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Abstract

Neighborhood racial and class segregation continue to be major social problems within America's metropolitan areas. Segregation has been linked to a whole host of racial and class inequalities, including access to jobs, schooling and single parenthood, and future earnings. One factor accounting for segregation is the inability of black and lower income households to afford housing in white neighborhoods, where housing units historically have been largely owner-occupied single-family homes. In recent years there has been a shift in the housing makeup of many of these neighborhoods, with rentals and foreclosures increasing in share. This has made housing more affordable in these neighborhoods. In this paper we investigate the impact that foreclosures and three types of rentals (single-family, condominium, and apartments) have on neighborhood racial and income integration using data from Miami-Dade County, Florida. We find that neighborhoods have become more racially and socially integrated as rentals have increased as a share of the housing stock.

Keywords: racial segregation, income segregation, rental housing

JEL Classification: J15, R21, R23, R31

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1. Introduction

Neighborhood racial and class segregation continue to be major social problems within America's metropolitan areas. Segregation has been linked to a whole host of inequalities, including access to jobs (Weinberg, 2000), schooling and single parenthood (Cutler and Glaeser, 1997), and future earnings (Chetty et al., 2014). One factor accounting for segregation is that housing in white neighborhoods tends to be relatively expensive, owner-occupied, single-family homes that many lower income and black households find unaffordable. Either these households cannot afford the price of these homes or they are unable to satisfy mortgage underwriting criteria. However, housing affordability is improving in many of these neighborhoods.

As is well known, foreclosures grew in the years following the Great Recession. Foreclosures tend to sell at a discount and their negative spillover effects tend to depress the values of neighborhood properties. Less well known is that there has been a remarkable shift in the housing tenure of many urban neighborhoods in favor of rentals, especially single-family rentals.[1] The latter is an important legacy of the housing market crash that accompanied the Great Recession that initiated the conversion of millions of single-family homes within America's neighborhoods that were once owner-occupied into rentals, as billions of dollars of private equity have poured into the single-family home rental business.[2] By making housing more affordable, the growth in foreclosures and rentals have expanded the housing opportunities of minority and lower income households, especially in higher quality neighborhoods.

The question we address in this paper is whether this increase in affordability has resulted in more racially and socially mixed neighborhoods. Extant evidence on this issue is thin and mixed; hence, there is

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[1] Based upon property tax records from Florida, Ihlanfeldt et al. (2018) show that on average across Florida's urban counties the share of housing units represented by rentals increased six percentage points from 42 percent in 2000 to 48 percent in 2014, while the growth in the single-family rental share increased seven percentage points from 15 to 22 percent. Based on their own calculations using data from the American Community Survey (ACS), the rentals share at the national level measured over a shorter period (2005—2014) grew from 33.1 to 36.9 percent, while the share of single-family rentals grew from 10.2 to 12.9 percent. They note that the ACS does not report the rental shares broken down by urban versus rural, but based on their findings from Florida the increase in the shares of rentals for urban areas nationally is expected to be much greater than for urban and rural areas combined.

[2] The lion's share of the investment in single-family rentals has originated from institutional investors (Smith and Liu, 2017; Mills et al., 2016). Mills et al. (2016) conclude that these investors are in for the long haul and do not intend to liquidate their rental holdings anytime soon. This suggests that single-family rentals in neighborhoods are here to stay.

a clear need for further investigation. Our analysis is based on a panel of all of the neighborhoods in Miami-Dade County, Florida covering the years 2006 through 2016. For the years 2006 through 2012 we relate shifts in a neighborhood's housing stock in the current year in favor of single-family REOs and three different types of rentals (single-family homes, apartments, and condominiums) to changes in the U.S. Census Bureau's estimated average racial and income composition within the neighborhood measured over the current and future four years.[3] Hence, the estimation allows changes in the neighborhood's shares of these housing types to have a long run effect on the percentage the neighborhood's residents who are black and the percentage of the households in the neighborhood who are poor.[4] Having found that rentals increase the black percentage, we then estimate the net number of non-Hispanic blacks, non-Hispanic whites, and Hispanics who move into or out of the neighborhood in response to changes in the housing types. This allows us to investigate whether a rental induced increase in the percentage of blacks living in the neighborhood is driven by black entry or the exit of the other two racial groups. If the rentals are increasing the black share of the neighborhood by inducing non-black flight, this would distract from the attractiveness of polices to place more rental housing in better neighborhoods. Similarly, we explore the extent to which the migrations of poor and non-poor households in response to changes in the different housing types account for their effect on the percentage of poor households living in the neighborhood. An advantage of our approach is that we estimate control function models that enable us to test for the endogeneity of each of the housing types and instrument them where necessary using a convincing identification strategy.

Separate models are estimated using all of the county's neighborhoods, neighborhoods where blacks or poor households are underrepresented, and neighborhoods found within white suburban jurisdictions. Poor households are divided into poor (household income less than the county median) and very poor (household income less than one half of the county median). The results show that an increase in the share of the

[3] REO is the acronym for Real Estate Owned properties, which are homes that are owned by financial institutions as a result of foreclosure proceedings.
[4] An alternative approach to capturing the long run effect of an increase in a housing type on neighborhood segregation would be to regress an annual measure of neighborhood racial composition on the share of the type of housing in the current year and in lagged years to obtain the long run propensity. The ACS, however, does not report neighborhood race annually at the block group level (our neighborhood unit). Only a five-year average is reported by the ACS.

neighborhood's housing units that are apartments increases the black, poor, and very poor percentages across all three samples. Condominium rentals have similar effects, while single-family rentals affect only the income and not the racial mix of neighborhoods, suggesting that racial discrimination may play a more important role in the single-family rental market in comparison to the other rental markets. An increase in the share of REOs generally fails to affect either the racial or the income mix of the neighborhood. Our estimated migration models show that within neighborhoods where blacks are underrepresented the positive effects that apartments and condominium rentals have on the black percentage of the neighborhood come largely from black entry rather than non-black exit. In contrast, in neighborhoods where poor and very poor households are underrepresented, the apartments and condominium rentals induced rise in the poor and very poor percentages are found to come from both poor entry and non-poor exit, while the positive effect that single-family rentals have on the poor percentages in these neighborhoods is entirely the result of non-poor exit.

2. Literature Review

On the surface the idea that improving housing affordability in better neighborhoods should result in a residential reshuffling of a community's disadvantaged households in favor of these areas may seem self-evident. However, there are reasons to suggest that such resorting may not occur. First, blacks and lower income households may encounter housing market discrimination that limits their ability to take advantage of cheaper housing in neighborhoods where high percentages of whites and higher income households reside. Second, the locational decisions of disadvantaged households may be governed by other stronger needs that keep them living within segregated neighborhoods. These include access to jobs, welfare services, or public transportation. Finally, it has been suggested that the distance of the move, which may be quite lengthy involving a move from one side of the urban area to the other, may be too arduous for some families to overcome (Ellen et al., 2016). Long distances may also limit the knowledge that disadvantaged household have of the affordable housing opportunities found within better neighborhoods.

Hence, the correct answer to the question of whether housing segregation declines in response to more affordable housing within white or higher income neighborhoods is unclear, a priori.

To our knowledge, no study has investigated the effect that rentals have on neighborhood housing integration. There are, however, a number of studies whose results relate to the issue of whether minority and low income families move to better neighborhoods in response to an increase in housing affordability. One set of papers investigates whether lower income and minority families residentially relocate so that their children can attend a better school if housing within that school's attendance zone becomes more affordable (Ihlanfeldt and Mayock, 2018a,b; Ihlanfeldt, 2019). Using data on Florida school districts these papers are differentiated by their focus on different disadvantaged groups, the use of alternative measures of affordable housing and their adoption of different identification strategies. Remarkably, they share a common conclusion; namely, that placing a larger share of a district's affordable housing within better school zones reduces racial and income school segregation. A limitation of this research is that only short run outcomes are analyzed. If the entrance of minority or lower income families into the neighborhood or their children into the neighborhood school induces whites to flee, the long and short run outcomes may be very different and it is the long run effect that garners the greater interest.

Evidence inconsistent with that from the above studies comes from the Moving To Opportunity (MTO) experiment (Sanbonmatsu et al., 2011) and national data on the residential locations of housing voucher recipients (Horn et al., 2014). The MTO experiment offered housing vouchers to two groups of eligible households. One group had no constraint on where they could locate, while the other group was required to locate in a low poverty neighborhood. Interestingly, the uptake rates (i.e., the percentage of households accepting the vouchers) were low. The percentages were 63% and 48% for the locational unconstrained and constrained groups, respectively. These results suggest that many disadvantaged households may not move to better neighborhoods when the opportunity presents itself. Similar findings are provided by Horn et al. (2014). Using confidential data on the residential locations of voucher recipients from the Department of Housing and Urban Development (HUD), the authors find that the children of voucher holders were more likely to attend low-performing schools than the children of households that did not receive housing vouchers.

