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Yingying Dong
University of California, Irvine

Dennis Yang
University of Virginia

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Yingying Dong
University of California, Irvine
e-mail: yyd@uci.edu

Dennis Yang
University of Virginia

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ABSTRACT

This paper investigates household consumption changes at retirement by utilizing a comprehensive, diary-based household survey from China. The survey contains both consumption quantity and price information, which permits separating quantity changes from price changes. The mandatory retirement policy in China provides a quasi-experimental setting for identification of the true causal effects of fully anticipated retirement. Using regression discontinuity models, we show that food expenditure declines at retirement, particularly among the low-education group, and that the decline is driven by price declines instead of quantity declines. Shopping time for food increases at retirement, consistent with the price and quantity changes.

JEL Classification Codes: J26, C21

Key Words: Retirement-consumption puzzle, Mandatory retirement, Regression discontinuity, Consumption vs. expenditure, Time use, Home production.

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Mandatory Retirement and the Consumption Puzzle: Prices Decline or Quantities Decline?

Yingying Dong, Dennis Yang[‡]
University of California, Irvine, and University of Virginia

February 2016

Abstract

This paper investigates household consumption changes at retirement by utilizing a comprehensive, diary-based household survey from China. The survey contains both consumption quantity and price information, which permits separating quantity changes from price changes. The mandatory retirement policy in China provides a quasi-experimental setting for identification of the true causal effects of fully anticipated retirement. Using regression discontinuity models, we show that food expenditure declines at retirement, particularly among the low-education group, and that the decline is driven by price declines instead of quantity declines. Shopping time for food increases at retirement, consistent with the price and quantity changes.

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1 Introduction

Many empirical studies show that consumption (typically food consumption) drops significantly at retirement. This finding is referred to as the "retirement-consumption puzzle," because a systematic fall in consumption is inconsistent with consumption

[*]Yingying Dong, Department of Economics, University of California, Irvine, yyd@uci.edu; Dennis Yang, Darden School of Business, University of Virginia, YangD@darden.virginia.edu.

[†]The authors would like to thank James Heckman, Michael Hurd, Robert Pollak, Josh Angrist, Daron Acemoglu, Amy Finkelstein, Michael Greenstone, David Autor, Esther Duflo, Jim Poterba, Kerwin Charles, seminar participants at MIT, and participants at the CEA, WEAI, CES annual conference for helpful comments. All errors are our own. This research is partly supported by an Early Career Research Grant from the W. E. Upjohn Institute for Employment Research.

smoothing suggested by the life cycle/permanent income hypothesis (PIH, Friedman [1956], Hall [1978]). According to the PIH, the marginal utility of consumption should stay constant over the life cycle, so rational people are expected to smooth consumption through borrowing and (dis)saving, and consumption should not fall when retirement is anticipated.

This "puzzle" has been documented using various data sets from many different countries, including the United Kindom, the United States, Italy, Canada, and Germany.[1] These studies all use expenditure data. Whether this finding can be interpreted as showing that households are ill-prepared for retirement depends upon how expenditures map into consumption as well as the ability of researchers to delineate between expected and unexpected retirement.

In an influential paper, Aguiar and Hurst (2005) suggest that retired households have considerably more leisure time, so they can shop for bargains and thereby pay lower prices for the same quantity and quality of goods. In this case, expenditure declines at retirement, but actual consumption does not. They also argue that retired households can engage in more home production, which enables them to substitute home meals for restaurant meals. In another paper, Aguiar and Hurst (2007) show that average prices paid for goods decrease while shopping time increases over the life cycle, and that purchase prices decline significantly around retirement age.

Following the insights of these two influential papers, we provide the first direct causal evidence on actual consumption changes at retirement. In particular, we separate quantity changes from price changes at retirement, utilizing the mandatory retirement policy in China for identification and a large confidential dairy-based consumption data set for estimation. We also document changes in household shopping time for food around retirement, which are consistent with the documented price and quantity changes.

