NBER WORKING PAPER SERIES

DEMOGRAPHY AND POPULATION LOSS FROM CENTRAL CITIES, 1950-2000

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Working Paper 16435
http://www.nber.org/papers/w16435

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
October 2010

We thank Pascaline Dupas and Fran Kobrin Goldscheider for motivating conversations. Adriana Lleras-Muney, Stewart Tolnay and participants at the 2010 PAA meetings and the KALER group at UCLA provided helpful suggestions. The views expressed herein are those of the authors and do not necessarily reflect the views of the National Bureau of Economic Research.

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NBER Working Paper No. 16435
October 2010, Revised September 2011
JEL No. J11,N92,R23

ABSTRACT

The share of metropolitan residents living in central cities declined dramatically from 1950 to 2000. We argue that cities would have lost even further ground if not for demographic trends such as renewed immigration, delayed child bearing, and a decline in the share of households headed by veterans. We provide causal estimates of the effect of children on residential location using the birth of twins. The effect of veteran status is identified from a discontinuity in the probability of military service during and after the mass mobilization for World War II. Our results suggest that these changes in demographic composition were strong enough to bolster city population but not to fully counteract socio-economic factors favoring suburban growth.

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I. Introduction

The share of the metropolitan population in the United States living in a central city fell from 58 percent in 1950 to 36 percent in 2000, with the balance residing in the suburban ring. Suburbanization intensified residential segregation by race and income, hastened the contraction of the urban tax base, and augmented disparities in access to education and other locally-provided public services (Baumol, 1967; Benabou, 1996; Fischer, et al., 2004).

Many economic, political and sociological trends contributed to the rapid growth of the suburbs. These factors include rising real incomes among American households after World War II (Margo, 1992); road building programs that reduced the time cost of commuting from bedroom communities to the central city (Baum-Snow, 2007); federal subsidies for the purchase of single-family homes through the underwriting of mortgages and the mortgage interest deduction (Jackson, 1985); the relocation of employment opportunities to the suburban ring (Boustan and Margo, 2010); and changes in the perceived benefits of urban residence due to racial diversity, income disparities between cities and suburbs, and heightened crime rates (Cullen and Levitt, 1998; Boustan, 2007, 2010).

Alongside these forces favoring suburbanization, a series of countervailing changes in the demographic composition of the metropolitan population bolstered the size of central cities. This paper identifies four such shifts: the growing share of the metropolitan population living in a household with a foreign-born or African-American household head; the declining share in a household headed by a veteran of the Armed Forces; and the declining share of households containing a child under the age of 18. We also consider the life-cycle mobility of the large Baby

Boom cohort from city to suburb (and back again) but find that it did not have a quantitatively meaningful effect on residential patterns.[1]

Central to our argument is the claim that demographic characteristics help to determine residential location.[2] However, the causal relationship could go in the other direct as well, with residential location influencing mutable characteristics like family size and veteran status. We therefore employ instrumental variables to identify the causal effect of having an additional child or serving in the military on place of residence. In particular, we instrument for household size with the occurrence of twins on either the first or the second birth (Angrist and Evans, 1998). We identify the effect of military service by comparing cohorts who came of age during and just after mass mobilization for World War II (Bound and Turner, 2002; Page, 2008; Fetter, 2010).

In the final section, we use the estimated determinants of living in the central city to consider a series of demographic counterfactuals. Overall, we find that, absent these changes in demographic composition, the share of the metropolitan population living in the central city would have declined by an additional 10 to 32 percent from 1960 to 2000. However, these demographic changes were only strong enough to partially reverse but not to overcome the strong economic and social forces in favor of suburbanization.

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[1] A series of recent essays have pointed to the effect of demographic shifts on city growth (Ehrenhalt, 2008; Leinberger, 2008). Ehrenhalt (2008) lists the "increased propensity to remain single, the rise of cohabitation, the much later age at first marriage for those who do marry, the smaller size of families for those who have children, and at the other end, the rapidly growing number of healthy and active adults in their sixties, seventies and eighties."

[2] The relationship between demographic characteristics and residential location derive from a complex interaction between household preferences and institutional and social constraints. For example, veterans bought new homes in the suburbs not only because of their own preference for suburban residence but also because new housing construction after World War II was disproportionately located in the suburban ring. In documenting these relationships, we do not aim to distinguish between these demand-side and supply-side mechanisms.

II. Residential mobility and the decline of central city population

A. Trends in city and suburban population, 1940-2000

Figure 1a documents trends in city and suburban growth from 1940 to 2000 for the 103 metropolitan areas anchored by the largest central cities in 1940.[3] Over the second half of the twentieth century, the share of metropolitan residents living in the central city fell from 58 percent to 36 percent.[4] With the exception of the 1970s, cities experienced positive population growth in each decade.[5] However, the suburban population grew at a substantially higher rate throughout the period, leading to a steady decline in the share of the metropolitan population living in central cities. Suburban growth was driven by a combination of city-to-suburban migration, natural increase and in-migration, both from other countries and from non-metropolitan areas. The difference between city and suburban growth rates in each decade reflect net out-flows from cities to suburbs as well as differences in the rates of in-migration and natural increase between cities and suburbs.

According to Figure 1a, the popular conception of a recent urban revival is unfounded; indeed, cities continued to grow at a slower pace than the surrounding suburbs from 1990 to 2000.[6] However, many downtown areas, which historically have been industrial or commercial centers, did experience rapid population growth in the 1990s. Birch (2005) reports that the

[3] This sample includes all metropolitan areas anchored by a city that had at least 50,000 residents in either 1940 or 1970. We note that the core analysis relies instead on a varying number of metropolitan areas whose residents can be identified by location (city or suburb) in the Census micro data in each year. See section IIIa for a discussion on the construction of the main sample and Table 2 for results restricted to a consistent set of 109 metropolitan areas that can be identified in 1980 and 2000.

[4] The share of the total population living in a central city only declined from 31 percent in 1950 to 25 percent in 2000 because the metropolitan shift to the suburbs was partially offset by rural-to-urban migration.

[5] The growth of central cities in this figure is in part driven by the expansion of city land area via annexation. In 1940, the average city in this sample was 48 square miles and by 2000 it had grown to 117 square miles.

[6] We are not the first scholars to point this out. Rappaport (2003) documents that, with the exception of a few large, coastal cities such as New York, Boston and San Francisco, urban areas did not experience a reversal of fortunes in the 1990s; rather, most cities either grew continuously or declined continuously since 1950.

population in the average downtown area grew by 13 percent in the 1990s – faster than the rest of the central city or the suburban ring. However, by Birch's definition, the downtown core consists of a few Census tracts in each metropolitan area and, therefore, is not a bellwether of general urban health.

The metropolitan areas represented in Figure 1a are anchored by two very different types of cities: around two-thirds cities experienced positive population growth from 1940 to 2000 while the remaining one-third declined in size. Figures 1b and 1c display separate patterns of growth by city type. Despite differing *levels* of growth over this period, the *time pattern* of city and suburban growth is very similar across these two groups. The fastest rates of city growth were posted in the 1940s, when expanding cities grew by nearly 30 percent and declining cities experienced their last decade of positive growth. The 1970s was the low point of city growth in both categories. In the 1980s and 1990s, expanding cities again experienced positive growth and the rate of population loss slowed in declining cities. Overall, the share of the metropolitan population living in the central city declined at a similar pace in both cities types.

B. Related literature

This paper contributes to two related literatures – one on the causes of urban decline and another on the determinants of residential mobility. Studies in both of these areas emphasis the roles of race, nativity and household structure in the determination of residential location.

There is an extensive body of work documenting that African-Americans and the foreign-born are more likely than native-born, non-Hispanic whites to live in central cities (Massey and Denton, 1993; Portes and Rumbaut, 2001). Blacks moved in large numbers from the rural South to metropolitan areas from 1940 to 1970 (Gregory, 2005; Boustan, 2010). Over 80 percent of these black migrants settled in central cities. Although black suburbanization began in earnest in the 1970s, sizeable gaps in the residential locations of blacks and whites remain (Frey, 1985; Schneider and Phelan, 1993). 61.6 percent of the black metropolitan population still lived in central cities in 2000, compared to 26.1 percent of whites.

Since the passage of the Immigration and Nationality Act of 1965, a large inflow of immigrants have settled in central cities (Martin and Midgley, 2003; Singer, 2004; Frey, 2005). However, recent scholarship has emphasized that, unlike European immigrants of the early twentieth century, new immigrants groups are increasingly "bypassing central cities and settling directly in suburbs" (Alba and Logan, 1991, p. 432). Despite this trend, immigrants from every sending country are still more likely than native-born whites to live in the central city; in 1990, for example, only 33 percent of white metropolitan households lived in the city, compared to numbers ranging from 40.8 for Asian Indians to 82.9 percent for Dominicans (Alba, et al. 1999).