In summary, the results of some studies suggest that improved housing affordability can reduce neighborhood segregation, while the results of other studies suggest that disadvantaged households do not move or locate in favor of better neighborhoods in response to offered or accepted housing vouchers. As noted earlier, a limitation of the first group of studies is that they focus exclusively on short run effects. In addition, only the movement of families with children are considered. Studies in the second group are limited in that only the location decisions of subsidized households are considered. Our analysis is not subject to either of these limitations. We investigate the long run changes in neighborhood racial and income composition that result from an improvement in housing affordability, as captured by a neighborhood's shift in housing tenure in favor of different types of rentals, where all blacks and lower income households are included in the analysis.

3. Conceptual Framework[5]

Extant evidence indicates that a shift of a neighborhood's housing stock in favor of REOs or rentals has two effects. On the one hand, they lower neighborhood quality in two ways: 1) they produce negative sight externalities due to their lower upkeep and 2) they raise the level of neighborhood crime.[6] On the other hand, the deferred maintenance of REOs decreases their price and this combined with their positive impact on crime creates negative spillover effects that increases the affordability of housing.[7] Rentals may improve affordability by producing similar negative spillover effects; however, the principal affordability advantage

[5] Our conceptual framework, presented descriptively in this section, is based on a static general equilibrium model of the impact of foreclosures on neighborhood residential sorting that we developed in previous work (Ihlanfeldt and Mayock, 2018b). We refer readers to this paper for a more formal theoretical model underlying our empirical work.
[6] Evidence on the poor physical condition of foreclosures is provided by (Gerardi et al., 2015) and the National Fair Housing Alliance (2012). Many studies have found that rentals are less well maintained than owner-occupied housing (Galster, 1983; Gatzlaff et al., 1998; Harding et al., 2000; Shilling et al., 1991). Ellen et al. (2013) find that neighborhood crime is positively associated with foreclosures, while Ihlanfeldt et al. (2018) demonstrate a relationship between neighborhood crime and rentals.
[7] A plethora of studies have found that foreclosures lower the rents and values of nearby properties. For a review of these studies see Ihlanfeldt and Mayock (2016b).

of rentals comes from their lower cost of neighborhood entry in comparison to owner-occupied units.[8] In sorting across neighborhoods households weigh the affordability advantage provided by REOs and rentals against their negative impact on neighborhood quality. The relative importance of these two factors may vary with the race and income level of the household. There are many possible scenarios. Our interest is whether blacks and lower income households give more weight to the affordability advantage and thereby sort into neighborhoods with larger shares of REOs or rentals. If so, this has the potential of increasing the percentage of the neighborhood's households who are black or low income. However, these percentages may also rise if whites or high income households place more weight on neighborhood quality than affordability in making their location decisions, causing a net outmigration from neighborhoods with larger shares of REOs or rentals. This underscores the importance of empirically addressing not only the change in the racial or income composition of neighborhoods in response to an increase in their REO or rental shares but also the magnitudes of the in-migration and out-migration of racial and income groups in possibly accounting for these results.

4. Data

Our study area is Miami-Dade County, Florida. The County has a large population (2.5 million, according to 2010 Census), which includes large percentages of all three racial groups---non-Hispanic blacks, non-Hispanic whites, and Hispanics and high levels of both racial and income neighborhood housing segregation.[9] In addition, over the course of our panel the housing stocks in a majority of the neighborhoods within the County experienced a shift in favor of REOs and rentals, where a neighborhood is defined as a census block group.[10] Hence, it is an ideal area for study given the purpose of our

[8] The entry cost on rental housing is the security deposit and required renters' insurance, while the upfront cost of obtaining owner-occupied housing is the down payment on the mortgage and associated fees.
[9] The St. Louis Federal Reserve Bank reports the white to non-white racial dissimilarity index for Miami-Dade County equaled 48.83 in 2016: https://fred.stlouisfed.org/series/RACEDISPARITY012086. The Racial Dissimilarity Index measures the percentage of the non-Hispanic white population in a county which would have to change Census tracts to equalize the racial distribution between white and non-white population groups across all tracts in the county. The Pew Research Center lists the Miami Metro Area as having the tenth worst neighborhood income segregation among the Nation's largest 30 areas: https://www.pewsocialtrends.org/2012/08/01/the-rise-of-residential-segregation-by-income/.
[10] The block group is a geographical subunit of a census tract. Typically, block groups have a population of 600 to 3,000 people.

investigation. Two main sources of data were used to construct our panel of neighborhoods. To construct our housing variables, we used property tax records from the Florida Department of Revenue for the years 2006—2012. These records provide a complete count of all housing types within each neighborhood. The types include single-family, condominium and apartment homes and REOs.

Importantly for our study, the tax roll data also contains fields that indicate whether or not a property was granted a property tax homestead exemption. This exemption is available to a person who makes the property his or her permanent residence. We use the presence of a homestead exemption to classify a property as owner-occupied, and housing units without a homestead exemption are classified as renter-occupied. Because the exemption provides significant tax savings, owner-occupants have strong financial incentives to file for the exemption, and we are thus confident that owner-occupied units will generally be correctly classified based on homestead status.[11] Properties that are not covered by a homestead exemption are primarily either rental units or second homes. The fraction of non-homesteaded single-family homes that are second homes is expected to be small because in Florida most vacation homes are condominiums. For condominiums we cannot rule out the possibility that some of the properties we label as rentals may in fact be second homes not available for rent.

Our second main source of data comes from the American Community Survey (ACS). For each neighborhood (block group) the ACS reports a five-year average of the number of each racial group residing in the neighborhood. Also provided is the average number of households falling within 17 different income groups. From these data we calculated the average percentages of the neighborhood's population who are black, who are poor (income less than the county median), and who are very poor (income one-half of the county median). We match each year for which we measure the composition of the housing stock with the ACS averages. For example, for 2006, the first year of our panel, each of the housing types within the neighborhood are counted on January 1, 2006. The associated five-year averages from the ACS covers the

Of the county's 1,534 block groups, 55% and 65% experienced an increase in REOs and rentals as a share of the housing stock, respectively. The housing stock is defined as including REOs, single-family rental and owner-occupied homes, condominium rental and owner-occupied homes, and apartments. Rentals are identified based on the homestead exemption as described below. Corresponding percentages for single-family rentals and condominium rentals equal 58% and 42%, respectively.

[11] A homestead exemption decreases a property's taxable value by as much as $50,000.

years 2006—2010. At the end of our panel, the housing counts are for January 1, 2012 and the ACS averages are for the years 2012—2016.

Another source of data is DataQuick, which we used to get annual counts of the number of REOs in the neighborhood.[12] We also drew from the property tax records to include in our panel the number of different types of non-residential land uses for each neighborhood/year observation. These included stores, service stations, restaurants, clubs and bars, office buildings, and churches. We used these variables as control variables in testing the robustness of our housing shares findings to the inclusion of properties other than housing units that may impact the attractiveness of the neighborhood to different racial and income groups.

5. Empirical Methodology

We are interested in empirically addressing two issues with our panel of Miami-Dade County neighborhoods. First, how do increases in REOs and different types of rentals as a share of the neighborhood's total housing units affect the racial and income composition of the neighborhood. Subsection 5.1 describes the neighborhood composition models we estimated to shed light on this issue. Second, to what extent are changes in the composition of neighborhoods due to black/low income entry versus white/high income exit in response to changes in the housing shares. The migration models we estimated to address this question are described in Subsection 5.2.

5.1 Neighborhood Racial and Income Composition Models

To address the first issue, we first constructed a variety of neighborhood samples. These included all neighborhoods ($n = 1{,}534$), neighborhoods where black people were underrepresented ($n = 1{,}120$), neighborhoods where poor households were underrepresented ($n = 788$), neighborhoods where very poor households were underrepresented ($n = 883$) and neighborhoods within suburban white jurisdictions ($n = 408$). A group is considered underrepresented within a neighborhood if at the beginning of our panel

[12] The algorithm we used to identify REOs with these data is described in detail in our earlier paper (Ihlanfeldt and Mayock, 2016a).

the neighborhood percentage of the group is less than the county-wide percentage.[13] The neighborhoods within the collection of white suburban jurisdictions are known to have the highest quality neighborhoods and schools within the county.[14] Blacks are especially underrepresented in these places, with the average black neighborhood share at the beginning of our panel equaling only 1.68%. The corresponding poor (49.7%) and very poor (26.2%) household percentages, while lower than the county-wide percentages, are much closer to the county-wide averages. The suburban jurisdictions sample was employed to determine if a shift in favor of REOs or rentals enable blacks and the poor (very poor) to move into the highest quality neighborhoods.

