THE EFFECT OF RESIDENTIAL LOCATION ON RETIREMENT AGE: THEORY AND SOME EVIDENCE ON MALE BEHAVIOR

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We introduce a retirement age choice into a conventional model of residential location to show that the retirement age of workers who do not move after retirement can be influenced by the urban variation of leisure and housing prices. A retirement demand equation that incorporates commute characteristics is estimated on a sample of urban male owners interviewed by the Panel Study of Income Dynamics. We find that commuters retire around 2 years earlier than comparable home-based workers, although the evidence provides no support for this gap to be the result of differences in leisure or housing prices across both groups.

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1. INTRODUCTION

When residential and job locations do not coincide and commuting is a necessity, a worker incurs commuting costs. These costs mainly include monetary, travel time, and psychological costs (e.g., risk of accident, discomfort), and the available evidence suggests that they are substantial. Small et al. (2005), for example, estimate that the marginal costs of one hour of commuting are $26.73, whereas the figure obtained in Van Ommeren and Fosgerau (2009) is $31.15.[1] In the case of a U.S. male worker aged 55-64, whose one-way travel to work time is 27 minutes,[2] these estimates imply that weekly commuting costs are around $130 if ten trips per week are assumed.

A branch of the literature on the theory of labor supply has stressed that commuting costs (or, more generally, fixed work costs) affect the opportunity price of leisure, leading, in turn, to adjustments in the labor supply. This insight, considered when analyzing the supply of daily hours and workdays (Oi 1976) and when examining labor market entry decisions (Cogan 1981), likewise seems relevant to study retirement, as yearly commuting costs lower the opportunity price of a year retired. Commuting costs also play an important role in urban economics theory, where the price of housing is generally determined by access to work (see, e.g., Alonso 1964, Muth 1969). Furthermore, as Oi (1976) recognized, the urban variation in housing prices opens up the possibility that the demand for leisure is influenced by the worker's residential location, if leisure and housing are related in consumption.

This paper first seeks to integrate the two preceding insights into a theoretical analysis of the decision to retire. We do this by introducing a retirement age choice into a conventional model of residential location. We find that the retirement age of workers who

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[1] These monetary quantities are expressed in 2009 U.S. dollars. I thank Mogens Fosgerau and Jos Van Ommeren for clarification on the nature of their estimate.

[2] Author's calculation with data from the American Time Use Survey 2003.

do not move after retirement can be influenced by traditional income and substitution effects created by the urban variation of leisure and housing prices. Kolodziejczyk (2006) related the age of retirement to fixed work costs using an alternative argument. He assumed that fixed work costs increase the marginal utility from consumption, so that if consumption and leisure are direct substitutes (in the sense that an increase in leisure diminishes the marginal utility from consumption), higher fixed work costs will lead workers to retire earlier in order to smooth consumption. Kolodziejczyk, however, did not incorporate space into his theoretical considerations, so that neither the level of fixed costs can be chosen by the worker nor the labor supply can be influenced by the price of housing.

Our second purpose is to explore empirically how residential location affects older men's retirement age. The existence of an empirical relationship between these two variables is relevant from a policy perspective, as public authorities would be offered a new course of action to alleviate projected financial pressures associated to population aging. In addition, it is also relevant from a historical viewpoint. Over the last decades, a much investigated topic in retirement literature has been the role played by the increasing generosity of Social Security in the long-term decline of older men's labor force participation (LFP) (see, e.g., Boskin 1977, Hurd and Boskin 1984, Burtless 1986, Krueger and Pischke 1992, Gruber and Wise 1999 and 2004, French 2005, Blau and Goodstein 2010). Although there remains considerable disagreement as to the magnitude of the effect, a recent survey concluded that studies that use a more plausible identification strategy tend to find a very modest impact of Social Security wealth on labor supply (Krueger and Meyer 2002), implicitly raising the question about the reason behind the LFP decline. Figure 1 shows that, in the U.S., the downward trend in LFP observed during the 1970s and 1980s coincided with a tendency to reside farther from the workplace by older workers.[3] This fact, coupled with the theoretical possibility that, among workers with high moving costs, those living farther from the place of work retire earlier, suggest that the increasing home-workplace separation could be partly responsible for the LFP decline.

This paper contributes to a field of research comprising the intersection of labor and urban economics that is reviewed and united in Simpson and van der Veen (1992), Crampton (1999), and Zenou (2009). Section 2 develops the basic theoretical model and discusses some extensions. Section 3 describes the data and econometric methodology used in the empirical analysis. The main empirical results and a number of robustness checks are presented in Section 4. Section 5 provides a summary of the analysis and discusses certain limitations of the data.