[1]See for example, Banks, Blundell, and Tanner (1998), and Smith (2006) for evidence in the United Kingdom; Bernheim, Skinner, and Weinberg (2001), Aguila, Attanasio, and Meghir (2011), Ameriks, Caplin, and Leahy (2007), Haider and Stephens (2007), and Hurd and Rohwedder (2008) for evidence in the United States; Battistin et al. (2009), Miniaci, Monfardini, and Weber (2009), and Borella, Moscarola, and Rossi (2011) for evidence in Italy; Robb and Burbridge (1989) for evidence in Canada; and Schwerdt (2005) for evidence in Germany.

The mandatory retirement policy provides a unique quasi-experimental setting for nonparametric identification of the true causal effects of fully anticipated retirement. This policy requires workers to retire at a certain age. Intuitively, workers who have just reached the retirement age and those who are just under are comparable except for their retirement status, so their mean outcome difference should be induced by retirement at that age. The identified effects are for those compliers whose retirement is precipitated by their age exceeding a predetermined threshold rather than by endogenous factors, such as health shocks or being laid off. Furthermore, retirement at the required age is fully anticipated by Chinese workers because the policy has been in effect since the 1950s. This helps clarify consumption changes when retirement is foreseen versus when it is unforeseen.

Compared with most of the pension eligibility rules in the West, the retirement mandate in China induces a larger change in the retirement probability, implying a more pronounced first stage. The policy therefore can help more precisely identify the impacts of retirement and provide implications for a broader group of compliers. In addition, the mandatory retirement polcy in China is entirely based on age. Age is well defined and is not easily susceptible to individuals' manipulation, as discussed later, which makes our regression discontinuity design highly credible.

Existing studies are almost exclusively based on developed Western countries. The case of China is interesting because of its unique social, cultural, and economic environment, which differs in many ways from developed Western countries. For example, Hurd and Rohwedder (2013) show that in the United States the consumption decline at retirement is mainly induced by work-related expenditures, such as clothing and transportation, and that once this is taken into account, there is no significant decline in expenditures among those who retire voluntarily. In contrast, in China workplaces typically do not impose a dress code, and most Chinese rely on cheap public transportation or bicycle to commute to work, so cessation of work may not induce significant declines in these work-related expenditures. In addition, Chinese people have the tradition of saving for old age, which is reflected in their very high saving rates. Because of these

dramatically different social norms and deeply rooted traditions, it is unclear whether existing conclusions would apply to China.

Using data from the Urban Household Survey (UHS) in China, we document that the mandatory retirement policy in China induces a significant sharp increase in the retirement rate. We futher show that food expenditures decline significantly at retirement, particularly among household heads without a college education (roughly three-fourths of our sample). Existing studies on expenditure also show that expenditure declines concentrate in the disadvantaged group, such as the low pre-retirement wealth group, blue-collar workers, workers with low education or low retirement savings (see, e.g., Bernheim, Skinner and Weinberg [2001], Borella, Moscarola, and Rossi [2011], Robb and Burbridge [1989], and Schwerdt [2005]).

More importantly, we show that the observed food expenditure decline is driven by reduced spending on eating out and by price declines instead of quantity declines. In addition, household shopping time for food increases significantly at retirement among the group that reduces its spending on food. We interpret these results as being consistent with a life-cycle model augmented with home production.

The rest of the paper proceeds as follows. Section 2 relates this study to existing studies; Section 3 describes the institutional background and the data; Section 4 discusses identification and the empirical specification; Section 5 investigates expenditure changes; Section 6 investigates price and quantity changes; Section 7 discusses the food consumption index and predicted food expenditure, holding prices fixed; Section 8 investigates food shopping time; Section 9 discusses the time-varying retirement effects. Section 10 examines the validity of our RD design. Brief concluding remarks are provided in Section 11.