With our samples in hand, we then regressed in turn the percentages of blacks, the poor, and the very poor, measured over the years $t = 1$ to $t = 5$ on the $t = 1$ shares of the neighborhood's housing stock represented by single-family rentals, condominium owner-occupied and rental units, apartment rentals, and REOs.[15] The excluded housing share from the equation is single-family owner-occupied units. Separate models were estimated for all neighborhoods, neighborhoods where the group was underrepresented and neighborhoods within the suburban jurisdictions. Formally, our neighborhood composition models can be expressed as:

$$y_{igt} = \boldsymbol{s}'_{igt}\boldsymbol{\beta} + \boldsymbol{x}'_{igt}\boldsymbol{\delta} + \alpha_g + \theta_t + \varepsilon_{igt}, \tag{1}$$

where i denotes the i^{th} neighborhood (block group), g denotes the g^{th} census tract, and t is time period index, for $i = 1, 2, \dots, n$; $g = 1, 2, \dots, G$; and $t = 1, 2, \dots, T$. The dependent variable y_{igt} is the five-year average covering years t through $t + 4$ of the percentage of non-Hispanic blacks or the percentage of the poor (very poor) residing in neighborhood i within tract g, for the racial and income composition models, respectively. $\boldsymbol{s}_{igt} = (s_{igt,1}, s_{igt,2}, \dots, s_{igt,J})'$ is the $J \times 1$ vector of possibly endogenous housing shares

[13] The county-wide percentage of black people equaled 17.7%. The county-wide percentages of poor and very poor households equaled 55.5% and 29.6%, respectively.
[14] The jurisdictions are Aventura, Bal Harbour, Bay Harbor Islands, Biscayne Park, Coral Gables, Doral, Golden Beach, Hialeah, Hialeah Gardens, Indian Creek, Medley, Miami Beach, Miami Lakes, Miami Springs, North Bay Village, Palmetto Bay, Pinecrest, Sunny Isles Beach, Surfside, Sweetwater, The Village of Key Biscayne, Virginia Gardens, and West Miami.
[15] Means and standard deviations of the housing shares are reported in Table A.1 of Appendix A.

measured for neighborhood i of tract g on January 1 of year t, with the associated vector of constant coefficients $\boldsymbol{\beta}$. $\boldsymbol{x}_{igt}$ is a $K \times 1$ vector of strictly exogenous control variables (conditional on α_g) with the associated parameters $\boldsymbol{\delta}$.[16] α_g represents the unobserved time-invariant tract-specific heterogeneity, which can be correlated with all explanatory variables. θ_t signifies the year fixed effects. ε_{igt} is the idiosyncratic time-varying error component.

To obtain unbiased and consistent estimated effects of the housing shares on the racial and income composition of the neighborhood, the condition of strict exogeneity must be satisfied (Wooldridge, 2010, p.322). Strict exogeneity can be violated by reverse causation or by unobservables that effect neighborhood composition and are correlated with the housing shares. Because we are estimating how the housing shares drive the future racial composition of the neighborhood, it is unlikely that our estimated effects are biased by reverse causation. Unobservables that may affect both the dependent and independent variables that are unchanging over time are captured by our inclusion of tract fixed effects. There remains the possibility that strict exogeneity is violated by time-varying unobservables. In this case, the affected housing shares require instrumentation. Econometrically, the challenge is in determining which shares to instrument.

We estimate model (1) using the control function (CF) approach. The CF approach is an alternative estimation strategy to the standard instrumental variable (IV) methods, including two stage least squares (2SLS) and generalized method of moments (GMM), which deal with the endogeneity problem. In linear (in parameter) models where the endogenous variables also show up linearly, the CF approach yields identical point estimates to the 2SLS estimates. However, the CF approach offers appealing advantages for models nonlinear in endogenous variables and/or nonlinear in parameters. As we will see, the Poisson model that we used in the study of migration in the next subsection is an example of the latter case and the endogeneity of the housing share variables cannot be readily handled by the standard IV methods.

[16] The control variables are building types that may affect a neighborhood's housing costs or its amenity value, apart from the effects of the housing shares. There are seven variables, which are the number of stores, service stations, restaurants, clubs/bars, office buildings, industrial properties, and churches. While we include the control variables in (1), they are not included in the models estimated that yielded the results reported in our tables. Our results are robust to their inclusion and they are generally statistically insignificant either together or alone. Details are provided below.

Even in the case of linear models as in (1) where the CF and 2SLS estimates coincide, the CF method provides an attractive estimation strategy. By including as regressors the residuals from the reduced form models for the endogenous variables, it not only corrects for endogeneity but also produces a simple variable addition test that is easily made robust to unknown heteroscedasticity and serial correlation.[17] Hence, for both our neighborhood composition and migration models the CF method addresses the issue of which shares should be treated as endogenous. Therefore, we estimated both composition and migration models by the CF approach. The CF approach requires the availability of valid instruments. Before describing our choice of instruments, we first explain the CF estimation of model (1).

Assume that $\boldsymbol{z}_{igt}$ is an $L \times 1$ vector of strictly exogenous variables (conditional on α_g), where $\boldsymbol{z}_{igt} = (\boldsymbol{x}'_{igt}, \boldsymbol{z}'_{igt2})'$ with $\boldsymbol{z}_{igt2}$ being exogenous variables excluded from equation (1). Suppose that a linear reduced form of $\boldsymbol{s}_{igt}$ is given by

$$\boldsymbol{s}_{igt} = \boldsymbol{\Pi}\boldsymbol{z}_{igt} + \boldsymbol{a}_g + \boldsymbol{\lambda}_t + \boldsymbol{e}_{igt}, \quad (2)$$

where $\boldsymbol{\Pi}$ is a $J \times L$ matrix of parameters, $\boldsymbol{a}_g$ and $\boldsymbol{\lambda}_t$ are $J \times 1$ vectors of tract and time fixed effects, respectively, $\boldsymbol{e}_{igt}$ is a $J \times 1$ vector of idiosyncratic errors.[18] Then we model the time-invariant heterogeneity by the Mundlak (1978) device:

$$\alpha_g = \psi_1 + \bar{\mathbf{z}}'_g \boldsymbol{\xi}_1 + r_{g1}, \quad (3)$$

$$\boldsymbol{a}_g = \boldsymbol{\psi}_2 + \Xi_2 \bar{\mathbf{z}}_g + \boldsymbol{r}_{g2}, \quad (4)$$

for $g = 1, 2, \dots, G$, where $\bar{\mathbf{z}}_g = (1/Tn_g)\sum_{t=1}^{T}\sum_{i=1}^{n_g} \boldsymbol{z}_{igt}$, n_g is the number of block groups within the g^{th} tract, ψ_1, $\boldsymbol{\psi}_2$, $\boldsymbol{\xi}_1$, Ξ_2 are constant parameters of conformable dimensions, and r_{g1} and $\boldsymbol{r}_{g2}$ are the error terms. Note that we assume the time-invariant heterogeneity exists primarily at the tract rather than the block group level since census tracts "are designed to be relatively homogenous with respect to population characteristics, economic status, and living conditions" (U.S. Census Bureau definition[19]). In addition, using

[17] It is more difficult to make the traditional Hausman test that compares 2SLS and OLS robust to arbitrary heteroscedasticity and serial correlation.
[18] We are assuming the usual rank condition for identification holds.
[19] Available at https://factfinder.census.gov/help/en/census_tract.htm.

averages at the tract level helps alleviate multicollinearity between the housing shares and their time means.[20] Substituting (3) into (1) yields

$$y_{igt} = \psi_1 + \boldsymbol{s}'_{igt}\boldsymbol{\beta} + \boldsymbol{x}'_{igt}\boldsymbol{\delta} + \bar{\mathbf{z}}'_g\boldsymbol{\xi}_1 + \theta_t + u_{igt}, \tag{5}$$

where $u_{igt} = r_{g1} + \varepsilon_{igt}$. Assume that $Cov(\boldsymbol{z}_{igt}, r_{g1}) = \mathbf{0}$ for all t, and note that $Cov(\boldsymbol{z}_{igt}, \varepsilon_{igt'}) = \mathbf{0}$ for all t and t' by the strict exogeneity of $\boldsymbol{z}_{igt}$, we have $Cov(\boldsymbol{z}_{igt}, u_{igt'}) = \mathbf{0}$ all t and t'. Then inserting (4) into (2) yields

$$\boldsymbol{s}_{igt} = \boldsymbol{\psi}_2 + \boldsymbol{\Pi}\boldsymbol{z}_{igt} + \Xi_2\bar{\boldsymbol{z}}_g + \boldsymbol{\lambda}_t + \boldsymbol{v}_{igt}, \tag{6}$$

where $\boldsymbol{v}_{igt} = (v_{igt,1}, v_{igt,2}, \dots, v_{igt,J})' = \boldsymbol{r}_{g2} + \boldsymbol{e}_{igt}$. Assuming that $Cov(\boldsymbol{z}_{igt}, \boldsymbol{r}_{g2}) = \mathbf{0}$ for all t and $Cov(\boldsymbol{z}_{igt}, \boldsymbol{e}_{igt'}) = \mathbf{0}$ for all t and t', it is immediate that $Cov(\boldsymbol{z}_{igt}, \boldsymbol{v}_{igt'}) = \mathbf{0}$ for all t and t'. Now projecting u_{igt} onto $\boldsymbol{v}_{igt}$ gives

$$u_{igt} = \boldsymbol{v}'_{igt}\boldsymbol{\varphi} + \zeta_{igt}, \tag{7}$$

where $Cov(\zeta_{igt}, \boldsymbol{v}_{igt}) = \mathbf{0}$ by construction. Since both $u_{igt'}$ and $\boldsymbol{v}_{igt'}$ are uncorrelated with $\boldsymbol{z}_{igt}$, $\zeta_{igt'}$ is also uncorrelated with $\boldsymbol{z}_{igt}$ for all t and t'. Finally, plugging (7) into (5) leads to the following estimating equation:

$$y_{igt} = \psi_1 + \boldsymbol{s}'_{igt}\boldsymbol{\beta} + \boldsymbol{x}'_{igt}\boldsymbol{\delta} + \bar{\mathbf{z}}'_g\boldsymbol{\xi}_1 + \theta_t + \boldsymbol{v}'_{igt}\boldsymbol{\varphi} + \zeta_{igt}. \tag{8}$$

Note that ζ_{igt} is uncorrelated with all the variables on the right-hand-side of (8), including $\boldsymbol{s}_{igt}$. Let $s_{igt,j}$ denote the j^{th} element of $\boldsymbol{s}_{igt}$, where j indexes the type of housing share, $j = 1,2, \dots, J$. The inclusion of $v_{igt,j}$, in the original equation "controls" for the endogeneity of $s_{igt,j}$.