2. RESIDENTIAL LOCATION AND RETIREMENT AGE: ESTABLISHING A CAUSAL EFFECT

Consider a circular entity located on a uniform, featureless plain, with all employment concentrated in a central business district (CBD) of negligible size, and a transportation system infinitely-elastically supplied in any direction from the CBD. A worker must choose one residential location, represented by its distance to the CBD (D), the years in retirement (L) out of his remaining years of life (T), housing services (Q), and other consumer goods

[3] Over the same period, the percentage of older workers working primarily from home decreased from 12.5% to 4.6%. This tendency to live farther is not exclusive of U.S. older workers, but part of a general shift common to OECD countries (Schafer 2000). Doling and Horsewood (2003) relate another shared trend (the expanding home ownership sector) to early retirement.

(X). Moving costs are assumed to be so high that a worker's residential location remains fixed when he retires.

Preferences are represented by the utility function

$$U = U(X, Q, L), \tag{1}$$

while the budget constraint is given by

$$X + R(D)Q + C(D)H \leq WH + Y. \tag{2}$$

The twice continuously differentiable functions $R(D)$ and $C(D)$ stand, respectively, for the market price of Q and yearly commuting costs; H is the number of years working, W yearly earnings, and Y non-labor income. Since $H = T - L$, expression (2) can be rewritten as

$$X + R(D)Q + (W - C(D))L \leq F \equiv (W - C(D))T + Y, \tag{3}$$

where F stands for the worker's potential income if working all T periods. The (opportunity) price of a year retired equals foregone earnings minus avoided commuting costs, $W - C(D)$. If, as seems sensible, commuting costs increase with D $(C' > 0)$, the price of a year retired will decrease with distance to the CBD.

Following Muth (1969), the worker's decision-making process is decomposed into two stages. In the first stage, optimal quantities of X, Q, and L are chosen for a given distance, obtaining a short-run maximum of utility. In the second stage, distance itself is chosen so as to maximize utility. The necessary conditions for equilibrium are expression (3), a set of first-order conditions for the short-run optimum, and an equation for the second-stage optimum distance,

$$-QR' = HC'. \tag{4}$$

As is well-known, equation (4) states that the worker's full income does not change in the neighborhood of the equilibrium location. With $C' > 0$, the price of housing must decrease with D $\left(R' < 0\right)$ for equation (4) to be satisfied.

There is the stability condition (Muth, 1969; De Salvo, 1977) among the sufficient conditions for equilibrium:

$$-QR'' - R'\frac{\partial Q}{\partial D} - HC'' - C'\frac{\partial H}{\partial D} \leq 0\,, \tag{5}$$

which expresses that the net savings on the quantity of housing and transportation decline with distance. Upon rearrangement, it yields

$$R'' \geq -\frac{1}{Q}\left\{\left(\frac{\partial \bar{Q}}{\partial R}\right)R'^2 + \left(\frac{\partial \bar{H}}{\partial C}\right)C'^2 + HC''\right\}, \tag{6}$$

$\left(\frac{\partial \bar{Q}}{\partial R}\right)$ and $\left(\frac{\partial \bar{H}}{\partial C}\right)$ being the slope of the income-compensated housing demand and labor supply curves. If these curves were not perfectly inelastic, and if commuting costs increased at a non-increasing rate $\left(C'' \leq 0\right)$, the right-hand-side of (6) would be positive, meaning that housing prices must decrease at a numerically decreasing rate $\left(R'' > 0\right)$ for the equilibrium not be unstable.

The equilibrium conditions do not point out the equilibrium location, for any location can be of equilibrium provided that $R\left(D\right)$ satisfies (4)-(5). Location, indeed, is determined by such factors as income or preferences. Yet, as the price of housing and of leisure vary across the city, the worker's demand for L may be influenced, in turn, by his location. By totally differentiating the first-order conditions for the short-run optimum, and assuming the worker is initially in disequilibrium (i.e. $-QR' \neq HC'$), we obtain

$$\frac{\partial L}{\partial D} = -C's_{LL} + R's_{QL} - \left(QR' + HC'\right)\mu_L\,. \tag{7}$$

The term $-C's_{LL}$, where s_{LL} is the income-compensated own-price derivative of L, is positive. It represents the demand for L increasing with D as the price of L varies inversely with D. The varying price of housing can also influence the demand for L when s_{QL}, the income-compensated cross-price derivative of L with respect to the price of Q, is different from zero. Finally, the term $-\left(QR'+HC'\right)\mu_L$, where μ_L is the income derivative of L, is the result of the increased full income which results from the move towards the equilibrium distance when (5) holds. Although the sign of (7) is ambiguous, if L were a normal good and Q and L were net complements, expression (7) would be positive: a worker residing farther would want to spend more years in retirement. As years retired tend to be bunched (see, e.g., Lazear 1986 for reasons why this occurs), this behavior would induce a negative retirement age gradient in the urban space.