2 Relationship to Existing Studies

There is a large literature on the retirement-consumption puzzle.[2] Existing studies largely rely on structural models for identification. The only studies that, similar to ours, utilize an RD design are Battistin et al. (2009), and more recently, Li, Shi, and Wu (2015). Whether the observed consumption decline is a puzzle is highly debated. Hurst (2008), after reviewing the existing studies based on developed countries' data, concludes that the large body of existing work represents "the retirement of a consumption puzzle." We contribute to the literature by distinguishing between expenditures and consumption and by providing direct causal evidence on real consumption changes at retirement.

As noted in Hurst (2008), the documented consumption decline utilizing developed countries' data is mainly a decline in nondurable consumption, particularly food and work-related consumption. The estimated sizes of the decline vary dramatically across studies. They fall anywhere from less than 10% to more than 30%. There also appears to be great heterogeneity. Consumption generally declines more among the disadvantaged group. For example, in the United States, Bernheim, Skinner, and Weinberg (2001) show that the consumption decline is negatively correlated with retirement savings and income replacement rates. Borella, Moscarola, and Rossi (2011) find that consumption drops more among the low-wealth group in Italy. Robb and Burbridge (1989) find that consumption declines significantly only among the blue-collar households in Canada. Schwerdt (2005) shows that consumption drops more than 30% for the low-income replacement group in Germany, while it increases more than 10% for high-income replacement individuals.

Various explanations have been proposed to reconcile the observed expenditure decline. Except for those related to cessation of work (work-related expenditure declines) and time use changes (such as shopping for bargains and increased home production), French (2005) and Blau (2008) emphasize that retirees enjoy increased leisure time and

[2]See, e.g., Banks, Blundell, and Tanner (1998), Bernheim, Skinner, and Weinberg (2001), Schwerdt (2005), Smith (2006), Haider and Stephens (2007), Ameriks, Caplin, and Leahy (2007), Hurd and Rohwedder (2008), Wakabayashi (2008), Miniaci, Monfardini, and Weber (2009), Battistin et al. (2009), and Aguila, Attanasio, and Meghir (2011).

so can reduce consumption without having their utility level affected, assuming nonseparability of preferences for leisure and for consumption.

An alternative explanation emphasizes that workers undersave for their retirement, either because they lack self-control (Angeletos et al. 2001) or because they fail to adequately foresee the income decline at retirement (see, e.g., Bernheim, Skinner, and Weinberg 2001). Consistent with the latter, Gustman and Steinmeier (2001) find that misinformation or lack of information about retirement benefits is the norm among U.S. workers. Lusardi (1999) finds that, ceteris paribus, households who have given little thought to retirement have far lower wealth than those who have given the subject more thought.

Battistin et al. (2009) exploit the pension eligibility rule in Italy for identification and investigate expenditure changes at retirement. Unlike mandatory retirement, pension eligibility in Italy is determined by both age and contribution years, and the resulting constructed running variable for their RD design is found to have measurement errors. They carefully derive conditions under which the measurement error only leads to a fuzzier RD design. Interestingly, Battistin et al. show that in Italy retirement induces a significant drop in the number of grown children living with their parents, causing a change in household composition and household size, which partly causes consumption to decline. We show later that household size does not have a significant change when household heads retire in China. In addition, recently Li, Shi, and Wu (2015) explore expenditure changes at retirement in China using the UHS data.[3]

Unlike the existing studies, we focus on separating price changes from quantity changes at retirement when retirement is fully anticipated. We also investigate how food-shopping time changes. Our expenditure and consumption information is from the consumption diary instead of recall questions.

[3]Their data cover a slightly different time period, 2002-2009, and different geographic areas: Beijing, Liaoning, Zhejiang, Anhui, Hubei, Guangdong, Sichuan, Shaanxi, and Gansu.