Since $\boldsymbol{v}_{igt}$ is unobserved, a two-step CF estimation procedure can be implemented as follows: In the first step, we estimate (6) by a pooled OLS regression and obtain the residuals, $\hat{\boldsymbol{v}}_{igt}$. In the second step, we estimate (8) by pooled OLS, where $\boldsymbol{v}_{igt}$ is replaced with $\hat{\boldsymbol{v}}_{igt}$. To test if the j^{th} type of housing share, $s_{igt,j}$, is exogenous, we can easily apply a cluster-robust t test of H_0: $\varphi_j = 0$, where φ_j is the j^{th} element of $\boldsymbol{\varphi}$.

[20] Because our panel of seven years is relatively short in duration, an explanatory variable and its time mean will obviously be highly correlated, creating a possible multicollinearity problem and inefficient estimates. The precision of our estimates was substantially improved by computing the time means at the tract level.

Since the second step utilizes generated regressors, the standard errors need to be adjusted for the first-step sampling error. We obtain valid standard errors by bootstrapping based on 1,000 replications if exogeneity is rejected.[21]

To obtain an instrument where exogeneity is rejected, we first defined a base year preceding the beginning of our panel. We chose 2004, two years before the start of the panel, as the base year.[22] Using the entire county, we then calculated the percentage change in the housing share (s) at the county level between the base and current years, excluding the home neighborhood value. These percentage changes were then multiplied by the base year value of the share to obtain a prediction of the current year share value ($\hat{s}$), assuming the growth in the share followed the change that occurred at the county level.[23] Formally,

$$\hat{s}_{igt,j} = s_{igt_0,j} \times \frac{s_{ct,j}}{s_{ct_0,j}}, \quad (9)$$

where as before i, g and j index block group, tract, and the type of housing share, respectively. c represents the county, t refers to the current year, and t_0 signifies the base year (2004). While the racial mix within the neighborhood may affect $s_{igt,j}$, it should not have an effect on $\hat{s}_{igt,j}$. However, the validity of our instrument also depends on whether the neighborhood base year housing share values are exogenous. For example, there may be an omitted variable that is correlated with the base year shares (say the level of neighborhood crime) that has a delayed impact (two years) on the racial or income composition of the neighborhood. In that case, our instrument would, in part, be capturing the crime effect and would not be orthogonal to the error term of equation (1). However, evidence suggesting that this potential problem may not be of sufficient concern to invalidate our instrument is based on the robustness of our findings from

[21] Under the null, we do not need to adjust the standard errors by bootstrapping.
[22] The housing shares for each neighborhood are available from the property tax records for years preceding the beginning of our panel. The year 2004 was chosen as the base year because it resulted in the most favorable first-stage diagnostics in the 2SLS estimation of model (1). These diagnostics were employed to judge the quality of our instruments, because first-stage diagnostics are not provided by the CF method. As noted below, using years earlier than 2004 as the base year had little effect on our estimated share effects, but as expected there was some loss in their precision.
[23] The general idea underlying these instruments is that changes within a smaller geographic unit of a larger unit containing the smaller unit are driven by factors within both the smaller and larger units.

reaching backward in time in defining our base year values. For example, using t-3 or t-4 as our base year yields results that mirror the results we report below for t-2.

Besides $\hat{s}$ satisfying the assumption of strict exogeneity, its validity as an instrument hinges upon its correlation with s. As noted above, for first-stage diagnostics we relied upon the estimation of 2SLS models. The results show that for all of our housing shares treated as endogenous, $\hat{s}$ and s are strongly correlated, with the first-stage F-statistic significant at better than the 1% level.[24] Therefore, $\boldsymbol{z}_{igt} = (\boldsymbol{x}'_{igt}, \hat{\boldsymbol{s}}'_{igt})'$ was used in the estimation of (6) and (8), where $\hat{\boldsymbol{s}}_{igt} = (\hat{s}_{igt,1}, \hat{s}_{igt,2}, \ldots, \hat{s}_{igt,J})'$.

5.2 Migration Models

To address our second issue of the roles played by the entry and exit of different racial/income groups in response to changes in the housing shares in generating the changes in the black/poor percentage of the neighborhood, we estimated the following Poisson regression model:[25]

$$E(C_{igt} | \boldsymbol{s}_{igt}, \boldsymbol{Z}_{ig}, \alpha_g, \theta_t, \varepsilon_{igt}) = \exp\left(\boldsymbol{s}'_{igt}\boldsymbol{\beta} + \alpha_g + \theta_t + \varepsilon_{igt}\right), \tag{10}$$

where as before i, g and t index block group, tract, and time period, respectively, for $i = 1, 2, \ldots, n$; $g = 1, 2, \ldots, G$; and $t = 1, 2, \ldots, T$. C_{igt} is the five-year average covering years t through $t+4$ of the number of people of interest residing in neighborhood i within tract g. Specifically, C_{igt} refers to the number of non-Hispanic blacks, or non-Hispanic whites, or Hispanics in the racial migration models; and C_{igt} represents the number of poor, or very poor in the income migration models. $\boldsymbol{s}_{igt}$ is the vector of housing shares measured for neighborhood i within tract g on January 1 of year t. α_g is unobserved time-invariant tract-specific heterogeneity, θ_t is year fixed effects, and ε_{igt} represents time-varying omitted factors. $\boldsymbol{Z}_{ig} = (\boldsymbol{z}_{ig1}, \boldsymbol{z}_{ig2}, \ldots, \boldsymbol{z}_{igT})'$, where $\boldsymbol{z}_{igt}$ is a column vector of strictly exogenous variables (conditional on α_g),

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[24] We also conducted the Sanderson-Windmeijer tests of underidentification and weak identification, respectively, of individual endogenous regressors. They are constructed by "partialling-out" linear projections of the remaining endogenous regressors. The null hypothesis is that the particular endogenous regressor in question is unidentified. Test results based on robust F-statistics indicated that the null could be rejected at better than the 1% level for each of our endogenous variables.

[25] To be clear, when we refer to entry (exit), we mean that the 5-year average of the number of the group increased (decreased), which means that on average the number entering was greater than (less than) the number leaving. We make no allowance for births and deaths, so there is some random error in our entry and exit measures.

which consists of the set of instrumental variables constructed following (9). Since (10) is nonlinear in parameters, it is difficult to apply the standard IV methods. However, a two-step CF estimation can be readily performed analogously to the estimation of the linear composition model.

In specific, using the Mundlak device to model the tract fixed effects as in (3), we have

$$E\left(C_{igt}\middle|\boldsymbol{s}_{igt},\boldsymbol{Z}_{ig},\theta_t,u_{igt}\right)=\exp\left(\psi_1+\boldsymbol{s}_{igt}'\boldsymbol{\beta}+\bar{\mathbf{z}}_g'\boldsymbol{\xi}_1+\theta_t+u_{igt}\right), \tag{11}$$

where $u_{igt}=r_g+\varepsilon_{igt}$. In addition, recall that under (2) and (4), the reduced form of $\boldsymbol{s}_{igt}$ is given by (6). Assume that $(u_{igt},\boldsymbol{v}_{igt}')$ is independent of $\boldsymbol{Z}_{ig}$, where $\boldsymbol{v}_{igt}$ is the reduced form error defined in (6). Further assume that $E\left[\exp\left(u_{igt}\right)\middle|\boldsymbol{v}_{igt}\right]=\exp(\eta+\boldsymbol{v}_{igt}'\boldsymbol{\rho})$, which is implied by joint normality of $(u_{igt},\boldsymbol{v}_{igt}')$ but can be more general.[26] We then obtain the following estimating equation:

$$E\left(C_{igt}\middle|\boldsymbol{s}_{igt},\boldsymbol{Z}_{ig},\theta_t,\boldsymbol{v}_{igt}\right)=\exp\left(\mu+\boldsymbol{s}_{igt}'\boldsymbol{\beta}+\bar{\mathbf{z}}_g'\boldsymbol{\xi}_1+\theta_t+\boldsymbol{v}_{igt}'\boldsymbol{\rho}\right), \tag{12}$$

where $\mu=\psi_1+\eta$.