In absence of moving costs, a retired worker would move to the city boundary ($\bar{D}$) to take advantage of cheaper housing services at ages where he cannot be penalized by greater commuting costs.[4] Hence, for given earnings, the prices of L and Q that shape the demand for L would depend on $\bar{D}$, i.e. they would be constant everywhere in the city, and the two driving forces of the retirement age gradient would no longer apply. On the other hand, establishing a retirement age gradient does not require the assumption of centralized employment. Consider for example White's (1988) decentralized employment model, where a historic CBD co-exists with firms located in all directions around it, and workers settle farther from the CBD (but along the same ray) that decentralized firms. Distance to the workplace is $D-V$, where V denotes job location measured in miles from the CBD. A relationship between L and workplace distance arises out of White model's equilibrium, although its interpretation is complicated by the different relocation possibilities underlying a given change in $D-V$. Interestingly, when decentralized firms reap all of their workers'

[4] I thank an anonymous referee for raising this point.

commuting savings by paying lower wages, White (1988) model's relationship between L and workplace distance collapses to that of the centralized employment case.

3. DATA AND ECONOMETRIC FORMULATION

In order to determine empirically how responsive retirement ages are to workplace distance, consider the following retirement demand function,[5]

$$H = \mathbf{z}'\boldsymbol{\delta} + \alpha\left(D\right) + u \,, \qquad (8)$$

where H is retirement age; $\mathbf{z}$, a vector of personal characteristics that might simultaneously affect residential and retirement decisions; $\boldsymbol{\delta}$, a vector of unknown parameters; $\alpha\left(D\right)$, a scalar function of distance; and u, an unobserved variable assumed mean-independent of $\mathbf{z}$ and D. The function $\alpha\left(D\right)$ represents the relationship of interest.

Data to estimate (8) are from the Panel Study of Income Dynamics (PSID), a longitudinal survey that has interviewed a representative sample of the U.S. population mainly on a yearly basis.[6] The PSID is particularly unique as a rich source of information on labor force and commute characteristics. In nearly all the interviews conducted between 1969 and 1986, information on distance to the workplace ("About how many miles is it to where you work?") and travel to work time ("About how much time does it take you to get to work each day, door to door?") was collected from persons who stated they were working. Furthermore, the PSID labor force data make it possible to construct an empirical counterpart

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[5] I employed the P_E test suggested by MacKinnon et al. (1983) to choose between (8) and a model having the logarithm of H as regressand. For each specification presented in Table 3, neither model was rejected against the other at standard significance levels, although the non-rejection of (8) was established with much more margin.

[6] The PSID is primarily sponsored by the National Science Foundation, the National Institute of Aging, and the National Institute of Child Health and Human Development.

to our notion of retirement as a voluntary transition from work to complete labor force withdrawal.

As the availability of work travel information at the individual level depends upon the answer to "Are you working now, unemployed, retired, or what?", I will use this question to define retirement empirically: A person declaring to be "Working now"/"Only temporarily laid off" followed by "Retired" in two consecutive interviews ($t-1$ and t) will be considered as retiring at t. The employment status question has been asked of male heads of household since the inception of the PSID, but it was not until the 1976 interview that the answers "Retired" and "Permanently disabled" were codified separately. As this study is aimed at analyzing voluntary retirement decisions, I have limited the analysis to retirements occurring between (and including) 1976 and 1987. The 1987 limit is due to this wave being the nearest to the last interview asking about commuting. If an individual reports more than one transition into retirement during this period, only the latest will be included in the sample, for this is the closest transition to complete labor force withdrawal to be observed.

Retirement age is generated from the reported month and year of birth and the date of the t interview. It is measured in whole years. Commute characteristics are taken from the $t-1$ interview. Distance is measured in whole miles, which certainly introduces some gaps with the true measurement, whereas time is measured in minutes but takes on values in the set of rational numbers.[7] The extent of measurement error in commuting variables due to

[7] The PSID makes available the annualized hours spent traveling to and from work. I calculated the time of a one-way commute dividing annual commuting hours by the product of weeks worked per year and commuting trips per week. Trips per week were assigned on the basis of hours worked per week according to this rule: if hours were less than 5, 2 trips were assumed; if hours were between 6 and 10, 4 trips were assumed; if hours were between 11 and 20, 6 trips were assumed; if hours were between 21 and 35, 8 trips were assumed; if

respondents' misreports does not seem large: signal to total variance ratios for miles are in the neighborhood of 0.88, while for minutes they are around 0.83.[8] As workers located at the same physical distance to the workplace may bear different commuting costs—due for example to different levels of congestion, I shall measure D both in miles and minutes.

The theory in Section 2 did not indicate the shape of $\alpha(D)$. Graphical inspection of the (unconditional) relationship between H and D did not suggest a possible functional form either. This paper employs an information criterion to select a model for $\alpha(D)$ out of ten linear-in-parameters model types: polynomials of degree one, two, three, and four; piecewise constant functions with one, two, and three breakpoints; and piecewise linear functions with one, two, and three breakpoints. I have limited the choice set to models linear in parameters as the sample contains individuals living in different cities, and retirement age gradients might differ across cities. However, we can hope to identify an average retirement

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hours were between 36 and 47, 10 trips were assumed; for 48 or more hours, 12 trips were assumed. Commuting time was not asked in 1982, and is only available for those who worked during the previous calendar year.