A portion of these location differences by race and ethnicity can be explained by group disparities in socio-economic status. In general, poor households are more likely to live in cities (Glaeser, Kahn, and Rappaport, 2008). However, notable differences in residential location by race and nativity remain even after controlling for income and education. This residual gap can be explained, in part, by the historical processes by which immigrant enclaves and majority black neighborhoods developed within central cities. To this day, some blacks and immigrants self-select into these areas to take advantage of familial or social networks or to enjoy community-specific amenities (Thernstrom and Thernstrom, 1997; Ihlanfeldt and Scafidi, 2002). In addition, African-Americans and the foreign-born continue to face barriers that preclude suburban residence, including limited access to mortgage finance (Munnell, et al., 1996; Berkovec, et al., 1996; Ondrich, Stricker and Yinger, 1999).

Household structure is another important determinant of residential location. Married couples are more likely than other household types to live in the suburbs or to move to the suburbs in a given period conditional on living in a central city (Frey and Kobrin 1982; Alba and Logan, 1991; South and Crowder, 1997). The preference among married couples for suburban living is likely related to the association between marriage and child-bearing. A large majority of married couples either currently live with children, have lived with children in the past, or are planning for children in the future. Therefore, married couples may place a higher premium on the larger lot sizes and the bundle of public goods, including higher quality public schools, available in the suburbs.[7] The presence of children in a household is itself positively related to *living in* the suburbs though, in some cases, it is negatively correlated with the likelihood of *moving to* the suburbs from elsewhere. A long literature, beginning with Rossi (1955) and Long (1972), shows that households with children are less likely to move overall and, conditional on moving, are likely to move short distances.

At midcentury, veterans of the second World War had access to housing benefits that encouraged homeownership and relocation to the suburbs. The Servicemen's Readjustment Act of 1944, commonly known as the GI Bill, included a mortgage program that allowed veterans to purchase a home with little or no down payment (Fetter, 2010). The Veterans' Administration assisted 2.1 million veterans in purchasing a home between 1946 and 1950 alone, the majority of which were located in suburban areas (Bennett, 1996, p. 24). The civilian market for credit also expanded during this period, facilitated by the Federal Housing Administration.[8] Despite the

[7] The presence of higher quality public goods and more affluent neighborhoods in the suburbs is not an inherent feature of cities in developed economies. Indeed, many European cities are organized differently, with the most desirable neighborhoods located in the city center. On the US-European comparison, see Brueckner, Thisse and Zenou (1999).
[8] The Federal Housing Administration (FHA) began insuring mortgages initiated by private lenders in the mid-1930s. As a result, mortgage rates fell from 6-8 percent in the 1920s to 2-3 percent in the 1940s and the average down payment declined from around half to around 10 percent of the value of the property (Jackson, 1985, p. 205).

expansion of credit in the civilian market, Vigdor (2006) finds that eligible veterans were still seven percentage points more likely than non-veterans to own a home in 1970.

Journalists have speculated that the aging of the baby boom generation will lead retired couples to return to cities. Demographers, however, have been more skeptical (Nelson, 1988; Frey, 1993). Frey (2007), for example, argues that seniors are more likely to "age in place." Mobility rates among the elderly are very low; less than five percent of Americans older than 65 move in a given year, compared to nearly 30 percent of individuals in their early twenties. As a result, there is little increase in the probability of living in the central city after retirement. A related literature points out that the elderly have high rates of home ownership; in 2003, 78 percent of Americans over the age of 75 owned their own home (Jones, 1997; Myers and Ryu, 2008; Painter and Lee, 2009). Contrary to the life-cycle savings model, there is no empirical evidence that seniors sell their home in order to dis-save as they enter old age. Rather, home sales among the elderly are prompted by life transitions, including the death of a spouse or a change in health status. Thus, while we investigate the effect of the aging of the baby boom cohort on city population, the demographic literature leads us to believe that this force is unlikely to be particularly strong.

III. Demographic correlates of living in the central city

A. Estimating equation

The goal of this paper is to examine whether demographic trends are quantitatively large enough to have bolstered the population of central cities, despite the strong forces encouraging suburban growth. This section begins by presenting the demographic correlates of living in the central city. The analysis is based on individual records from the 1960 to 2000 Censuses compiled by the Integrated Public Use Micro-data Series or IPUMS (Ruggles, et al., 2008).[9] Our sample includes all residents of metropolitan areas for whom place of residence (central city versus suburbs) is reported in the data. We can identify place of residence for 76 percent of the metropolitan sample in 1960, 83 percent of the sample in 1980, and 61 percent in 2000. The fraction of the population that can be identified by place of residence shifts as the Census privacy requirements change over time.[10] For robustness, we present results with a constant-geography sample below.

Our dependent variable is an indicator equal to one if the respondent lives in a central city. We pool individual records from 1960 to 2000 and estimate equation 1 using a probit specification:

$$=1 \text{ if central city}_{iact} = \alpha + \Gamma' X_i + \nu_t + \mu_a + \theta_c + \varepsilon_{iact} \quad \textbf{(1)}$$

The subscript i indexes individuals who are a years old in Census year t and belong to birth cohort c. The regression includes fixed effects for Census years (ν_t), individual years of age (μ_a), and five cohort intervals, each representing roughly twenty years of birth cohorts (θ_c).[11] X_i is a vector of characteristics for the household in which individual i resides.

[9] We do not include data from 1950 because, in that year, veteran status was only asked of individuals on the sample line (five percent of the population), preventing the accurate identification of veteran-headed households.

[10] IPUMS does not report central city status if doing so would allow users to identify geographic areas with fewer than 100,000 to 250,000 residents; the exact requirements change by year. Depending on the year, we are able to identify observations from between 91 and 143 metropolitan areas. We analyze a consistent sample of 109 metropolitan areas in Table 2. To assess whether our inability to identify observations from small metropolitan areas affects our results, we also tried reproducing Table 1 separately by city size. The only notable difference is in the estimated relationship between living in an immigrant-headed household and living in a central city. However, given that, by definition, a small share of the metropolitan population lives in a small metropolitan area, adjusting for this difference would only reduce the coefficient on being foreign-born from 0.17 to 0.14.

[11] The five cohort groups in the main specification are 1869-1910, 1911-1930, 1931-1950 and 1951-1970, with the omitted category being those born after 1970. We are able to identify age, period, and cohort effects by constraining that cohort effect to be identical within these twenty year intervals. Results are robust to instead using finer cohort groups of either 13 or 16 years.

In the baseline equation, X_i contains indicators for the race, nativity, and veteran status of the household head and a dummy variable for the presence of children in the household. We define a child as anyone who is 18 years of age or less regardless of his or her relationship to the household head. In alternative specifications, we allow residential location to vary with the number of children in the household and add indicators for being married or being an "empty nester." Households are considered to be "empty nesters" if one member reports having had children but there are no children currently present.

To the best of our knowledge, our paper is the first to estimate the age profile of city residence within birth cohorts over time. Vigdor (2006) and others report age profiles of city residence constructed from single cross-sections. These profiles likely overstate the probability that the elderly will "return" to the central city by conflating age and cohort effects. For example, individuals who were 70 years old in 1970 were born in 1900 and came of age before the diffusion of the automobile and the large-scale suburban growth of the post World War II period. Therefore, the elderly in 1970 may have been more likely to live in central cities for both life-cycle and cohort-specific reasons.

B. Probit results

Our estimating equation produces two sets of results: the age profile of city residence over the life cycle and the relationship between the other demographic characteristics in the vector X and the probability of living in the central city. We report the age profile of city residence in Figure 2, which graphs the average marginal effects by single years of age from equation 1 (plus the constant). The probability that a metropolitan resident lives in the central city peaks between the ages of 20 and 22. Many individuals then leave the central city in their late twenties and thirties. The lowest probability of city residence occurs at the age of 55, an age at which households are likely to have children and to be able to afford the larger homes available in the suburbs.[12] After that point, individuals slowly return to the city.

Table 1 presents the average marginal effects relating the other demographic characteristics to the probability of living in the central city. Members of immigrant-headed households are 17 percentage points more likely than native-born whites to live in the central city. The excess probability of living in a city is even higher for individuals living with a black household head (37 percentage points). In contrast, members of households headed by a veteran of the US Armed Forces are 4.6 percentage points *less* likely to live in a central city.[13] Consistent with recent work on immigrant locations, we find that the relationship between immigration status and residential location changes over time. In 1960, immigrant households, most of whose household heads were European born, were only 14 percentage points more likely than the native born to live in a central city. By 1990, the immigrant-native gap increased to 20 points, before declining again to 16 points in 2000 as immigrants began to suburbanize or to bypass the central city altogether.

In the first column, we estimate the effect of children on residential location with a single dummy variable for the presence of any child in the household. Households with at least one child are 7.6 percentage points less likely to live in the central city. The second column replaces the indicator variable with a linear measure of the number of children in the household. Each child appears to depress the likelihood of living in the central city by 1.2 percentage points. Together, these estimates suggest that the relationship between the presence of children and

[12] The probability of city residence is also locally minimized at age 8 when many children live in the suburbs with their parents.