Now it is readily seen that (12) can be estimated by the following two-step CF method: In the first step, we obtain the residuals, $\hat{\boldsymbol{v}}_{igt}$, from a pooled OLS regression of $\boldsymbol{s}_{igt}$ following (6). In the second step, we apply a quasi-maximum likelihood estimation (QMLE) to (12), where the unobserved $\boldsymbol{v}_{igt}$ is substituted with the first-step residuals, $\hat{\boldsymbol{v}}_{igt}$. As for the composition models, to test if the j^{th} type of housing share, $s_{igt,j}$, is exogenous, the usual robust t test of H_0: $\rho_j=0$ can be easily applied, where ρ_j is the j^{th} element of $\boldsymbol{\rho}$. If exogeneity is rejected, we obtain valid standard errors for all coefficients by bootstrapping based on 1,000 replications.

6. Results from Estimating the Neighborhood Composition Models

Tables 1-3 present the results from estimating our percent black, percent poor, and percent very poor equations using as observations all neighborhoods within the county (Table 1), neighborhoods where each group is underrepresented (Table 2), and neighborhoods located within the white suburban jurisdictions of

[26] Note that in contrast with assumption (7) for the composition model, the linear projection argument no longer applies due to nonlinearity of the Poisson model.

the county (Table 3). For each housing share explanatory variable three numbers are given: the estimated coefficient, the cluster-robust standard error (in parentheses), and the change in the dependent variable induced by a within standard deviation change in the variable, holding all other variables constant (in brackets).[27]

Across all three samples of neighborhoods, the share of the neighborhood's housing units represented by apartments is found to have a positive and statistically significant effect on the percentage of the neighborhood's population who are black and the percentages of the neighborhood's households who are poor or very poor. Of particular interest are the magnitudes of the estimated effects obtained for the samples of neighborhoods where each group has been historically underrepresented. The change in percent black induced by a within standard deviation increase in the share of apartments is .093, which is a non-trivial increase when judged against the sample mean percent black (3.28%). The corresponding changes for the percentage of households who are poor or very poor are 1.32 and .72, which are also economically significant changes based on the dependent variable's mean value (reported at the bottom of the tables). These results suggest that increasing the share of apartments can result in greater racial and income integration within neighborhoods where blacks and lower income households are underrepresented. The results obtained for neighborhoods within the white suburban jurisdictions indicate an increase in the apartments share could also allow more integration of the county's higher quality neighborhoods. A within standard deviation increase in the apartment share increases percent black, percent poor, and percent very poor by .21, 1.88 and .88, respectively.

The impacts of condominium rentals as a share of a neighborhood's housing units on the neighborhood black, poor and very poor percentages closely match the results obtained for apartments when using the all neighborhoods and underrepresented neighborhoods samples. For the latter neighborhoods, the increase in percent black induced by a within standard deviation increase in the share of condominium rentals is .224,

[27] Because we are using panel data a standard deviation change can be computed between two randomly selected neighborhoods or between two randomly selected years within the same neighborhood. We use the within change rather than the between change because it is the more conservative number and has the more meaningful policy interest.

while the corresponding changes for the percentages of households who are poor or very poor are 1.82 and 1.03. However, in contrast to the apartments share, which is found to effect all three group percentages within neighborhoods located in the white suburban jurisdictions, the condominium rentals share only effects the poor percentage within these neighborhoods, registering an increase of .880 from a within standard deviation increase.

Shares of the final type of rentals, those provided by single-family homes, produce results that differ from those of the shares of apartments and condominium rentals. While the latter rentals have important effects on the percentage of a neighborhood's residents who are black, the share of single-family rentals has a statistically insignificant effect on the black percentage for all three neighborhood samples. However, the single-family rentals share is found to increase the percentage of a neighborhood's households who are poor or very poor within neighborhoods where these groups are underrepresented. A within standard deviation increase in a share yields increases of 1.82 and .59 in the poor and very poor household percentages, respectively. The failure of single-family rentals to racially integrate neighborhoods may reflect relatively greater discrimination in the renting of single-family homes. While we know of no evidence that this is the case, unlike multi-family rental housing, single-family rentals are not covered by federal fair housing laws. In addition, many single-family rentals are owned by small investors who have direct contact with their tenants.[28] If owners are prejudiced, they may choose not to rent to people of color. Another consideration might be an interest to maintain cordial relations with ex-neighbors, among those owners who previously lived in the neighborhood.

REOs potentially improve housing affordability within a neighborhood by selling at a discounted price and by depressing the values of nearby homes and rentals. While we find some evidence that an increase in REOs as a share of a neighborhood's housing units increase the percent black and percent poor of the neighborhood, the results are spotty in comparison to those obtained for the rental shares. In only two cases did we find REOs yielding a statistically significant effect. Using the full sample of neighborhoods, a within

[28] Based on data obtained from CoreLogic, Mills et al. (2016) calculate that 38% of single-family rentals are owned by investors holding one or two units, while 60% of the rentals are owned by investors having ten or less properties.

standard deviation increase in the REO share increases percent poor by .415. For the white suburban neighborhoods sample, the same increase in the REO share raises percent black by .181. In none of the underrepresented samples did REOs yield a significant effect. These results suggest that REOs may have a weaker effect on housing affordability in comparison to rental housing. Alternatively, the affordability advantage offered by REOs may be largely offset in the minds of blacks and lower income households by their negative impact on neighborhood quality.

Finally, in addition to the rental and REO share variables, all of our estimated models included the share of the neighborhood's housing units represented by owner-occupied condominium units. An increase in the share of these homes is found to have no effect on neighborhood racial integration, but across all three neighborhood samples the effect is positive and significant for percent poor and percent very poor. While this unexpected finding merits additional inquiry, one testable hypothesis is that these are actually rental units where the owner has illegally received the homestead exemption. Another possibility is that these units are being sold under the county's first-time homebuyer program, which provides mortgage loan subsidies to lower income households.[29]

As a check on the robustness of our findings to the inclusion of non-residential properties, we reran our neighborhood composition models including, along with the housing shares, the neighborhood number of each of the following types of land uses: stores, service stations, restaurants, clubs and bars, office buildings, and churches. Different racial and income groups may vary in their preferences for having these land uses within their neighborhood and the presence of these land uses may be correlated with our housing shares. Hence, their inclusion may result in omitted variable biases. Generally, these land uses yielded only a few significant effects on the racial or income mix of the neighborhood. Moreover, there inclusion had little impact on the results obtained for the housing shares. The land uses that were sometimes significant in affecting racial or income segregation were stores, restaurants and office buildings.

[29] A description of this program is provided at https://www8.miamidade.gov/global/housing/affordable-homeownership-program.page.

The results from estimating our racial and income neighborhood composition models suggest that the percentage of a neighborhood's population who are black and the percentage of a neighborhood's households who are poor depend on the share of the neighborhood's housing units that are rentals. The increase in these percentages from a larger share of rentals may be driven by black/low income in-migration, non-black/higher income exit, or some combination of both entry and exit. In the next section we report the results obtained from estimating our migration models, which are relevant to this issue.

7. Results from Estimating the Migration Models

We estimated separate Poisson Control Function (PCF) models explaining the five-year average (t to $t + 4$) neighborhood net migration of non-Hispanic blacks, non-Hispanic whites, and Hispanics using all three of our neighborhood samples. We also estimated PCF models explaining the neighborhood net migrations of poor, very poor, and non-poor households for all samples. The explanatory housing share variables, measured for t, are identical to those entering the neighborhood income and racial composition models. Our purpose was to explore, in those cases where a housing share had a statistically significant effect on the share of the neighborhood's residents who are black, the extent to which the change in the black share was driven by the housing share's effect on black entry versus its effect on the exit of whites or Hispanics. Similarly, we wished to learn the extent to which an induced housing share increase in the share of a neighborhood's households who are poor or very poor was the result of poor entry versus non-poor exit. The complexity of the task and the volume of results required summarizing the results in an easily readable format. Hence, we report the detailed results from the estimation of the PCF models in Appendix B and "comparison of effects" tables in this section (Tables 4-6). To illustrate the construction of these tables, we will focus on apartment rentals and neighborhood racial integration. To assess the importance of black entry we calculated how the mean black share of the neighborhood would be altered from the change in the number of blacks induced by a one percent increase in the apartments share, assuming there was no change in the number of whites or Hispanics. To accomplish this, we first calculated the black share as $B/(B + W + H)$, where B, W, and H are the sample mean numbers of each group. For the full sample of neighborhoods, this equaled 283/ (283+1068+297) or .1717. From the PCF model estimated for blacks the change in B from a one percent increase in the share of apartments is 5.6. So the new mean black share is now 288.6/ (288.6+1068+297) or .1745. The implied percentage increase in the mean black share is 1.63. The PCF model estimated for whites yields an exit of 2.6, so the new black share mean from this change alone is 283/ (283+1068+294.4) or .17199. The implied percentage increase in the mean share black is .17. Replicating this same procedure for Hispanics resulted in no increase in the mean black share. So we conclude from this experiment that the increase in the black share of the neighborhood's population from a one percent increase in the share of apartment rentals is being driven by both black entry and white exit, but the former effect is many times more dominate. The changes in the black share from black entry/exit, white entry/exit, and Hispanic entry/exit are labeled in the first comparison table (Table 4) the "black effect", "white effect", and "Hispanic effect" which are reported in Columns (1), (2), and (3), respectively. The fourth column of the table indicates the sign and statistical significance of the estimated effect of each of the housing share variables in the racial composition model. We are most interested in the comparison of effects (i.e., the importance of black entry relative to non-black exit) where the rentals shares have a positive significant effect on the share (percent) black.[30]