[8] To calculate these figures, I followed the procedure in Bound et al. (2001, p. 3729). I constructed a sample of pairs of consecutive interviews ($t-1$ and t) pertaining to heads reporting the same residence, job, and commuting mode. Commute responses at $t-1$ and t were considered error-ridden indicators of the same true workplace separation: $D_{t-1} = D + m_{t-1}$ and $D_t = D + m_t$, where m_{t-1} and m_t are measurement errors assumed uncorrelated with D and with each other. Since the covariance between D_{t-1} and D_t equals the variance of D, the signal to total variance ratio for D_{t-1} (respectively, D_t) can be estimated as the sample covariance between D_{t-1} and D_t divided by the sample variance of D_{t-1} (D_t).

age gradient provided city-specific parameters enter $\alpha(D)$ linearly, and these parameters are independent and identically distributed across cities with distribution that does not depend on D and $\mathbf{z}$ (Cameron and Trivedi, 2005, p. 94). The limitation to models linear in parameters is less restrictive than it may initially appear, for polynomials and piecewise constant/linear functions are popular tools for identifying unspecified nonlinearity in regression functions.

Most of the postulated model types for $\alpha(D)$ nest more than one model. For example, the degree two polynomial,

$$\alpha(D) = \alpha_1 D + \alpha_2 D^2, \tag{9}$$

nests a one-parameter model (if α_1 is put to zero) and a model with two parameters. Models nested within each piecewise constant/linear type are not distinguished by the number of parameters, but by the position of the breakpoints over the range of D. For example, a piecewise linear function with one breakpoint,

$$\alpha(D) = \alpha_1 (d_0 D + d_1 s) + \alpha_2 d_1 (D - s), \tag{10}$$

where $d_0 = 1$ if $D < s$, $d_1 = 1$ if $D \geq s$, and s denotes a known breakpoint, has two parameters, but its fit to data will crucially depend on where the breakpoint is placed. I estimate all possible models within each model type. In the case of piecewise constant/linear functions, breakpoints are placed successively at every decile of the sampling distribution of D. To select among models, I use Schwarz's (1978) Bayesian Information Criterion,

$$\mathrm{BIC} = \ln \mathrm{SSR} + \frac{\mathrm{K} \ln \mathrm{N}}{\mathrm{N}}, \tag{11}$$

which is preferred to other popular criteria when some modeling alternatives are nested (Nishii, 1988). In expression (11), SSR denotes a model's sum of squared residuals, K is the total number of parameters in (8), and N represents the sample size. The model with the lowest BIC is favored. Granger et al. (1995) provide further details about, as well as a critical assessment of, model selection procedures based upon information criteria.

The specification of $\mathbf{z}$ is kept the same throughout the selection process. Besides an intercept, included in $\mathbf{z}$ are: labor income (in thousands), asset income categories (corresponding to quartiles in the asset income distribution),[9] the number of weeks of work missed through own illnesses (divided by 4.3), educational category, whether the individual is black, whether the individual is currently married with spouse present or living with a partner, whether the wife works in the market, whether the family unit comprises three or more members, whether the individual has a white collar occupation, and a linear cohort trend. All these characteristics are taken from the $t-1$ interview. The cohort time trend, calculated as the individual's year of birth minus the minimum birth year in the sample plus one, simultaneously controls for cohort and year effects, for the interview year can be inferred from information on age and year of birth. Thus, the cohort trend will not only absorb any secular trend toward earlier retirement and increased home-workplace separation present in the data, but also the influence on the timing of retirement of institutional features (such as Social Security parameters and mandatory retirement legislation) prevailing in a given year.

While some worker's characteristics can be considered as predetermined to the timing of retirement, the income and health declared at the $t-1$ interview are age-dependent, and therefore endogenous in a regression for retirement age. To correct for this, I firstly regressed the income and health variables on a complete set of age dummies using Ordinary Least Squares (OLS), obtained the residuals (i.e. income and health net of age effects), and replaced

[9] I here follow Burtless (1986), who found a non-linear relationship between wealth holdings and retirement age. Wealth holdings are proxied with asset income because information on wealth was not collected by the PSID until 1984.

the original income and health variables in $\mathbf{z}$ with their corresponding residuals. The inclusion of these generated residuals in $\mathbf{z}$ does not invalidate the standard errors of OLS estimates of (8) (Pagan, 1984).