[13] A portion of the relationship between veteran status and residential location appears to be driven by an association between having served in the military and being married. Controlling for marital status in column 5 cuts the effect on veteran status in half, but the coefficient is still large and statistically significant.

residential location is non-linear and, in particular, that the first few children are most strongly associated with leaving the central city. To further explore this non-linearity, the third column adds dummy variables for having exactly one child and for having two or more children in the household. Relative to households with no children, households with one child are 5.9 percentage points less likely to live in the central city and households with two or more children are 8.7 percentage points less likely to live in the central city.[14]

Households may not instantaneously adjust their residential location decisions on the basis of current composition. Rather, some households may move to the suburbs in anticipation of having children, while some households that once contained children may remain in the suburbs even after the children leave home. We provide evidence consistent with this life cycle perspective in columns 4 and 5. Column 4 uses two indicator variables to summarize household composition: one for the current presence of children and another for empty nesters who once had children living at home. The omitted category is individuals living in households that have never (to date) contained children. Compared to this omitted category, households with children present are 7.6 points and empty nesters are 1.2 points more likely to be suburban residents.[15]

Column 5 proxies for the full life-cycle effect of having children by adding an indicator for being married, relative either to never having been married or to being divorced or widowed.[16] Married individuals are 11.4 percentage points more likely than singles of the same age to live in the suburbs.[17] Because marital status is highly correlated with the presence of

[14] We experimented with adding a richer set of dummy variables and found no statistical difference between having two versus three children, three versus four children, etc.

[15] The "empty nest" indicator is only available from 1960 to 1990. In these years, the Census added a question about children ever born for all women who were at least 15 years old.

[16] We include divorced and widowed household heads in the control group because these marital transitions often lead to residential mobility. However, results are similar if we instead compare the ever and never married (results not shown).

[17] Married men have higher labor market earnings than their single counterparts, which may allow married couples to afford a suburban lifestyle (Korenman and Neumark, 1991; Ginther and Zavodny, 2001). When we control for

children, the independent effects of both the current and prior presence of children in the household decline substantially.

In the final column, we expand the sample to include non-metropolitan households. In this case, the coefficients can be interpreted as the effect of each demographic factor on the probability of living in the central city, relative to living either in a suburb or in a non-metropolitan area. The largest change in this expanded sample is the effect of veteran status on residential location. Although veterans are four percentage points more likely to live in a *suburb* relative to a city, they are only one point more likely to live in a *non-city* relative to a city. In other words, veterans are unlikely to live in non-metropolitan areas; in this specification, the lack of veterans in non-metropolitan areas offsets the concentration of veterans in the suburbs. This pattern is consistent with military exemptions for farm workers during World War II, which we discuss in more detail below.

Thus far, we have relied on the set of metropolitan areas for which residential locations (city versus suburb) is known in the Census micro data. This geography presents two concerns. First, information on place of residence is available for a varying set of metropolitan areas in the micro data in each year. Secondly, the boundaries of each metropolitan area, which is composed of one or more contiguous counties, can expand over time as the Census Bureau adds peripheral counties to existing area definitions.[18] Table 2 presents results using two samples that impose consistent geography in the 1980 and 2000 Census years.[19]

household income, the effect on marital status declines by 20 percent (results not shown). We interpret the results in column 5 as the total effect of marriage on residential location, including a potential earnings channel.

[18] For example, between 1980 and 1990 the number of counties included in the average metropolitan area increased from 4.6 to 7.5.

[19] We exclude 1960 and 1970 because of additional data restrictions in these years. In 1960, the micro data does not report metropolitan area of residence and, in 1970, either place of residence (central city versus suburb) or metropolitan area is known but not both. We then omit 1990 as well for temporal balance.

For comparison, the first column of Table 2 uses all metropolitan observations in 1980 and 2000. Metropolitan areas are made up of sets of counties, some of which are themselves large enough to be identified in the micro data. Column 2 restricts the sample to the 109 metropolitan areas for which at least one underlying county is large enough to be separately identified in the micro data in both years; in this specification, we include all counties in the 109 metropolitan area in each year even though this set of counties can change over time. The coefficients are qualitatively unchanged, suggesting that the results are not sensitive to the set of metropolitan areas included in the analysis. If anything, demographic characteristics are slightly stronger predictors of city residence in these (larger) metropolitan areas, perhaps because they have a sharper distinction between city and suburb. In column 3, we create consistent boundaries for the same set of 109 metropolitan areas by restricting our attention to the set of counties that are identified in both 1980 and 2000. That is, we exclude counties that were added to the area between 1980 and 2000 or that grew large enough over this period to be identified in the micro data only in the year 2000. Again, results are nearly identical to the full sample, suggesting that residence patterns for peripheral counties are similar to those in core suburban counties.

IV. Instrumental variables: The effect of children and veteran status on city residence

In the previous section, we estimated the effects of five demographic characteristics – age, race, nativity, veteran status and the presence of children – on residential location. Ultimately, our goal is to use these estimates to infer how changes in the demographic composition of the metropolitan population have affected the population of central cities over the past half-century. However, before doing so, we must determine that the estimates indeed reveal the effect of personal characteristics on residential location and not the other way around. In so doing, we distinguish between predetermined characteristics, such as race and nativity, and mutable characteristics like veteran status and child bearing. For example, suburbanites may be encouraged by their friends and neighbors or by the child-friendly environment to have an additional child. Furthermore, both child-bearing and veteran status may be correlated with other characteristics that are associated with living in the suburbs. As a result, we employ instrumental variables to estimate the causal effect of having an additional child or serving in the military on place of residence.

A. Veteran status

According to our probit estimates, veterans are less likely than non-veterans to live in the central city. One explanation for this pattern is that veterans were offered generous housing benefits that provided the resources necessary to buy single-family housing in the suburbs. However, this relationship could also be generated by omitted variables that are correlated with veteran status. Veteran status is determined by a combination of individual initiative (whether or not to enlist) and military selection. Men who suffer from health ailments or are cognitively impaired are less likely to serve in the military. At various points, the military also offered deferments to men who were enrolled in college, employed in a war industry or working in the agricultural sector. Any of these factors may be correlated with later residential location.

Our goal is to estimate the direct effect of veteran status on residential location while minimizing these other confounding factors. To do so, we focus on the era of mass mobilization for World War II between 1940 and 1945, a period in which the probability of military service was strongly influenced by external events. Figure 3 reports the share of white, native-born men from the 1915 to 1934 birth cohorts who served in the Armed Forces. The probability of military service increased from 50 percent for men born in 1915 to over 80 percent for the men born between 1919 and the third quarter of 1927. Men in this group were 18 to 26 years old at the end of World War II. The probability of military service then declined from 83 to 70 percentage points between the 1927 and 1928 birth cohorts. Men born after the third quarter of 1927 were too young to participate in World War II, although many of them served in Korea.

With this history in mind, we compare the veteran status and residential location of these two adjacent birth cohorts who faced different probabilities of military service due to the timing of World War II. In this setting, we treat quarter of birth as an instrument for the probability of military service.[20] In particular, in our first stage equation, the probability of military service is a function of a linear trend in quarter of birth and a dummy variable for being born before the fourth quarter of 1927. We estimate:

$$=1 \text{ if veteran}_{it} = \alpha + \beta(=1 \text{ if born before } 4^{\text{th}} \text{ Q } 1927)_{it} + \gamma(\text{quarter of birth})_{it} + \nu_t + \varepsilon_{it} \qquad \textbf{(2)}$$

The linear trend accounts for other factors that may have increased homeownership over time in the second stage equation, such as rising real incomes. The sample is restricted to white, male, native-born heads of household from the birth cohorts of 1919 through 1932. As a robustness exercise, we consider different birth years as starting and ending points of the comparison window.[21]

[20] Bound and Turner (2002) and Page (2008) use a similar approach to study the effect of the GI Bill on educational attainment. Because quarter of birth is only available in the Census from 1960 to 1980, we focus on these years in our IV estimation.

[21] For brevity, we report results in which the treated group consists of men born between 1919 and 1927 and the comparison group is either men born in 1928 and 1929 or between 1928 and 1932. Results are similar when we shorten the treatment window to 1921-1927 or 1923-1927 or lengthen the comparison window to 1935.

Table 3 presents the coefficients from our first and second stage equations, which are estimated using linear probability models.[22] As is clear in Figure 3, men born before the fourth quarter of 1927 were 13 percentage points more likely than men born after that period to have served in the Armed Forces.[23] In this restricted sample, our OLS estimates suggest that being a veteran increases the probability of living in the suburbs by 3 percentage points (a slightly smaller effect than in the full sample in Table 1). When we instrument for veteran status, the relationship between military service and suburban residence increases to 6.7 percentage points. However, we note that, given the size of the standard errors, we cannot reject that the OLS and IV coefficients are the same.

