-09)	(0.000000569)	(2.44E-08)
Population Density 000s	-0.00372***	-0.000826***	-0.0198***	-0.00141***	-0.0151***	-0.00157***
per Sq Mi in Base Year	(0.000793)	(0.0000649)	(0.00492)	(0.000273)	(0.00468)	(0.000164)
Homeownership Rate			-0.0365***	-0.00115*		
in Base Year			(0.00733)	(0.000684)		
Share of Multi-Fam			-0.0159*	0.000327		
Dwellings in Base Year			(0.00864)	(0.00056)		
Median Rent in Base Year			-0.0131*	0.00282***		
			(0.00796)	(0.000628)		
Constant	0.21	-0.00112	2.803***	0.19	1.24	0.0873
	(0.167)	(0.0132)	(0.859)	(0.175)	(0.986)	(0.034)
City Fixed Effects	Included	Included	Included	Included	Included	Included

Note: Tipping points estimated using tracts where the share of non-white families < 60%. Regressions run on full sample.