In order to obtain a group of individuals for whom the theoretical model is likely to be relevant, I discarded farmers, individuals residing in census sampling units with no cities over 50,000, tenants (as these have lower moving costs than owners), and owners who, at t, declared that had moved since the last interview. I also discarded individuals retiring before age 51 or after age 75, reporting miles or minutes above the 95th percentile of the corresponding sampling distribution, reporting 0 miles but positive minutes or vice versa, or presenting missing or inconsistent data in some other personal characteristic. These restrictions left a sample of 201 men, of which 186 provide valid information on distance and 177 on time. Sample statistics are presented in Table 1. The average retirement age is 63.3 years, whereas the average one-way commute is 8.1 miles long and takes 22 minutes. About 7% of the sample works from home, a figure that, as defined in Christensen (1988, p. 2), includes individuals working exclusively in the home and workers whose commute varies widely (e.g., sales representatives). The average commuting speed of those who commute is 22 mph (90% commute by car). Figure 2 plots the distribution of retirement ages. The pronounced peaks at ages 62 and 65 are well-known facts of retirement in the U.S. The peak at age 62 (the age of early eligibility for Social Security benefits and the median normal retirement age for pensions) has been related to the tax and accrual aspects of pensions and to liquidity constraints (Stock and Wise, 1990; Rust and Phelan, 1997; French, 2005), whereas the high job exit rate at age 65 is viewed as an artifact of incomplete health insurance markets and the actuarial unfairness of Social Security (Burtless, 1986; Madrian et al., 1994; Rust and Phelan, 1997).

Before proceeding with the empirical results, an issue requires some discussion. In practice, not all sample commutes will end up in the CBD (Gobillon et al., 2007, found that the proportion of jobs located in U.S. central cities in 1980 was 57%). To assess the consequences that this may have on the estimation, consider for example White's (1988) decentralized employment model, where workers in-commute, one-way commuting distance is $D-V$, and the price of housing depends only on D. Since miles/minutes to the workplace measure $D-V$, they properly represent commuting costs (and thus the price of L for given yearly earnings). On the other hand, miles/minutes to the workplace measure D with error, and thus misrepresent the price of housing implicit in a residential location. In White's model, the error of measurement, $-V$, may influence D, suggesting that the lack of information on residential location may bias the parameter estimates. Since V is chosen by the worker and could be therefore correlated with $\mathbf{z}$, the direction of any bias is difficult to predict. To assess the robustness of the results to the lack of accurate information on residential location, I will proxy the price of housing with a price-per-room variable constructed as reported house value divided by the number of rooms (Kohlhase, 1986).

4. EMPIRICAL RESULTS

Table 2 presents the BIC values achieved by the ten model types for $\alpha(D)$ listed in the preceding section. For model types nesting more than one model, the BIC value corresponds to the best fitting alternative, which is indicated in the table either by the variables included in the polynomial function or by the position of the breakpoints over the range of D. Whether D is measured in miles or minutes, the piecewise constant function with one breakpoint placed at the first decile of D is clearly the best data fitting option. This result implies that the urban retirement age gradient is best modeled as a level shift. In the case of miles, where the breakpoint is at one mile and the only data value below one is zero miles, the level shift is between home-based workers and commuters. With D measured in minutes, the breakpoint is at 5 minutes, the level shift thus being between workers whose commute takes less than 5 minutes (including zero) and workers having 5-minutes or longer commutes.

The magnitude of the shift is shown in Table 3. This table presents OLS estimates of equation (8) with $\alpha(D)$ specified as an indicator variable: In column (1), it equals one if miles are positive and equals zero otherwise, whereas in column (2) it equals one if the worker's journey to work takes five or more minutes and equals zero otherwise. Heteroskedasticity robust standard errors are shown in parentheses, and probability values in brackets. The estimated coefficient on $\alpha(D)$ presented in column (1), -2.6, suggests that, on average, commuters retire 2.6 years earlier than comparable home-based workers. A similar pattern emerges from the estimated coefficient on $\alpha(D)$ presented in column (2), -1.6: on average, workers whose commute takes 5 minutes or longer retire 1.6 years earlier than comparable workers with shorter commutes. (Placing the break at 0 minutes, the estimated coefficient is -1.8.) Although both estimates are not very precise, they are nonetheless statistically different from zero at .05 level.

The other estimated parameters presented in Table 3 seem reasonable. More weeks of work missed through own illnesses reduce the age of retirement. Increases in labor income or in education are positively associated with retirement age. Individuals with a college degree, for example, retire some 2.5 years later on average than individuals not having a high-school diploma. The coefficient associated to the linear cohort trend is negative and very precisely measured. Given the way in which we have defined cohorts, with bigger values corresponding to younger cohorts, it implies that, *ceteris paribus*, each cohort reduces its average retirement age by 0.6 years solely as a consequence of the cohort and time effects prevailing in the study period.