The fact that the IV estimates are larger (in absolute value) than their OLS counterparts implies that, at least during World War II, veterans were selected on attributes that were positively correlated with living in the central city. This pattern is consistent with draft exemptions for farmers and agricultural workers. In other words, we suspect that military service is correlated with living in a central city precisely because draft exemptions were most common among men living outside of cities.

This logic is consistent with a number of facts about this era. First, Acemoglu, Autor and Lyle (2004) show that, across states, mobilization rates of prime-age men varied between 40 and 55 percent, with the lowest call-up rates in the plains states of North and South Dakota and Iowa and in the agricultural South. Furthermore, in the 1950 Census, young veterans of World War II

[22] We also estimated this system using a seemingly unrelated bivariate probit model. The IV probit results are quite similar to the two-stage least squares estimates (coeff. = -0.076, s.e. = 0.030).
[23] The validity of our instrument depends on the assumption that being born after the third quarter of 1927 only affects residential location via veteran status. We check this assumption by defining an alternate birth indicator that ends in the third quarter of 1929. We know of no event in military history that would suggest that men born after this period were more or less likely to have served in the Armed Forces. Indeed, the first stage coefficient for this placebo indicator is five times smaller than the estimate for our variable of interest (coeff. = 0.024, s.e. = 0.007).

were far less likely than non-veterans to live on a farm (6.7 versus 15.7 percent).[24] Given these facts, we suspect that the larger IV coefficients provide the most reasonable estimate of the effect of veteran status on residential location. However, we note that the estimate is derived from a particular time period and, therefore, we use both the probit and IV coefficients to conduct the counterfactual exercises.

B. The presence of children

In our probit estimation, we find that households with children are less likely to live in the central city. One explanation for this result is that having children increases the demand for certain aspects of suburban life, including the larger housing units, presence of open space and higher quality public schools. However, this finding could be contaminated by either omitted variables bias or reverse causality. During this period, rich households had fewer children and were more likely to live in the suburbs, which may bias downward the relationship between having children and living in the suburbs. On the other hand, suburban residence could directly influence a household's preferences for optimal family size. The attitudes of friends and neighbors in the suburbs may encourage households to have an additional child. In this case, our estimate would overstate the effect of having children on moving to the suburbs.

We use the birth of twins as an instrument for the number of children in a household. Angrist and Evans (1998) argue that, conditional on the age and race of the mother, twinning is an exogenous event. We focus on the period 1960 to 1980 because of the availability of quarter-of-birth data used to identify twin pairs in these years. In addition, these years pre-date the development of infertility treatments that have increased the probability of twinning for (the non-random set of) mothers who seek medical intervention. A large literature uses twinning to study

[24] Authors' calculations from the 1950 IPUMS.

the effect of family size on women's and children's outcomes (Bronars and Grogger, 1994; Rosenzweig and Wolpin, 1980a, 1980b; Angrist and Evans, 1998; Black, Devereaux and Salvanes, 2005).

We define two children in the same household with the same quarter and year of birth as a pair of twins. 0.5 percent of households with at least one birth have twins on the first birth. Our first-stage equations relate the presence of twins to various measures of the number of children in the household.[25] For example, for households with at least one birth, we estimate:

$$\text{Number of children}_{it} = \alpha + \beta(=1 \text{ if twin on first birth})_{it} + \nu_t + \varepsilon_{it} \qquad \textbf{(3)}$$

Alongside the standard controls included in equation (1), we also control for the race and age of the mother. One limitation of this approach is that households must have at least one birth event in order to be at risk for having twins. Table 1 demonstrates that the shift from zero to one child is an important determinant of residential location; however, twinning cannot be used as an instrument for the presence of the first child in a household.

Table 4 presents the coefficients from our first and second stage equations, estimated using linear probability models.[26] The raw data indicates that, among households with at least one birth, those with a singleton on the first birth have an average of 2.58 children, whereas those with a twin on the first birth have 3.34 children. Accordingly, we estimate that having a set of twins on the first birth event increases household size by 0.7 children (column 2). Much of the

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[25] The validity of this instrument depends on the assumption that twinning only influences residential location through its effect on family size. We find no effect of having twins on the probability of living in the central city after controlling for the number of children in the household.

[26] We also experimented with appropriate probit IV methods. However, the resulting estimates were either implausibly large or implausibly small. In particular, the estimated effect of the number of children in the household was very large, while the coefficient on the indicator for having at least two children was very small. These two findings appear to be logically inconsistent. Therefore, we follow Angrist and Pischke (2009, p. 204-05) in reporting the more robust 2SLS coefficients.

difference in completed family size arises from the (obvious) fact that the vast majority of households with twins on the first birth have at least two children, while only 73.4 percent of households with a singleton first birth have an additional child. Consistent with this figure, we estimate that having twins increases the probability of having two children by 25.6 percentage points (column 4).

In the restricted sample of households with at least one child, the OLS estimate implies that each additional child reduces the likelihood of living in the central city by 0.5 percentage points. Note that this effect is smaller than the coefficient for the full sample reported in Table 1. The effect of family size on residential location more than doubles when we instrument for the number of children with the occurrence of twins on first birth. In this case, each additional child reduces the probability of living in the central city by 1.9 percentage points. The larger IV estimates suggest that households with many children have unobserved characteristics that are otherwise positively associated with living in the central city; for example, large households may have a lower socio-economic status. Furthermore, in this case, the OLS and IV estimates lie outside of each other's confidence intervals. However, we note that the IV results are derived from the subset of households with at least one child and may not be generalizable to the full population. We therefore conduct our counterfactual simulations using both the probit and IV estimates.

IV. Counterfactual effects of demographic composition on central city population

In this section, we document trends in the demographic composition of metropolitan area residents from 1940 to 2000. We then use a series of counterfactual exercises to assess the contribution of changes in demographic composition to the maintenance of city population over this period.

A. Demographic characteristics of the metropolitan population, 1940-2000

Figure 4 displays demographic characteristics of the metropolitan population over time. With the immigration restrictions of the mid-twentieth century, the share of the metropolitan population living in an immigrant-headed household fell from 30 percent in 1940 to ten percent in 1970. After the expansion of immigration quotas in 1965, this share returned to nearly 30 percent by 2000. Given that the foreign born are more likely to live in central cities, we expect this pattern to contribute to population growth in central cities from 1970 onward. As blacks from the rural South migrated to industrial cities, the share of the metropolitan population living with a black household head increased from eight percent in 1940 to 17 percent in 2000. Again, this pattern would likely bolster city population.

After servicemen returned from World War II, the share of the metropolitan population living in a veteran-headed household spiked from less than five percent in 1940 to nearly 50 percent in 1960 and 1970. Since 1970, the veteran share has declined to just over 10 percent in 2000. Because veterans are more likely to live in the suburbs, the recent reduction in the number of veterans in the population favors the city relative to the suburbs. The share of households with a child present has also declined substantially since mid-century. While nearly 80 percent of households had at least one child in 1950, this share declined to 50 percent in 2000. Again, because households with children are more likely to live in the suburbs, the growth of childless households favors the city.

The age structure of the population also notably shifted from 1940 to 2000 with the birth and aging of the large baby boom cohort (born between 1946 and 1964). Yet, because of the rapid swings in the age profile of city residence (Figure 2), we find that the aging of this cohort had little effect on city growth. In essence, there is no decade in which the baby boom generation has been clustered in either a peak or a valley of the city residence profile. In 1980, for example, many of the baby boomers were in their early twenties and lived in central cities. However, at the same time, others in the cohort were still in their teenage years or had entered their early thirties and therefore tended to live in the suburbs.

We explore the possibility of age effects on city growth more systematically by imposing a counterfactual flat age profile and predicting the resulting city share of the metropolitan population in each decade. In particular, we allow each age between zero and 70 to contain 1.3 percent of population and constrain older ages to each hold 0.3 percent of population. Using the estimated age effects in Figure 2, we then predict the share of the population that would be living in the central city under this counterfactual age profile. We find no meaningful difference between the predicted and the actual city shares, leading us to turn our attention to the other set of demographic factors (results available upon request).

B. Counterfactual simulations

Table 5 uses the coefficient estimates relating demographic trends to city residence to provide counterfactual statements about how much further the share of metropolitan residents living in central cities *would have declined* between 1960 and 2000 if not for these demographic moderators.[27] In the simplest exercise, we use the probit or IV coefficients to consider the extent

[27] We select 1960 as our starting point because a number of our variables reached their minimum or maximum point in that year. We discuss the sensitivity of this result to varying starting points below.

to which center city population increased through each of the four demographic channels. We then allow for the fact that new arrivals may lead to the departures of some existing residents, either directly through white or native flight or indirectly via an increase in city housing prices.