3 Institution Background and Data Description

In China, the official retirement age is 60 for male workers, 55 for white-collar female workers, and 50 for blue-collar female workers, with some exceptions applying to certain occupations and to disabled workers.[4] These mandatory retirement ages have not changed since they were established in the 1950s. Mandatory retirement is strictly enforced in the state sector, including the government organizations and state-owned enterprises (SOEs), whereas workers in the private sector have more flexibility. In our sample, the majority (78%) of workers around retirement age work for the state sector. Working for the private sector is more common among younger workers, since the private sector in China virtually did not exist until the 1980s.

Workers may retire earlier before they reach the mandatory retirement age. Retirees may also take a new job or may even be rehired by the same employer after their official retirement, so the change in the retirement rate is less than 1 at the mandatory retirement age, which entails a fuzzy RD design.

In China, the replacement rate (pension as a fraction of a worker's pre-retirement income) depends on the duration of pension program participation and on pre-retirement occupation. To be eligible for a pension, one must participate in the program for a minimum of 10 years.[5] Typically, a worker with 10 years of participation receives 60% of the pre-retirement wage, and the replacement rate goes up to 70% for workers with 15 years or more. The maximum replacement rate for civil servants is 88%, and for government institution workers it is 90%. A small number of workers, those who started working for the Communist Party before 1945, get a 100% replacement rate.

We use data from the China Urban Household Survey (UHS). The UHS is an ongoing national annual survey of urban households conducted by the China National Bureau of

[4]Those who have jobs that are risky, harmful to their health, or extremely physically demanding can retire five years before the official retirement ages—45 for blue-collar female workers and 55 for male workers. Male workers who become disabled and hence are unable to do their work can apply to retire at 50, while disabled female workers can retire at 45. Civil servants also qualify for early retirement if they have worked for 30 years and are within five years of their retirement age.

[5]For those joining the workforce after 1993, the minimum number of years of contribution required is 15, but this is irrelevant for the retiree cohorts we are looking at.

Statistics (NBS). The first wave of the UHS was conducted in 1988. The UHS surveys a large representative sample of urban households (30,000–50,000 every year for our sample period) and provides detailed information on household consumption and income as well as household members' education, employment, and demographic information. One unique feature of the UHS data is that the food consumption is collected through a consumption diary, and so it has not only expenditure information but also information on quantities purchased for detailed consumption categories. Unlike data collected through recall questions, the data are less likely to have recall errors. The UHS data have been used to compile CPI and monitor consumption changes over time. The rich information in the UHS allows us to consistently investigate consumption expenditure, quantity, and price changes for refined categories and conduct RD analyses.

Limited by data availability, we use a subset of the UHS sample, representing urban households in five provinces and one municipality.[6] Selection into the sample of urban residents at retirement is not an issue here. Because of the restricted household registration or HuKou system in China, workers rarely ever move to rural or other areas upon retirement. We choose the sample period 1997–2006, mainly because the UHS questionnaires changed a few times over the years, but the questionnaires are largely consistent for the period of 1997–2006. In addition, the pension system in China changed in 1997. Starting that year, the Chinese government adopted a system that combines individual accounts and social pooling to provide retirement funds. Before that, pensions were provided entirely by employers.[7]

We focus on male workers who are household heads for clean identification. Female workers' labor supply is more complicated, and their mandatory retirement age varies across occupations. However, this may cause another sample selection issue—i.e., the reported household head may change with the head's retirement status, which would invalidate our research design. If this is true, we would expect the probability of being

[6]The five provinces are Liaoning, Zhejiang, Guangdong, Shanxi, and Sichuan, and the one city is Beijing.

[7]Since we are looking at individuals close to retirement age, even if this change has any effects on the size of the compliers at mandatory retirement, it should not invalidate our identification. We find that including or dropping data from 1997 does not affect our results much.

a household head to change at the mandatory retirement age of 60. To make sure the sample selection does not undermine our identification, we plot the fraction of household heads at each age among the full UHS sample of males. As shown in Figure 1(a), the probability of being a household head changes smoothly with age, and there do not appear to be visible changes at the mandatory retirement age of 60.