These results appear to be robust to a variety of alternative specifications. There was no change in the best data fitting model and little effect on parameter estimates when the 24 individuals reporting more than one transition into retirement were removed from the sample (the estimated coefficient on $\alpha(D)$ became -2.45, *SE*=0.78, for miles, and -1.56, *SE*=0.60, for minutes). I have also experimented by including price per room (in real terms) in $\mathbf{z}$. In this alternative specification, the effect on retirement age exerted by the price of housing is disentangled from that exerted by the price of leisure, which is now captured by $\alpha(D)$. The estimated coefficient associated to price per room was in no case statistically different from zero, and no change was observed in the best data fitting model and in the point estimates on $\alpha(D)$. Due to concerns regarding the endogeneity of hours worked (implicitly included in labor income), the model was also re-estimated using the hourly wage (calculated as labor income divided by the product of annual weeks worked times average hours per week) as the empirical counterpart to W. The main findings were preserved, although there was some evidence that the coefficient on the hourly wage could be attenuated.

The commuting time and distance of the self-employed is considerably less than those of the employees (van Ommeren and van der Straaten, 2008). Hence, if the self-employed differed from employees in unobserved characteristics relevant to the age of retirement (e.g., pension status), D would be endogenous in (8). On the other hand, the self-employed are much more likely to continue working than comparable employees (Fuchs, 1982), which questions the inclusion of a self-employed indicator in $\mathbf{z}$. To deal with this issue, I have re-estimated model (8) including a self-employed indicator in $\mathbf{z}$, but excluding from the sample individuals reporting more than one transition into retirement. The exogeneity of the self-employed indicator seems more tenuous for these workers, as they retire much later and are much more likely to be self-employed than individuals reporting only one transition (these differences are statistically significant at .01). Again, there was no change in the best data fitting model and little effect on parameter estimates: -2.25, *SE*=0.92, for miles, and -1.80, *SE*=0.81, for minutes. The estimated coefficient on the self-employed indicator ranged from -

0.45 to 0.35. However, if individuals reporting more than one transition were included, the estimated effect of distance would be lower (-1.94, *SE*=0.89, for miles, and -0.89, *SE*=1.01, for minutes),whereas the effect of being self-employed would be around 1.

One might also wonder whether the set of models postulated for $\alpha(D)$ is sufficiently wide to contain a good approximation to the truth. Besides, variables such as labor income, the number of weeks of work missed through own illnesses, and the cohort trend were simply entered linearly in $\mathbf{z}$. To detect possible functional form misspecification in the regression for H, I applied Ramsey's (1969) regression specification error test (RESET). RESET adds powers of the fitted values to equation (8) and then tests their joint statistical significance. Under the assumption of no functional form misspecification, the powers of the fitted values will be jointly insignificant. I have added $\hat{H}^2$, $\hat{H}^3$, and $\hat{H}^4$, and tested their joint significance with a heteroskedasticity-robust Wald statistic, which is asymptotically distributed as χ^2 with three degrees of freedom. Each column of Table 3 presents the *p*-value for this test. In both cases, the claim of no functional form misspecification is well within confidence bounds.

The preceding results suggest that the retirement age gradient presents a discontinuous drop at a distance of some five minutes to the place of work, remaining flat thereafter. In this and the following paragraphs, we subject this finding to further scrutiny. *A priori*, the flat portion of the gradient could be indicative of years retired and housing services being net substitutes: As workers move farther, the increased demand for leisure caused by the own-price effect would be offset by a cross-price effect stemming from reductions in the price of housing. However, the insignificance of the price per room variable in the retirement age regression suggests that the age of retirement is unaffected by residential location for distances greater than five minutes.

The drop at five minutes could be the result of walking being substituted by car driving starting from that distance: Workers' marginal costs of commuting would suddenly rise as a consequence of the fixed costs implicit in owning a car (technological depreciation, insurance, etc.), thus lowering the price of leisure and the retirement age. Yet, the proportion of drivers among those whose commute takes less than 5 minutes is quite high (63 percent). More importantly, the separation at 5 minutes appears to be an artifact of model selection, where breakpoints were placed at deciles of D only. Separating individuals reporting no travel to work time from those reporting positive commuting time produces a model whose fit to data (BIC=7.640) is practically as good as the model with breakpoint at 5 minutes. These results, coupled with that obtained when D is measured in miles, strongly suggest that the retirement age drop is associated to the fact of non working from home.

Another consideration is the reason behind the retirement age gap between commuters and home-based workers. Male home-based workers have been found to differ in characteristics such as education, race, marital status, or unearned income from commuters (Kraut and Grambsch, 1987), but these factors have been essentially taken into account in the regression analysis. Since, for given earnings, yearly commuting costs lower commuters' opportunity price of a year retired, the different price of leisure could be playing some role in generating the gap. To test this hypothesis, one could compare the retirement age of home-based workers with that of workers reporting walking as their means of commuting. The assumption behind this comparison, that the latter group bears low commuting costs and thus has a similar price of leisure than the former, does not seem too unreasonable. In our sample, for instance, the average one-way journey to work time of those walking is 5 minutes, whereas in Smith's (1991) modal choice study the psychological cost of walking was perceived as slight by those walking to get to work. To perform the test, I regressed H on $\mathbf{z}$ and a series of indicator variables reflecting the means of transport chosen (excluding the category of those walking). The estimated coefficient on the home-based worker indicator, 2.15, suggests that home-based workers retire much later than comparable individuals who walk to get to work, questioning the role that leisure prices play in generating the retirement age gap between commuters and home-based workers. Yet, the associated standard error is large, 1.81, and the estimate is not statistically significant.