The first row of Table 5 presents the actual change in the share of metropolitan residents living in central cities from 1960 to 2000. The share of the metropolitan population living in a central city declined by 17.8 percentage points over this period. The magnitude of this decline reflects a combination of socio-economic forces favoring suburban residence, on the one hand, and the demographic trends favoring city residence on the other.

The second row of Table 5 isolates the role of changes in demographic composition. In particular, we ask how the share of the metropolitan population living in a central city would have changed if changes in demographic composition had been the only relevant factor over this period and if new arrivals did not generate corresponding departures. In this case, the city share of metropolitan population would have actually increased by 6.4 to 8.3 percentage points from 1960 to 2000. The low and high points of this range measure the strength of the relationship between demography and residential location using the probit or IV coefficients, respectively.[28]

Rows 3a through 3d illustrates how each demographic characteristic contributes to the total counterfactual change in city population. That is, the sum of the entries in rows 3a through 3d is equal to the total counterfactual change in the city population share in row 2. To generate these values, we multiply the total change in the variable in question from 1960 to 2000 by the estimated effect of that variable on the probability of living in the central city. For example, the share of households headed by an immigrant increased by 16.3 percentage points from 1960 to

[28] Because our IV regressions are estimated on selected samples, we use the ratio between the OLS and IV estimates in Tables 3 and 4 to scale the coefficients for the whole population from Table 1. Specifically, we augment the veterans coefficient by 2.3 (= -0.067/-0.029 from Table 3, columns 1 and 2) and we augment the any child coefficient by 1.6 (= -0.052/-0.032 from Table 4, columns 3 and 4).

2000 and, as a result, the share of metropolitan residents who lived in the central city rose by 2.8 percentage points (= 16.3 · 0.170, coefficient estimate from Table 1, column 1).

In the probit-based counterfactual, the most important demographic trend contributing to the growth of city population is renewed immigration, followed by the increase in black population, the shrinking number of veterans and the declining share of households with children. The IV estimates suggest a greater role for the decline in veteran status and in the share of households with children present. However, we caution that the IV estimates are less precisely estimated. Therefore, we report 95 percent confidence intervals around each of the counterfactual effects calculated from the IV estimates. Note that if we, instead, had selected 1940 or 1950 as a starting point for this exercise, the role of racial composition and the presence of children would have been larger.

The simple counterfactuals discussed thus far do not account for the possibility that some existing residents might have left central cities as new households arrived, either due to a direct distaste for living near black or immigrant neighbors or due to the indirect effect of these arrivals on city housing prices. The fourth row of Table 5 presents a modified counterfactual using estimates of white/native flight and housing price responses from the literature to calculate the *net effect* of these demographic shifts on center city population.

Recent studies of white and native flight find nearly one-for-one displacement rates from cities or urban neighborhoods. Boustan (2010) shows that, over a single decade, one southern black arrival into a central city led to 2.5 non-black departures. Over the long run, new arrivals partially compensate for initial white flight and one black arrival is associated with one non-black departure. Saiz and Wachter (2011) find that one immigrant arrival into a Census tract leads to the departure of 0.68 native, non-Hispanic whites. Borjas (2006) presents a similar estimate for the native workforce at the metropolitan level (0.61 departures).[29] Adjusting for native white departures completely offsets the effect of black arrivals and reduces the net effect of immigrant arrivals by 70 percent. In this case, the 4.4 percentage point gross increase in the share of the metropolitan population located in the central city due to black and immigrant arrivals would have only lead to a 0.9 percentage point net increase in the city share (= 2.8 points · 0.32 + 1.6 · 0.0).

The other two demographic groups – veterans and households without children – are not likely to prompt specific outflows. However, their presence could affect the probability that other households remain in (or move to) a city indirectly via the housing market.[30] A larger population in the city can increase housing prices and, in turn, higher prices may encourage existing residents to leave or deter other new residents from moving in.[31] The best empirical evidence on the effect of population growth on housing prices is based on variation in immigration across metropolitan areas.[32] Saiz (2007) finds that a one percent *net* increase in metropolitan population increases housing prices by 0.77 percent.

Assessing the effect of higher housing prices on out-migration is an empirical challenge. In the raw data, there is no correlation between housing rents and net migration at the state level (Coen-Pirani, 2010). However, this null effect is likely due to the fact that states with higher housing prices also have either a more productive set of industries (and hence higher wages) or a

[29] Similar work by Crowder, Hall and Tolnay (2011) shows that, for the native born, the log odds of moving increases with the foreign-born share of the neighborhood.

[30] The presence of non-veteran households or households without children need not be the result of new in-migration to the central city. However, these changes in demographic composition may instead lead fewer households leave the city for the suburban ring or elsewhere. As a result, the central city will have a larger population through retention rather than through new in-migration.

[31] The extent of price increases depends on the elasticity of housing supply; at the extreme, prices will not respond to in-migration if each new arrival is met with the construction of a new housing unit.

[32] An earlier literature used the coming of age of the baby boom cohort to assess the relationship between population growth and housing prices. Mankiw and Weil (1989) document large effects of the entry of baby boomers into the housing market in the US, while Engelhardt and Poterba (1991) find no effect in Canada. This evidence is hard to interpret because it is based on national time series.

more valuable set of amenities (Roback, 1982). Saks (2008) is one of the few papers to (indirectly) measure the migration response to an exogenous change in housing prices, in this case driven by variation in zoning regulations. She finds that, in places with strict zoning rules, in-migration is dampened by a corresponding increase in housing prices; her estimates imply that a one percent increase in housing prices reduces any given in-migration flow by 0.4 percent.

According to our probit estimates, the three demographic forces under consideration increased gross city population by 7.5 percent.[33] As a result, housing prices would have increased by 6 percent, leading to a subsequent 2 percent population decline. In other words, a 7.5 percent increase in *gross* city population would have resulted in only a 5.5 percent *net* increase in the number of city residents due to the effect of population growth on housing prices. Converting these percentages back into shares, these calculations imply that, net of white flight and price-induced out-migration, demographic factors would have increased the city share of the metropolitan population by 2.1 percentage points (Table 5, row 4). Rows 5a through 5d use a similar logic to illustrate how each demographic trend contributes to the net counterfactual.

Overall, we conclude that, absent these demographic shifts, central city population would have declined by 10 to 32 percent more than it did between 1960 and 2000. The range of these estimates depends on the estimation method used (probit or IV) and assumptions about how existing residents would have reacted to these new arrivals. Our preferred estimate uses the IV coefficients to calculate the effect of each demographic factor on city population, while also accounting for responsive out-migration. By this estimate, the share of metropolitan residents living in the city would have declined by an additional 3.4 percentage points, or 16 percent, from

[33] The demographic forces include a declining veteran share, an increase in households without children, and the net increase in population due to foreign arrivals. Together, these forces would have increased the share of metropolitan residents living in the central city by 3 percentage points. Over this period, around 40 percent of metropolitan residents lived in a central city. Therefore, city population would have needed to increase by 7.5 *percent* in order for the share of the metropolitan area living in a central city to increase by 3 *percentage points*.

1960 to 2000 if not for the demographic counterweight (16 percent = 3.4 points/[3.4 points + actual 17.8 points]). In other words, shifts in demographic composition helped to maintain city population but were not strong enough to compensate for the powerful forces favoring population growth in the suburbs.

V. Conclusion

The share of the metropolitan population living in central cities has declined sharply over the past sixty years. This paper shows that, absent changes in the demographic composition of the metropolitan population, this share would have fallen even further. In particular, city population has been bolstered by the in-migration of southern blacks from 1940 to 1980, the expansion of international immigration after 1965, and a decline in the share of households with children or headed by a veteran.

We provide new estimates of the relationship between each of these demographic characteristics and the likelihood of living in the central city. Our analysis distinguishes between predetermined characteristics, such as race and nativity, and endogenous characteristics like veteran status and child bearing. We instrument for veteran status by comparing birth cohorts of men coming of age during and just after the mass mobilization for World War II. We use the arrival of twins to instrument for the number of children in a household. In both cases, the IV estimates are larger in absolute value (more negative) than the corresponding probits.

Two counterfactual simulations are used to assess the effect of these demographic factors on city population from 1960 to 2000. Our simplest approach predicts the gross number of new residents in central cities using our coefficient estimates and the trends in each characteristic over time. We then allow for the fact that new arrivals may lead to the departures of some existing residents, either directly through white or native flight or indirectly via an increase in city housing prices. The counterfactuals indicate that changes in demographic composition increased the share of metropolitan population living in the central city between 10 and 32 percent. Our preferred estimate, which relies on the IV coefficients and allows for responsive out-migration, suggests that the share of the metropolitan population living in the central city would have fallen by 16 percent more than it did if not for these demographic trends. Changes in demographic composition were only strong enough to attenuate, not to reverse, a relative decline in city population driven by economic and social factors favoring suburbanization.