In addition to the share of apartments, the results from the racial composition model indicated that the condominium rentals share also has a positive and significant effect on the black share. As for apartments, the results from the PCF models based on the full sample of neighborhoods indicate that the condominium effect comes *almost entirely* from black entry (Panel A, Table 4). Using the sample of neighborhoods where blacks are underrepresented, the shares of apartments and condominium rentals are again found to have positive and statistically significant effects on the black share. The results from the estimated PCF models show that the increase in the black share from apartments and condominiums is coming *entirely* from black entry and not from either white or Hispanic exit (Panel B, Table 4). In the black share equation estimated

[30] For completeness, the tables also include a comparison of effects for owner-occupied condominium units and REOs. We restrict our discussion, however, to just rentals in light of their stronger performance in estimating the neighborhood racial and income composition models. The large magnitudes of the estimated effects of REOs in these tables and the tables in Appendix B are due to a one percentage point increase in an REO share representing a very large change relative to the means of the REO shares, which range from 0.13% to 0.28% in different samples (see Table A.1 in Appendix A).

for the neighborhoods within the white suburban jurisdictions only the share of apartments has a positive, significant effect. Again, the results from the estimation of the PCF models reveal that this effect comes entirely from black entry (Panel C, Table 4).

Turning to the results for the share of the neighborhood's households who are poor (Table 5), from the estimation of the neighborhood income composition model, increases in all three types of rentals increase the poor share and this is true across all three of our samples. In the case of an increase in the share of single-family rentals, the migration results for all three samples indicate that the induced increase in the poor share is coming entirely from the exit of non-poor households. In contrast, based on the full sample of neighborhoods increases in the poor share from apartments and condominium rentals comes entirely from poor entry (Panel A, Table 5). However, according to the migration estimates obtained from the underrepresented (Panel B, Table 5) and suburban samples (Panel C, Table 5), both poor entry and non-poor exit are contributing to the increase in the poor share from these two type of rentals. The results for the very poor household share of the neighborhood closely match those obtained for poor households (Table 6), with the exception that the share of single-family rentals increases the very poor share only from the underrepresented sample, but again the effect is coming almost entirely from non-poor exit.

8. Conclusion

In this paper our interest was in whether an increase in a neighborhood's share of housing units represented by REOs or different types of rentals results in a neighborhood becoming more racially or income integrated. Contributing to this interest is the recent growth in these shares, especially in single-family rentals. We offered a conceptual framework for why resorting across neighborhoods might occur in response to changes in these shares. REOs and rentals increase household utility by reducing the costs that must be paid to move into and live in the neighborhood, but reduce utility by eroding the quality of the neighborhood. Possible variance across racial and income groups in the strength of the pull of enhanced affordability versus that of the push of lower neighborhood quality may result in changes in the racial and income composition of neighborhoods. To address this issue empirically, we estimated control function models using panel data on Miami-Dade County neighborhoods. The CF models controlled for unobserved heterogeneity across neighborhoods, uniform time effects affecting all neighborhoods, and the possible endogeneity of the housing unit shares. CF models were also used to explore whether a change in the racial or income composition of a neighborhood in response to changes in the housing shares is due to black/lower income entry versus non-black/higher income exit.

Our results suggest that increases in the shares of apartments and condominium rentals increase both neighborhood racial and income integration. Moreover, it is black entry and not non-black exit that accounts for the improvement in racial integration. These findings support policies directed toward creating more of these types of rentals in an effort to achieve greater racial integration at the neighborhood level. The improvement in income integration that comes from these rentals is the result of both poor entry and non-poor exit. Hence, if the goal is to increase neighborhood income integration, these two effects need to be carefully considered. An important issue for further inquiry is why higher income households avoid neighborhoods with larger shares of these rentals. Is it the crime they engender, the negative sight externalities they emit, or perhaps the type of tenant that they attract? A neighborhood's share of single-family rentals has decidedly different effects on integration and the migration of groups accounting for any effect. While these rentals improve neighborhood income integration, they have no effect on racial integration. Moreover, in the former case, the improvement is entirely the result of higher income households exiting the neighborhood. Again, these results point to the need for further research. We have suggested that discrimination may play a more important role in the single-family rental market and that this may account for the failure of their neighborhood share to affect the racial mix. We encourage tests of this hypothesis. In addition, future research should focus on the factors underlying higher income flight in response to single-family rentals, along with our suggestion above for studying flight in response to the other two types of rentals. A key question is whether these factors differ across the different types of rentals.

The overarching policy implication that can be drawn from our findings is that increasing housing affordability within neighborhoods where blacks and lower income households have historically been underrepresented can result in more racially and income integrated local housing markets. As we alluded to in the introduction of this paper, there is a plethora of likely positive outcomes from achieving this result. Market trends in favor of rentals representing a larger share of the housing within America's urban neighborhoods are likely to continue, especially from the conversion of owner-occupied single-family units. This will enhance the affordable housing options available to disadvantaged households in better neighborhoods. Policies can also be implemented to further this expansion. The approach most frequently recommended is inclusionary zoning, where developers of multi-family housing projects are required to set aside a percentage of their units at below market prices. However, there is the risk that the penalty imposed on developers will kill the entire project. Another concern is the possible opposition of residents within better neighborhoods to any type of affordable housing. NIMBYism is the reality and local governments have offered little resistance. We have offered some encouraging results from our previous research where we find that the imposition of impact fees by suburban jurisdictions increases the quantity of multifamily housing receiving project approval (Burge and Ihlanfeldt, 2006). In essence the fees represent compensation to the community to offset the perceived negative externalities emitted by the housing, including the fiscal deficit that it allegedly creates. Perhaps, similar compensation at the neighborhood level would also raise the share of rental units.

Table 1 Estimated effects of housing tenure on racial and income composition in all neighborhoods

	Percent Black	Percent Poor	Percent Very Poor
SF	0.049	0.584***	0.121
	(0.105)	(0.100)	(0.081)
	[0.155]	[1.827]	[0.377]
APT	<u>0.095</u>***	0.469***	0.284***
	(0.035)	(0.029)	(0.027)
	[0.471]	[2.318]	[1.405]
CONDO	0.068*	0.353***	0.169***
	(0.038)	(0.049)	(0.037)
	[0.355]	[1.834]	[0.877]
CONDO OO	0.001	0.206***	0.102***
	(0.025)	(0.031)	(0.024)
	[0.005]	[1.469]	[0.731]
REO	4.116	2.103**	1.174
	(2.563)	(1.052)	(0.813)
	[0.813]	[0.415]	[0.232]
Mean of Dep. Var.	17.829	54.420	29.532
Tracts	498	498	498
Neighborhoods	1,534	1,534	1,534
Observations	10,688	10,686	10,686

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) Each equation also contains year dummies, time-means of exogeneous explanatory variables at the tract level, and control variables for endogenous explanatory variables (estimates not reported).
(ii) An underline indicates that the variable is treated as endogenous.
(iii) Standard errors are in parentheses. They are clustered at the tract level if all explanatory variables are treated as exogeneous; otherwise they are computed based on 1,000 bootstrap replications. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.
(iv) The figure in square bracket is the change in the dependent variable induced by a within standard deviation change in an explanatory variable (holding all other variables constant).

Table 2 Estimated effects of housing tenure on racial and income composition in neighborhoods where black people, poor households, and very poor households are underrepresented, respectively

	Percent Black	Percent Poor	Percent Very Poor
SF	0.013	0.495***	0.161**
	(0.042)	(0.104)	(0.066)
	[0.036]	[1.828]	[0.591]
APT	0.019*	0.308***	0.152***
	(0.010)	(0.049)	(0.029)
	[0.093]	[1.318]	[0.724]
CONDO	0.042***	0.307***	0.177***
	(0.014)	(0.051)	(0.031)
	[0.224]	[1.815]	[1.031]
CONDO OO	0.004	0.133***	0.065***
	(0.011)	(0.032)	(0.023)
	[0.027]	[1.263]	[0.589]
REO	1.378	1.627	1.257
	(1.030)	(1.312)	(1.028)
	[0.201]	[0.273]	[0.216]
Mean of Dep. Var.	3.277	37.900	18.738
Tracts	398	354	395
Neighborhoods	1,120	788	883
Observations	7,824	5,506	6,167

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) Each equation also contains year dummies and time-means of explanatory variables at the tract level. All explanatory variables are treated as exogeneous.
(ii) Standard errors clustered at the tract level are in parentheses. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.
(iii) The figure in square bracket is the change in the dependent variable induced by a within standard deviation change in an explanatory variable (holding all other variables constant).