5. CONCLUSION

This paper has shown that the urban distribution of workers with high moving costs can generate differences in their age of retirement. These differences stem from the urban variation in the prices of leisure and of housing, which are the result, in turn, of the varying cost of transport. Thus, for example, if years retired are a normal good and housing and years retired are net complements, workers living farther from the place of work will retire earlier. This theoretical prediction opens the possibility that the increased home-workplace separation observed among U.S. older workers in the 1970s and 1980s were partly responsible for the contemporaneous reduction in older men’s labor force participation. In addition, the existence of an empirical relationship between residential location and retirement age would open a new avenue for intervention in the labor force decisions of older workers.

We have also explored empirically the relationship between residential location and retirement age. An information criterion has been used to choose the best data fitting model for that relationship out of a relatively wide pool of popular linear-in-parameters model types. We have found evidence of a significant retirement age gap between commuters and home-based workers in a group of urban male owners interviewed by the Panel Study of Income Dynamics: On average, a commuter retires some 2 years earlier than a comparable home-based worker. However, the evidence provides no support for the role of leisure or housing prices in generating this retirement age gap.

The validity of these empirical results rests of course on a number of assumptions. First, it has been assumed that the efficiency of the urban transportation system does not influence workers' residential locations. Yet, there may be cases in which the modes of transport exert an influence of their own that alters the structure of a community (Stucker, 1975). Second, distance to the workplace has been used to approximate distance to the CBD, what has likely introduced measurement error. Although a price-per-room variable has been used to assess the sensitivity of our results to the lack of accurate location data, price per room could be also measured with error, as it was imputed from reported house value (Bound et al., 2001, p. 3780). Third, the empirical specification has missed key pieces of information on pension wealth and medical care, which, as suggested in Muller (1988), could be important for explaining home-based workers' higher retirement age.

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Table 1. Descriptive statistics: 1975-1987 Panel Study of Income Dynamics.

Variable	Mean	Std dev	Minimum	Maximum
Year of retirement	1982	3.5	1976	1987
Age at retirement	63.3	5.0	51	74
Miles to workplace (one-way)	8.1	6.6	0	32
Minutes to workplace (one-way)	21.8	15.2	0	60.5
Labor income ($1,000)	21.7	14.7	0	79.3
Asset income ($1,000)	2.8	5.1	0	26.8
Weeks ill	2.1	5.6	0	52

Variable (%)	Mean	*Variable (%)*	Mean
Less than a high school diploma	40.8	Married	90.5
High school graduate	15.9	Wife works	42.8
Some college	29.4	Family unit > 2 persons	35.8
College graduate	13.9	White collar worker	42.3
Black	26.4		

Notes: Data are of 201 male heads of household reporting one transition from employment into retirement in two consecutive interviews ($t-1$ and t) of the period 1975-1987. Labor and asset income are yearly rates expressed in 1982-1984 dollars. Labor income comprises wages and salaries, bonuses, overtime, commissions, and the labor part of business income. Asset income is made up of rents, interest payments, dividends, trust funds, royalties, the asset part of business income, and, if an individual is married or cohabiting, his wife's income from assets. The wife works in the market when she reports positive labor income during the calendar year of the $t-1$ interview. White collar workers are professional and technical workers, managers, and clerical and sales workers.

Table 2. BIC values.

Model type for $\alpha(D)$	Miles	Model	Minutes	Model
Polynomial of degree 1	7.481		7.661	
Polynomial of degree 2	7.486	D^2	7.663	D^2
Polynomial of degree 3	7.486	D^3	7.663	D^3
Polynomial of degree 4	7.485	D^4	7.663	D^4
Piecewise constant, 1 breakpoint	7.422	$s=1$	7.638	$s=5$
Piecewise constant, 2 breakpoints	7.442	$s=1,7$	7.660	$s=5,30$
Piecewise constant, 3 breakpoints	7.449	$s=1,5,7$	7.676	$s=5,20,30$
Piecewise linear, 1 breakpoint	7.438	$s=3$	7.657	$s=10$
Piecewise linear, 2 breakpoints	7.442	$s=5,7$	7.682	$s=14,15$
Piecewise linear, 3 breakpoints	7.461	$s=1,5,7$	7.696	$s=10,20,25$

Notes: The specification of $\mathbf{z}$ is the same in all models. Values of s are breakpoints.