Through the 1990s, central cities continued to experience positive population growth but their expansion continued to be outpaced by more rapid growth in the suburbs. Although we find no evidence of a recent urban revival, this national focus may miss localized instances of gentrification in certain neighborhoods or within particular cities that have been fueled in part by demographic trends (Vigdor, 2002). For example, the interaction between demographic forces, particularly delayed child-bearing and longer life expectancies, and rising incomes in top income brackets may have contributed to the renaissance of "super star" cities like New York City and San Francisco (Gyourko, Mayer and Sinai, 2006). Understanding variation in the role of demography across different types of cities remains an important avenue for future research.

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Table 1: Demographic characteristics and the probability of living in the central city, 1960-2000

	Dependent variable = 1 if live in central city					
RHS variable	(1)	(2)	(3)	(4)	(5)	(6)
Any children in HH	-0.076			-0.079	-0.042	-0.074
	(0.001)			(0.001)	(0.001)	(0.001)
Number children		-0.012				
		(0.000)				
One child			-0.059			
			(0.001)			
2+ children			-0.087			
			(0.001)			
Empty-nester				-0.012	0.009	
				(0.001)	(0.001)	
Married					-0.114	
					(0.001)	
Head is foreign born	0.170	0.169	0.171	0.172	0.179	0.234
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Head is black	0.367	0.369	0.367	0.372	0.351	0.281
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Head is veteran	-0.046	-0.048	-0.045	-0.045	-0.018	-0.008
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)
Constant	0.564	0.519	0.506	0.567	0.629	0.391
	(0.003)	(0.003)	(0.003)	(0.004)	(0.004)	(0.003)
N	5,742,225	5,742,225	5,742,225	4,536,450	4,536,450	8,406,585
Pseudo R-squared	0.99	0.99	0.99	0.99	0.99	0.99

Notes: Coefficients from probit estimation of equation 1 with standard errors in parentheses. Regressions also contains four birth cohort dummies, four census year dummies, dummies for single years of age between 1 and 90 and an indicator for being older than 90. Household members are assigned the race, nativity and veteran status of the household head. Cox-Snell pseudo R-squared statistics are reported in the last row. In columns 1-5, the sample contains all residents of metropolitan areas for whom place of residence (central city or suburb) is known. Column 6 adds all non-metropolitan residents to the sample. Columns 4 and 5 do not include data from the year 2000 because the variable “child ever born” used to construct the “empty nester” indicator is not available in that year.

Table 2: Demographic correlates of living in the central city using constant city and metropolitan area samples, 1980 and 2000

Dependent variable = 1 if live in central city

RHS variable	All metros (1)	Consistent set of metros (2)	Consistent set of metros with consistent boundaries (3)
Any children in HH	-0.079	-0.084	-0.088
	(0.000)	(0.000)	(0.000)
Head is foreign born	0.163	0.176	0.178
	(0.000)	(0.000)	(0.000)
Head is black	0.340	0.373	0.363
	(0.000)	(0.000)	(0.000)
Head is veteran	-0.055	-0.062	-0.063
	(0.000)	(0.000)	(0.000)
Constant	0.450	0.613	0.561
	(0.003)	(0.003)	(0.004)
N	11,938,826	8,830,474	8,393,237
Pseudo R-squared	0.87	0.92	0.90

Notes: Coefficients from probit estimation of equation 1 with standard errors in parentheses. Regressions also contains four birth cohort dummies, one census year dummy and dummies for single years of age between 1 and 90 and an indicator for being older than 90. Household members are assigned the race, nativity and veteran status of the household head. Cox-Snell pseudo R-squared statistics are reported in the last row. In column 1, the sample contains all residents of metropolitan areas for whom place of residence (central city or suburb) is known. In column 2, the sample contains residents of the 109 metropolitan areas for which at least one underlying county is large enough to be separately identified in the micro data in both years. In the third column, the sample contains the same 109 metropolitan areas and imposes the 1980 metropolitan area county definitions in both 1980 and 2000.

Table 3: IV estimates of the effect of veteran status on place of residence, 1960-80

	Birth cohorts: 1919-1929		Birth cohorts: 1919-1932	
RHS variable	OLS	IV	OLS	IV
	A. First stage. Dependent variable = 1 if veteran			
=1 if born between 1919-1927		0.132		0.130
		(0.003)		(0.003)
	B. Second stage. Dependent variable = 1 if live in city			
=1 if veteran	-0.029	-0.067	-0.029	-0.069
	(0.003)	(0.029)	(0.002)	(0.028)
N	188,734	188,734	237,968	237,968
R-squared	0.02	0.02	0.03	0.03

Notes: The sample is restricted to white, native-born male heads of household for whom place of residence (central city or suburb) is known. Regressions include a linear trend in quarters of birth, an indicator for children present in household, and dummies for 1970 and 1980 Census years.

Table 4: IV estimates of the presence of children in the household on place of residence, 1960-80

	Households with at least 1 birth				Households with at least 2 births			
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	A. First stage. Dependent variable = Number children (or indicator)							
=1 if twins on 1st (2nd) birth		0.705		0.256		0.924		0.408
		(0.012)		(0.004)		(0.010)		(0.004)
	B. Second stage. Dependent variable = 1 if live in city							
Number children	-0.005	-0.019			0.000	0.004		
	(0.000)	(0.006)			0.000	(0.004)		
=1 if 2+ children			-0.032	-0.052				
			(0.001)	(0.015)				
=1 if 3+ children							-0.006	0.008
							(0.001)	(0.009)
R-squared	0.10	0.10	0.11	0.10	0.10	0.10	0.10	0.10

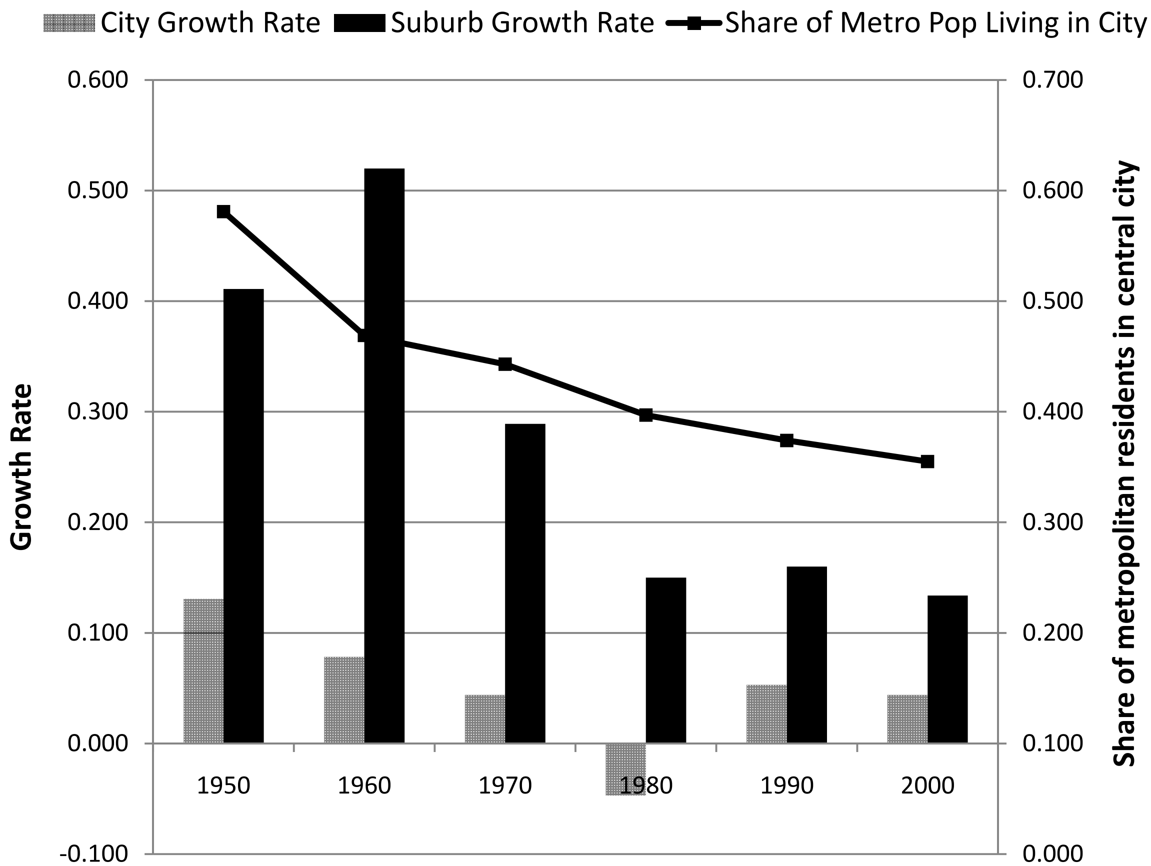
Notes: N = 2,372,595 for one birth and 1,746,963 for two births. Twins are defined as two children in the same household with the same year and quarter of birth. Sample selection and specifications follow the notes to Table 1.