Eligible male workers can start their retirement paperwork at the beginning of the month they turn 60. Typically the paperwork is processed within the same month, and eligible workers start to receive a pension the following month after they turn 60. Retirement status in our study is obtained from the UHS survey question on individuals' employment status. A household head is retired if his employment status is "retiree." We do not include in our sample those who are not labor force participants. Consumption outcomes we look at include expenditures, quantities, and prices. All categories of expenditures and prices are adjusted for regional specific inflation and are in 1996 constant Chinese Yuan.

4 Identification and Empirical Specification

4.1 RD Model Identification and Specification

As a quasi-experimental approach, a standard RD design identifies the effect of a binary treatment when the assignment of treatment is determined by an observed covariate, the so-called "running variable," exceeding a known threshold. RD identification associates a discrete change in the treatment probability at the threshold with a corresponding discrete change in the mean outcome. Here the treatment is whether a male household head is retired or not. The running variable is the household head's age.

For now, we consider the standard static RD model, which identifies the immediate effect of retirement on household consumption at the mandatory retirement age of 60. Time-varying effects of retirement are discussed later in Section 9.

Let Y be household consumption. Let T be the binary indicator that equals 1 if

a household head is retired and 0 otherwise. Let X be a household head's exact age relative to the mandatory retirement age of 60. Also define $D = I(X \geq 0)$, where $I(\cdot)$ is an indicator function that equals 1 if its argument is true and 0 otherwise. Therefore, $D = 1$ for household heads who are at or above the mandatory retirement age, and $D = 0$ otherwise. Assume household consumption changes smoothly with a household head's age, which may also depend on the head's retirement status. We can then write the consumption model as

$$Y = f(X) + \tau_0 T + \varepsilon, \tag{1}$$

where $f(X)$ is a low-order polynomial of X, and ε captures all other smooth factors that determine a household's consumption. Note that the standard argument applies—in Equation (1), τ_0 captures the average effect of retirement on consumption at the mandatory retirement age, even if in the true consumption model the retirement effect is heterogenous, so there are interaction terms between T and covariates (such as age X).

The retirement equation can similarly be written as

$$T = \sum_{j=0}^{J} a_j X^j + \sum_{j=0}^{J} b_j X^j D + v, \tag{2}$$

where J is the order of polynomial and v is a smooth regression error.

Plugging Equation (2) into Equation (1) yields the reduced-form consumption equation

$$Y = \sum_{j=0}^{J} c_j X^j + \sum_{j=0}^{J} d_j X^j D + u. \tag{3}$$

For simplicity, assume that it is a local polynomial regression of order J, though one could allow the order of polynomials to differ for the retirement and the consumption equations. In this case, J can be taken as the higher order of the two.

Both consumption Y and retirement T could depend on other covariates, which are suppressed for now, since a generic virtue of the RD approach is that inclusion of other covariates (assumed to be smooth) only affects efficiency but not consistency of estimated

RD treatment effects. The above equations allow the slopes and higher-order derivatives of retirement and consumption profiles to differ at either side of the age threshold.

Given smoothness of $f(X)$ and ε in Equation (1), any observed discontinuity in the mean consumption can be attributed to the change in the retirement rate. So the ratio of the mean consumption change to the retirement rate change at the mandatory retirement age identifies the average effect of retirement on consumption at that age (see, e.g., Hahn, Todd, and van der Klaauw [2001]). That is,

$$\tau_0 = \frac{b_0}{d_0}. \tag{4}$$

4.2 Issues with Using Age in Years

The UHS records age in years, similar to many surveys. Age in years can be seen as the exact age rounded down to the nearest integer, so, for example, a worker who is reported to be 60 in the survey can have a true age anywhere between 60 and 61 minus one day. However, RD model identification crucially relies on a continuous running variable. Using a rounded or discretized running variable may lead to biased estimates.