Table 3. The equilibrium effect of workplace distance on retirement age.

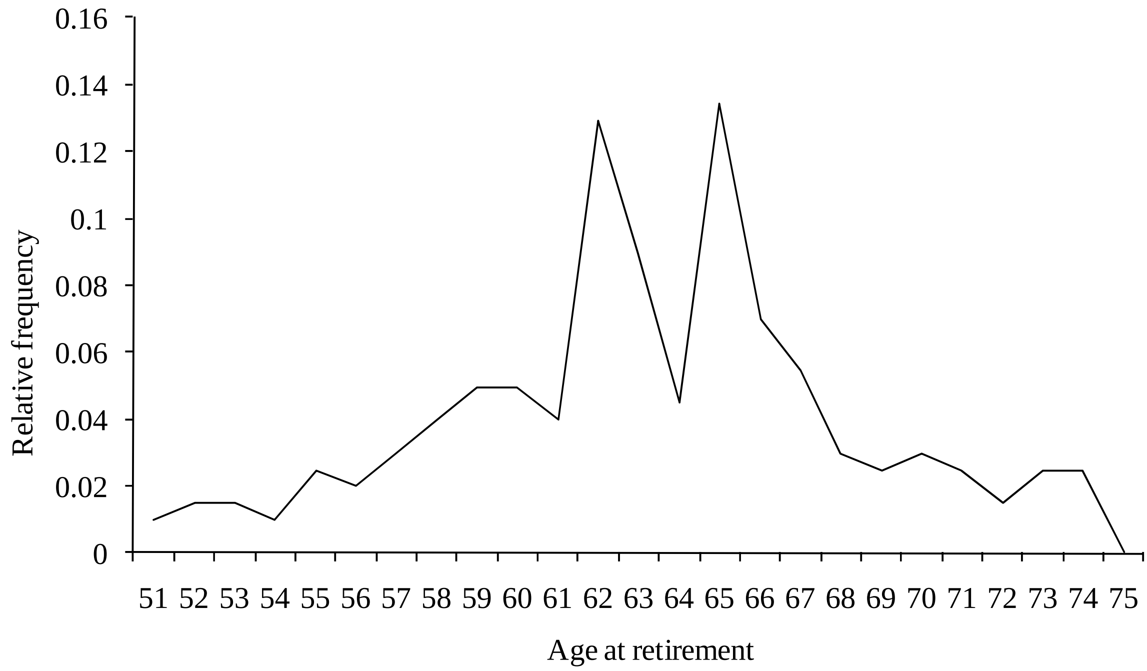
Independent variables	Dependent variable: age at retirement (1)	(2)
$1[\text{miles} \geq 1]$	-2.644 (0.698)***	
$1[\text{minutes} \geq 5]$		-1.600 (0.766)**
Labor income, $1,000	0.042 (0.014)***	0.019 (0.023)
Asset income in 2nd quartile	-1.314 (0.486)***	-1.234 (0.613)**
Asset income in 3rd quartile	0.042 (0.564)	-0.280 (0.635)
Asset income in 4th quartile	-0.926 (0.612)	-0.585 (0.764)
Weeks ill ($\div 4.3$)	-0.354 (0.162)**	-0.585 (0.247)**
High school graduate	1.322 (0.611)**	1.639 (0.694)**
Some college	1.464 (0.492)***	1.212 (0.569)**
College graduate	2.562 (0.696)***	2.229 (0.843)***
Black	1.277 (0.489)**	0.781 (0.554)
Married	0.251 (0.615)	-0.061 (0.633)
Wife works	-0.782 (0.399)*	-0.611 (0.455)
Family unit > 2 persons	-0.989 (0.440)**	-0.954 (0.508)*
White collar	-0.058 (0.466)	-0.080 (0.518)
Cohort time trend	-0.609 (0.031)***	-0.622 (0.035)***
Intercept	75.21 (0.87)***	75.36 (1.00)***
R^2	0.75	0.71
Ramsey's (1969) RESET	[0.69]	[0.69]
Observations	186	177

Notes: The estimation method is OLS in all columns. Heteroskedasticity robust standard errors appear in parentheses, and probability values in brackets. The function $1[\cdot]$ equals one if its argument is true and zero otherwise. Unreported categories: asset income in 1st quartile; less than a high school diploma. * Significant at 10%. ** Significant at 5%. *** Significant at 1%.

Figure 1. Labor force participation rate, U.S. males 55-64 (solid), and average home-workplace distance, U.S. male workers 55-64 (dotted).

Notes: Author’s calculations. Data on participation are from the Bureau of Labor Statistics. Data on distance for 1971-1986 are from the Panel Study of Income Dynamics, whereas for 1995 are from the Nationwide Personal Transportation Survey.

Figure 2. Retirement age distribution



Notes: Author’s calculation with data from the PSID, 1976-1987.