Table 5. Demographic contributions to city population growth: Counterfactual scenarios

	Share of metropolitan population living in central cities			
	Level in 1960	Level in 2000	Change (Probit)	Change (IV)
1. Actual city share	51.3	33.5	-17.8	-17.8
2. Counterfactual share, gross population flows	51.3	57.7	6.4	8.3 [5.4, 11.5]
3. Contributions to gross counterfactual				
a. Foreign born	*Increased 16.3 points*		*2.8*	*2.8*
b. Black	*Increased 4.3 points*		*1.6*	*1.6*
c. Veteran	*Declined by 25.8 points*		*1.2*	*2.5* *[0.4, 5.0]*
d. Children in HH	*Declined by 11.1 points*		*0.8*	*1.4* *[0.6, 2.1]*
4. Counterfactual share, net population flows	51.3	53.4	2.1	3.4
5. Contributions to net counterfactual				
a. Foreign born			*0.6*	*0.6*
b. Black			*0.0*	*0.0*
c. Veteran			*0.8*	*1.8* *[0.3,3.5]*
d. Children in HH			*0.5*	*1.0* *[0.4, 1.5]*

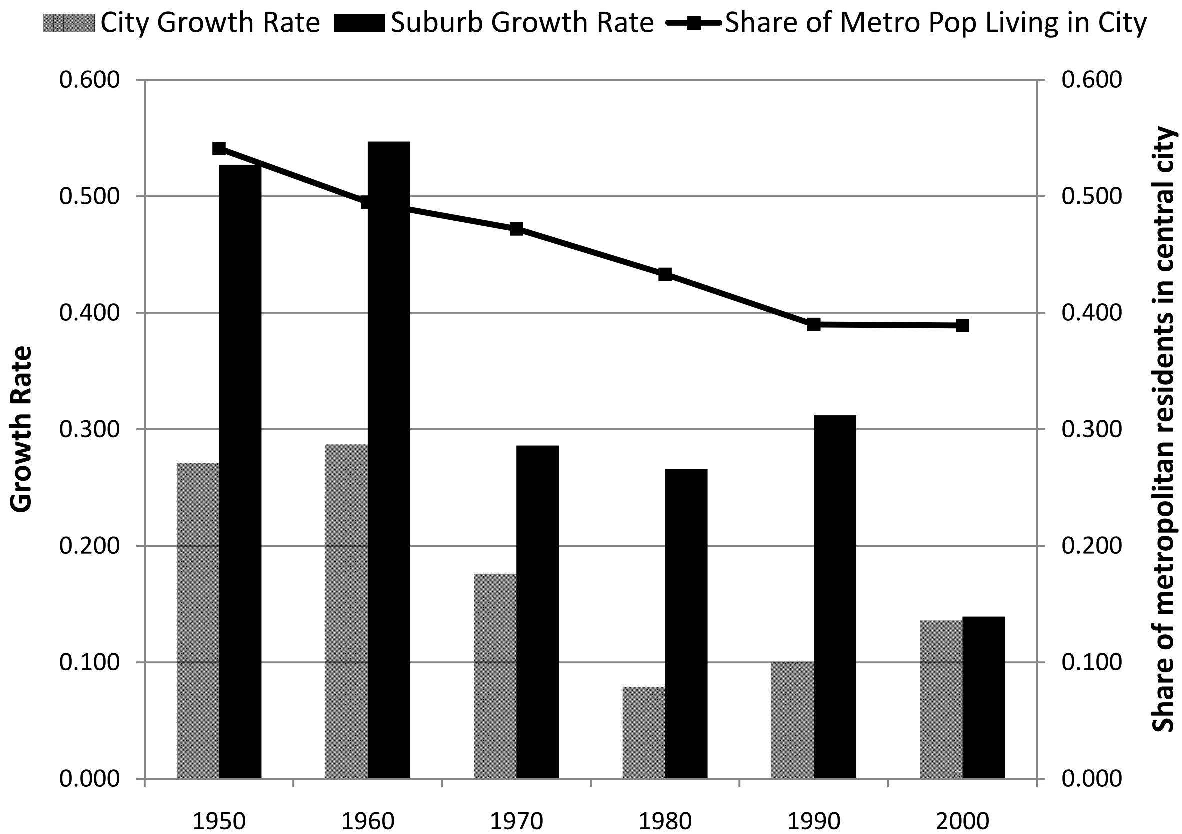
Notes: Row 1 reports the actual share of metropolitan area residents who report living the central city from IPUMS samples. Row 2 presents the counterfactual share of the metropolitan population living in central cities under a scenario in which demographic composition is the only factor allowed to change between 1960 and 2000. The counterfactual in column 3 is based on the probit regression in Table 1, while the counterfactual in column 4 is derived from the IV coefficients in Tables 3 and 4. See footnote 23 for the details on translating the IV coefficients for use in the counterfactual. For the IV-based counterfactuals, we report the 95-percent confidence interval in square brackets. Rows 3a-3d indicate the contribution of each demographic factor to the overall counterfactual in row 2. Row 4 reports the results of a modified counterfactual simulation that allows for the fact that new arrivals may lead to the departures of some existing residents, either through white/native flight or through an increase in housing prices. Rows 5a-5d indicate the contribution of each demographic factor to the net counterfactual in row 4.

Figure 1a. City and suburban population growth by decade, 1940-2000
103 metropolitan areas



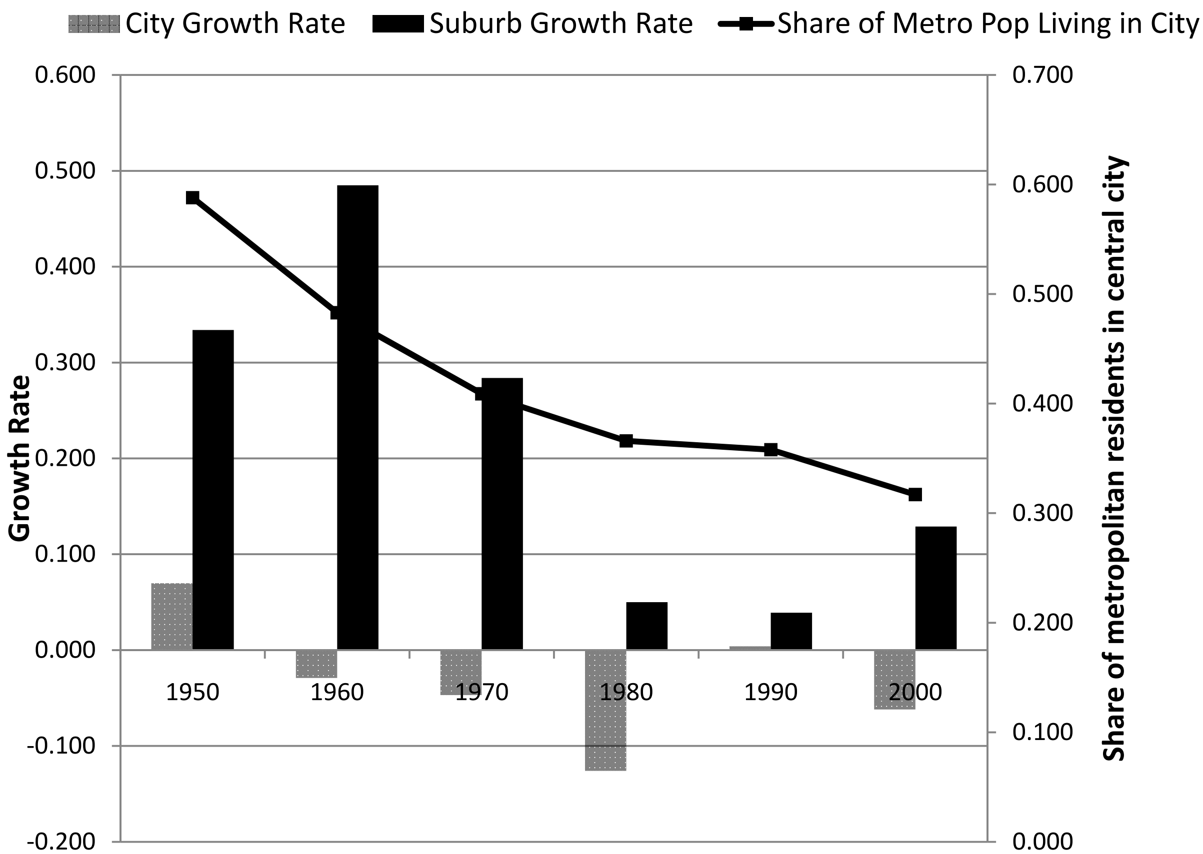
Notes: N=103 metropolitan areas. Values refer to the decade ending in the Census year on the x-axis. All metropolitan areas are anchored by a city that had at least 50,000 residents in either 1940 or 1970. City and county population are taken from the City and County Data Books. The 1970 county definitions of metropolitan areas are applied in all years. Suburban population is computed as the total metropolitan area population minus the city population.