Table 3 Estimated effects of housing tenure on racial and income composition in white suburban jurisdictions

	Percent Black	Percent Poor	Percent Very Poor
SF	0.069	0.508***	-0.078
	(0.066)	(0.188)	(0.240)
	[0.164]	[1.182]	[-0.182]
APT	0.045***	0.400***	<u>0.188</u>***
	(0.017)	(0.060)	(0.059)
	[0.213]	[1.875]	[0.883]
CONDO	0.029	0.209**	<u>-0.158</u>
	(0.018)	(0.082)	(0.098)
	[0.121]	[0.880]	[-0.665]
CONDO OO	<u>0.019</u>	0.341***	0.251***
	(0.022)	(0.075)	(0.065)
	[0.071]	[1.287]	[0.948]
REO	1.548*	<u>-21.825</u>*	<u>-20.334</u>
	(0.850)	(11.410)	(13.050)
	[0.181]	[-2.553]	[-2.379]
Mean of Dep. Var.	2.123	49.500	27.004
Tracts	148	148	148
Neighborhoods	408	408	408
Observations	2,831	2,829	2,829

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) Each equation also contains year dummies, time-means of exogeneous explanatory variables at the tract level, and control variables for endogenous explanatory variables (estimates not reported).
(ii) An underline indicates that the variable is treated as endogenous.
(iii) Standard errors are in parentheses. They are clustered at the tract level if all explanatory variables are treated as exogeneous; otherwise they are computed based on 1,000 bootstrap replications. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.
(iv) The figure in square bracket is the change in the dependent variable induced by a within standard deviation change in an explanatory variable (holding all other variables constant).

Table 4 Percentage change in non-Hispanic black share from a one percentage point increase in a housing type share

	(1) Black Effect	(2) White Effect	(3) Hispanic Effect	(4) Percentage Black
Panel A. All neighborhoods				
SF	+ 1.88	+ .87	0	+, I
APT	+ 1.63	+ .17	- .65	+, S
CONDO	+ 1.79	+ .08	- .65	+, S
CONDO OO	- .02	+ .14	- .27	+, I
REO	+ 104.38	- 18.57	- 51.31	+, I
Panel B. Neighborhoods where blacks are underrepresented				
SF	+ 1.90	+ 3.40	+ 2.53	+, I
APT	+ 1.34	+ .15	- .24	+, S
CONDO	+ 2.06	0	- .23	+, S
CONDO OO	+ .27	+ .10	+ .16	+, I
REO	+ 47.38	- 26.41	- 42.48	+, I
Panel C. White suburban jurisdictions				
SF	+ 4.89	+ .66	+ 1.01	+, I
APT	+ 3.16	0	0	+, S
CONDO	+ 1.95	0	+ .36	+, I
CONDO OO	+ 2.26	0	0	+, I
REO	+ 78.3	- 12.68	+ 1.71	+, S

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) A plus sign in Column (1) indicates black neighborhood percentage increasing from black entry. A negative sign in Column (1) indicates black neighborhood percentage decreasing from black exit.

(ii) A plus sign in Columns (2) and (3) indicates black neighborhood percentage increasing from white (Hispanic) exit. A negative sign in Columns (2) and (3) indicates black neighborhood percentage decreasing from white (Hispanic) entry.

(iii) Column (4) indicates the sign of the estimated effect of the housing share and whether the effect is statistically significant (S) or statistically insignificant (I) in the racial composition model.

Table 5 Percentage change in poor share from a one percentage point increase in a housing type share

	(1) Poor Effect	(2) Non-Poor Effect	(3) Percentage Poor
Panel A. All neighborhoods			
SF	0	+ .02	+, S
APT	+ 1.01	0	+, S
CONDO	+ 1.28	-.10	+, S
CONDO OO	+ .44	0	+, S
REO	+ 51.11	+ 3.27	+, S
Panel B. Neighborhoods where poor households are underrepresented			
SF	- 2.36	+ 6.33	+, S
APT	+ .20	+ .17	+, S
CONDO	+ .37	+ .88	+, S
CONDO OO	- .37	+ .82	+, S
REO	+ .57	- 37.15	+, I
Panel C. White suburban jurisdictions			
SF	- 1.63	+ 1.41	+, S
APT	+ .17	+ .18	+, S
CONDO	- .60	+ .25	+, S
CONDO OO	+ .60	+ .09	+, S
REO	+ .35	- 6.87	-, S

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) A plus sign in Column (1) indicates poor neighborhood percentage increasing from poor entry. A negative sign in Column (1) indicates poor neighborhood percentage decreasing from poor exit.
(ii) A plus sign in Column (2) indicates poor neighborhood percentage increasing from non-poor exit. A negative sign in Column (2) indicates poor neighborhood percentage decreasing from non-poor entry.
(iii) Column (3) indicates the sign of the estimated effect of the housing share and whether the effect is statistically significant (S) or statistically insignificant (I) in the income composition model.

Table 6 Percentage change in very poor share from a one percentage point increase in a housing type share

	(1) Very Poor Effect	(2) Non-Poor Effect	(3) Percentage Very Poor
Panel A. All neighborhoods			
SF	- .27	0	+, I
APT	+ 1.99	0	+, S
CONDO	+ 2.58	- .07	+, S
CONDO OO	+ .93	0	+, S
REO	+ 160.6	+ 2.49	+, I
Panel B. Neighborhoods where very poor households are underrepresented			
SF	+ .49	+ 4.46	+, S
APT	+ .75	+ 1.42	+, S
CONDO	+ 1.82	+ 1.59	+, S
CONDO OO	+ .22	+ .67	+, S
REO	- 6.40	+ 4.33	+, I
Panel C. White suburban jurisdictions			
SF	- 3.03	+ 1.24	-, I
APT	+ .33	+ .27	+, S
CONDO	- 1.06	+ .33	-, I
CONDO OO	+ 1.20	0	+, S
REO	+ 1.18	+ 5.39	-, I

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) A plus sign in Column (1) indicates very poor neighborhood percentage increasing from very poor entry. A negative sign in Column (1) indicates very poor neighborhood percentage decreasing from very poor exit.
(ii) A plus sign in Column (2) indicates very poor neighborhood percentage increasing from non-poor exit. A negative sign in Column (2) indicates very poor neighborhood percentage decreasing from non-poor entry.
(iii) Column (3) indicates the sign of the estimated effect of the housing share and whether the effect is statistically significant (S) or statistically insignificant (I) in the income composition model.

A. Appendix

Table A.1 Means and standard deviations of key variables

		Standard Deviation		
Variable	Mean	Overall	Between	Within
All neighborhoods				
Percent Black	17.829	28.346	27.953	5.046
Percent Very Poor	29.532	18.647	17.401	6.888
Percent Poor	54.420	22.921	21.823	7.232
SF	9.952	9.545	9.024	3.137
APT	23.613	28.926	28.554	4.945
CONDO	15.550	21.159	20.682	5.193
CONDO OO	14.552	20.125	18.826	7.139
REO	0.255	0.385	0.331	0.197
SF OO	36.078	31.938	31.098	7.375
Neighborhoods where blacks are underrepresented				
Percent Black	3.277	6.021	5.234	2.998
Percent Very Poor	26.636	17.445	16.137	6.660
Percent Poor	50.416	22.758	21.606	7.219
SF	8.399	8.519	8.020	2.878
APT	20.298	27.581	27.148	4.845
CONDO	18.462	22.556	22.004	5.290
CONDO OO	17.256	21.056	19.700	7.464
REO	0.194	0.289	0.249	0.146
SF OO	35.390	33.131	32.264	7.579
Neighborhoods where poor households are underrepresented				
Percent Black	11.437	21.648	21.172	4.490
Percent Very Poor	17.531	11.003	9.211	6.022
Percent Poor	37.900	16.160	14.330	7.477
SF	10.858	8.593	7.764	3.696
APT	8.609	15.973	15.390	4.273
CONDO	17.003	23.138	22.420	5.904
CONDO OO	16.157	21.366	19.172	9.481
REO	0.277	0.341	0.297	0.168
SF OO	47.096	33.363	31.937	9.767
Neighborhoods where very poor households are underrepresented				
Percent Black	12.362	23.052	22.613	4.643
Percent Very Poor	18.738	11.883	10.179	6.165
Percent Poor	41.844	18.814	17.322	7.421
SF	10.700	9.134	8.367	3.681

APT	11.044	19.084	18.488	4.775
CONDO	16.925	22.750	22.069	5.816
CONDO OO	16.249	21.460	19.454	9.097
REO	0.271	0.350	0.305	0.172
SF OO	44.811	33.399	32.089	9.369
White suburban jurisdictions				
Percent Black	2.123	4.828	4.033	2.760
Percent Very Poor	27.004	17.550	16.577	6.486
Percent Poor	49.500	23.945	23.101	6.841
SF	7.022	8.844	8.519	2.365
APT	20.367	24.884	24.637	4.708
CONDO	24.975	25.708	25.639	4.218
CONDO OO	17.472	18.553	18.187	3.775
REO	0.125	0.198	0.159	0.117
SF OO	30.039	33.216	33.030	3.588