Figure 1b. City and suburban population growth by decade:
62 metropolitan areas whose city gained population between 1940 and 2000



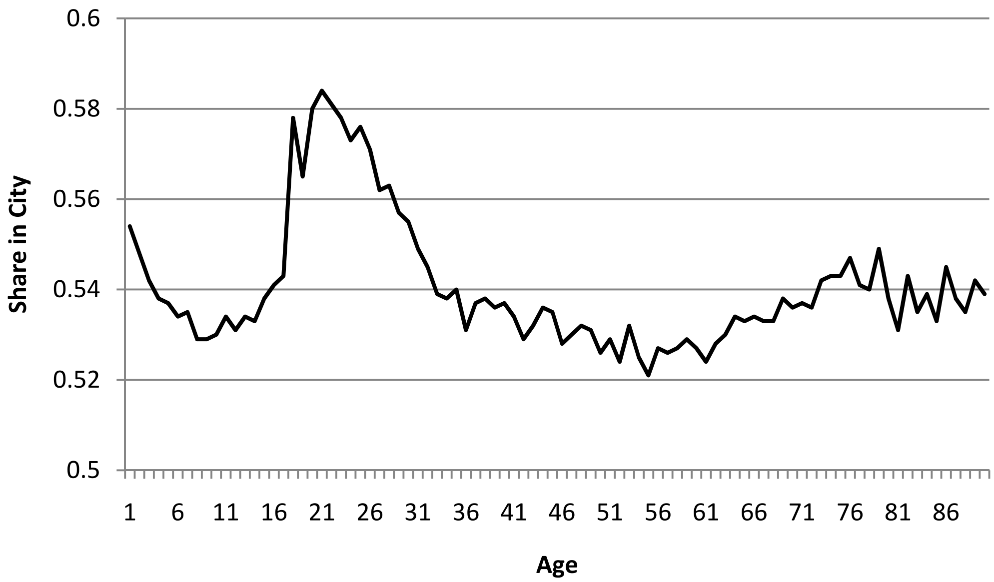
Notes: N = 62. See notes to Figure 1a.

Figure 1c. City and suburban population growth by decade
41 metropolitan areas whose city lost population between 1940 and 2000



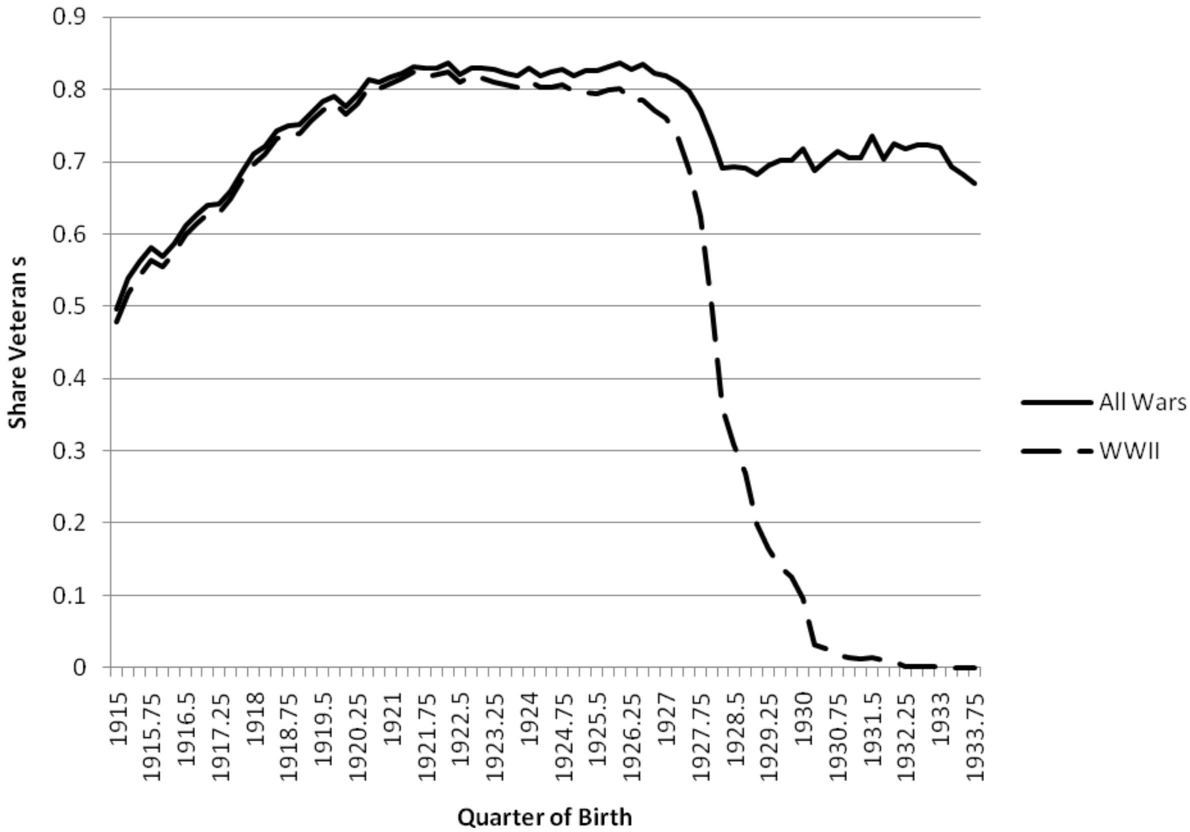
Notes: N = 41. See notes to Figure 1a.

Figure 2. Probability of living in city conditional on being in metropolitan area by age



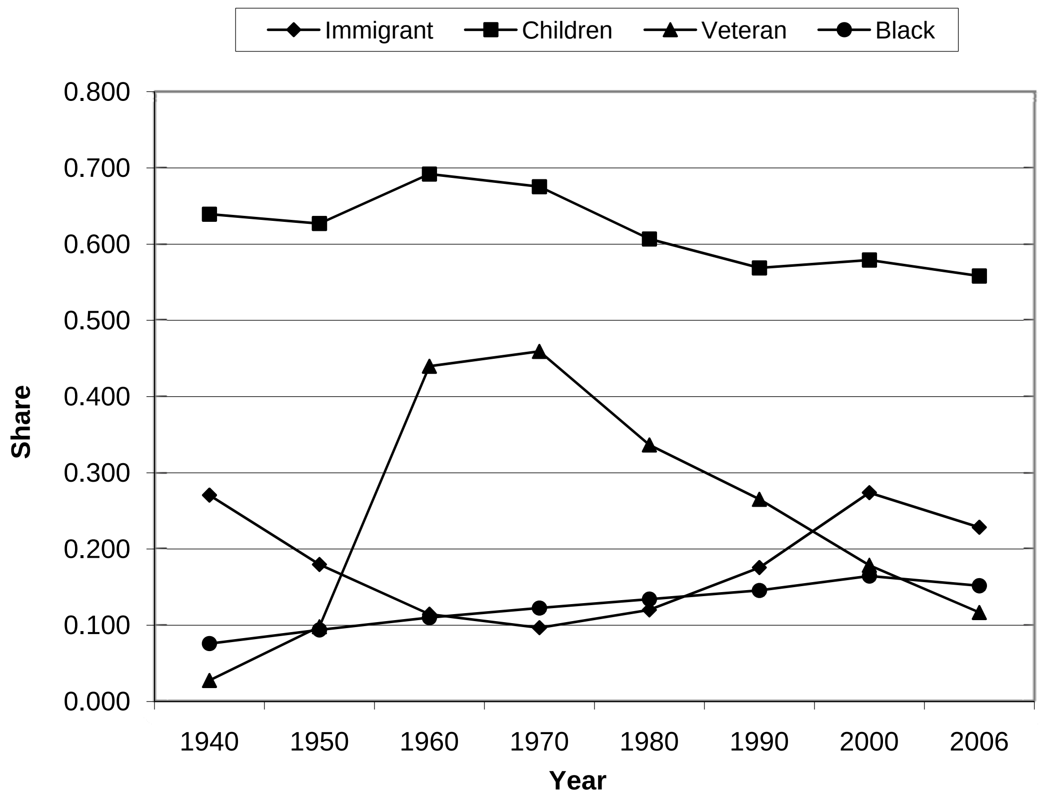
Source: IPUMS, 1960-2000.
Notes: We plot the constant plus the average marginal effects of the single years of age indicators in equation 1. The underlying regression equation also contains indicators for four birth cohorts, four census years, and controls for the presence of children in the household and the race, nativity and veteran status of the household head.

Figure 3. Share of white men serving in armed forces by year and quarter of birth



Notes: Sample includes all white, native-born men from the 1960-1980 1% IPUMS samples.

Figure 4. Characteristics of the metropolitan population, 1940-2000
Race, nativity, veteran status and presence of children



Source: IPUMS, 1940-2000.
Notes: Sample contains all metropolitan area residents for whom place of residence (central city versus suburb) is known in the given Census year. Household members are assigned the characteristics of the household head.