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium
SF OO = share of owner-occupied single-family rentals

B. Appendix

Table B.1 Estimated effects of housing tenure on migration by race in all neighborhoods

	Black	White	Hispanic
SF	<u>0.023</u>	<u>-0.055</u>**	-0.001
	(0.034)	(0.023)	(0.020)
	[6.366]	[-13.927]	[-0.560]
APT	0.020	-0.010	<u>0.010</u>
	(0.013)	(0.007)	(0.008)
	[5.620]	[-2.633]	[10.897]
CONDO	<u>0.022</u>	<u>-0.004</u>	0.010
	(0.023)	(0.012)	(0.012)
	[6.088]	[-1.112]	[10.951]
CONDO OO	<u>-0.0003</u>	<u>-0.008</u>***	0.004
	(0.007)	(0.003)	(0.004)
	[-0.096]	[-2.129]	[4.723]
REO	<u>1.611</u>	<u>1.489</u>	<u>1.626</u>
	(4.221)	(1.916)	(3.517)
	[455.648]	[376.060]	[1737.267]
R-squared	0.186	0.205	0.143
Tracts	498	498	498
Neighborhoods	1,534	1,534	1,534
Observations	10,688	10,688	10,688

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) Each equation also contains year dummies, time-means of exogenous explanatory variables at the tract level, and control variables for endogenous explanatory variables (estimates not reported).
(ii) An underline indicates that the variable is treated as endogenous.
(iii) The numbers in parentheses are standard errors based on 1,000 bootstrap replications. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.
(iv) The average marginal effects are in square brackets.

Table B.2 Estimated effects of housing tenure on migration by race in neighborhoods where blacks are underrepresented

	Black	White	Hispanic
SF	0.020*	<u>-0.089</u>	<u>-0.036</u>
	(0.012)	(0.170)	(0.126)
	[1.081]	[-26.067]	[-45.235]
APT	0.014***	-0.014***	0.002
	(0.004)	(0.005)	(0.011)
	[0.782]	[-4.007]	[2.464]
CONDO	0.022***	-0.009	<u>0.002</u>
	(0.007)	(0.007)	(0.025)
	[1.185]	[-2.686]	[2.397]
CONDO OO	<u>0.004</u>	<u>-0.011</u>***	-0.003
	(0.007)	(0.004)	(0.006)
	[0.207]	[-3.121]	[-4.039]
REO	0.500	<u>1.946</u>	<u>0.934</u>
	(0.369)	(9.408)	(9.049)
	[27.101]	[572.303]	[1169.058]
R-squared	0.120	0.201	0.152
Tracts	398	398	398
Neighborhoods	1,120	1,120	1,120
Observations	7,824	7,824	7,824

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) Each equation also contains year dummies, time-means of exogenous explanatory variables at the tract level, and control variables for endogenous explanatory variables (estimates not reported).
(ii) An underline indicates that the variable is treated as endogenous.
(iii) The numbers in parentheses are standard errors based on 1,000 bootstrap replications. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.
(iv) The average marginal effects are in square brackets.

Table B.3 Estimated effects of housing tenure on migration by race in white suburban jurisdictions

	Black	White	Hispanic
SF	0.048*	-0.021	-0.011
	(0.026)	(0.023)	(0.015)
	[1.349]	[-7.525]	[-12.690]
APT	0.030***	-0.001	0.002
	(0.008)	(0.005)	(0.003)
	[0.855]	[-0.492]	[2.599]
CONDO	0.018**	-0.0001	-0.003
	(0.009)	(0.005)	(0.004)
	[0.509]	[-0.029]	[-2.895]
CONDO OO	0.021*	-0.001	0.002
	(0.011)	(0.005)	(0.003)
	[0.599]	[-0.242]	[1.699]
REO	0.803*	0.607	-0.021
	(0.410)	(0.735)	(0.152)
	[22.633]	[220.928]	[-23.080]
R-squared	0.121	0.219	0.310
Tracts	148	148	148
Neighborhoods	408	408	408
Observations	2,831	2,831	2,831

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) Each equation also contains year dummies, time-means of exogenous explanatory variables at the tract level, and control variables for endogenous explanatory variables (estimates not reported).
(ii) An underline indicates that the variable is treated as endogenous.
(iii) The numbers in parentheses are standard errors based on 1,000 bootstrap replications. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.
(iv) The average marginal effects are in square brackets.

Table B.4 Estimated effects of housing tenure on migration by income in all neighborhoods

	Poor	Very Poor	Rich
SF	0.003	-0.003	-0.001
	(0.008)	(0.017)	(0.006)
	[0.833]	[-0.550]	[-0.156]
APT	0.023	<u>0.026</u>**	<u>-0.002</u>
	(0.017)	(0.013)	(0.002)
	[6.912]	[4.198]	[-0.538]
CONDO	<u>0.029</u>	<u>0.034</u>	<u>0.002</u>
	(0.027)	(0.028)	(0.004)
	[8.778]	[5.453]	[0.488]
CONDO OO	0.010*	0.012**	-0.0001
	(0.006)	(0.005)	(0.001)
	[2.956]	[1.960]	[-0.016]
REO	<u>3.200</u>	<u>4.023</u>	-0.071
	(4.082)	(5.095)	(0.054)
	[953.093]	[649.658]	[-16.894]
R-squared	0.270	0.303	0.275
Tracts	498	498	498
Neighborhoods	1,534	1,534	1,534
Observations	10,688	10,688	10,688

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) Each equation also contains year dummies, time-means of exogenous explanatory variables at the tract level, and control variables for endogenous explanatory variables (estimates not reported).
(ii) An underline indicates that the variable is treated as endogenous.
(iii) The numbers in parentheses are standard errors based on 1,000 bootstrap replications. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.
(iv) The average marginal effects are in square brackets.

Table B.5 Estimated effects of housing tenure on migration by income in neighborhoods where poor or very poor households are underrepresented

	Neighborhoods where the Poor are Underrepresented		Neighborhoods where the Very Poor are Underrepresented	
	Poor	Rich	Very Poor	Rich
SF	<u>-0.038</u>	<u>-0.097</u>	0.006	<u>-0.091</u>**
	(0.050)	(0.065)	(0.009)	(0.038)
	[-7.711]	[-30.866]	[0.661]	[-27.844]
APT	0.004	<u>-0.007</u>	0.009***	<u>-0.014</u>**
	(0.009)	(0.015)	(0.003)	(0.007)
	[0.711]	[-2.352]	[0.967]	[-4.269]
CONDO	<u>0.006</u>	-0.014	<u>0.013</u>**	-0.018**
	(0.014)	(0.021)	(0.006)	(0.009)
	[1.220]	[-4.524]	[1.382]	[-5.409]
CONDO OO	<u>-0.006</u>	-0.013	0.003	-0.015**
	(0.006)	(0.011)	(0.002)	(0.006)
	[-1.177]	[-4.242]	[0.310]	[-4.456]
REO	0.009	<u>0.965</u>	-0.077	-0.088
	(0.089)	(2.321)	(0.066)	(0.074)
	[1.884]	[307.728]	[-8.193]	[-27.039]
R-squared	0.152	0.098	0.195	0.123
Tracts	354	354	395	395
Neighborhoods	788	788	883	883
Observations	5,506	5,506	6,167	6,167

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) Each equation also contains year dummies, time-means of exogenous explanatory variables at the tract level, and control variables for endogenous explanatory variables (estimates not reported).
(ii) An underline indicates that the variable is treated as endogenous.
(iii) The numbers in parentheses are standard errors based on 1,000 bootstrap replications. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.
(iv) The average marginal effects are in square brackets.

Table B.6 Estimated effects of housing tenure on migration by income in white suburban jurisdictions

	Poor	Very Poor	Rich
SF	-0.034**	-0.040**	-0.029**
	(0.015)	(0.018)	(0.013)
	[-9.896]	[-6.474]	[-7.837]
APT	0.003	0.003	-0.004
	(0.004)	(0.005)	(0.003)
	[0.954]	[0.412]	[-0.964]
CONDO	-0.013	-0.015	-0.005
	(0.008)	(0.010)	(0.004)
	[-3.678]	[-2.470]	[-1.425]
CONDO OO	0.013***	0.014***	-0.002
	(0.004)	(0.004)	(0.003)
	[3.745]	[2.207]	[-0.534]
REO	-0.105	0.013	-0.137
	(0.196)	(0.236)	(0.118)
	[-30.780]	[2.165]	[-36.399]
R-squared	0.361	0.348	0.144
Tracts	148	148	148
Neighborhoods	408	408	408
Observations	2,831	2,831	2,831

Notes:
SF = single-family rental share
APT = apartment rental share
REO = completed foreclosure share
CONDO = condominium rental share
CONDO OO = share of owner-occupied condominium

(i) Each equation also contains year dummies, time-means of exogenous explanatory variables at the tract level, and control variables for endogenous explanatory variables (estimates not reported).
(ii) An underline indicates that the variable is treated as endogenous.
(iii) The numbers in parentheses are standard errors based on 1,000 bootstrap replications. *, **, *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.
(iv) The average marginal effects are in square brackets.

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