Given a discrete running variable, one does not observe data arbitrarily close to the cutoff even if one has an arbitrarily large sample. Extrapolation based on functional forms is unavoidable. Well-established nonparametric methods, such as the recent bias-corrected robust inference approach proposed in Calonico, Cattaneo, and Titiunik (2014), are not feasible. Intuitively, rounding down means that each age is not centered at the mean or midpoint of the corresponding age cell. Even if one can recenter the integer age to be the midpoint of the age cell (by adding 0.5 to each integer age), the curvature or nonlinearity of the age profiles can cause further problems. An illustration of the problem along with a description of a bias correction procedure utilizing the moments of the birth-date distribution within a year is provided in Appendix II. The bias correction procedure follows the general approach discussed in Dong (2014) but is adapted to facilitate obtaining standard errors directly.

4.3 The Retirement Rate Increase at the Mandatory Retirement Age

Figures 1(b), 1(c), and 1(d) show the age profiles of household heads' retirement rates, pensions, and wages, respectively. Dots in these figures represent mean values at each age. There is a clear jump in the retirement rate at age 60. Similarly, average pensions and wages change discontinuously at the same age. Consistency of all three figures suggests that household heads' retirement status is not systematically mismeasured.

The UHS is an annual survey. It records individuals' retirement status by the end of a survey year. In theory, a household head can retire at any time during the year, so household consumption at 60 is generally a mixture of pre- and postretirement consumption. When a household head retires at the end of the year, the consumption at 60 captures entirely pre-retirement consumption. We therefore drop observations at 60. This ensures that all observations below 60 are drawn from the premandatory retirement profile, and all observations above 60 are drawn from the postmandatory retirement profile. The difference in the two profiles evaluated at 60 yields the exogenous change induced by mandatory retirement policy. We accordingly estimate the polynomial regressions using data from ages 59 and below and ages 61 and above, and we evaluate changes at 60 by extrapolating these regression curves to the cutoff age of 60.[8]

We consider widely varying ranges of age for the retirement equation, 6, 10, and 15 years above and below the cutoff, corresponding to age ranges 54–66, 50–70, and 45–75. The sample sizes corresponding to the three windows are 12,050, 21,576, and 33,149, respectively. On average, there are more than 2,000 observations at each age. In practice, there is a tradeoff regarding what range of age around the threshold to include in the model. A wider range provides more observations, thereby adding to the

[8]We use a uniform kernel for convenience, since more complicated weighting or different kernels rarely make much difference in practice. The only difference between regressions using a uniform kernel and those using more complicated kernels is that the latter put more weight on observations closer to the cutoff. An arguably more transparent way of putting more weight on observations closer to the cutoff is simply to reestimate a model with a uniform kernel using a smaller bandwidth. If using different weights makes a difference, it likely suggests that the results are highly sensitive to the choice of bandwidth, a point made by Lee and Lemieux (2010).

precision with which the model coefficients can be estimated. However, the further away the included ages are from the threshold, the more likely the correct model specification for these distant observations will differ from the correct specification near the threshold, risking specification errors.

The estimated increases in the retirement rate at the mandatory retirement age are reported in Table 1a. For comparison purposes, we also report in Table 1b estimates using the observed age in years without any bias correction.

Overall we have 21 different specifications, depending on the order of polynomials, bandwidths, and whether to include covariates or not. The full set of covariates include year fixed effects, province fixed effects, year-province interactions, household size, household size squared, and heads' education levels in three categories: 1) college or above, 2) high school, and 3) less than high school (the default).[9] Goodness of fit measures (adjusted R^2 and AIC) suggest that the second order polynomial is preferable when using the short six-year window, while the third order polynomial fits better when using the 10- or 15-year window.

The estimates do not vary much across specifications. It is estimated that close to 30% of male workers retire at age 60. In contrast, the estimates in Table 1b without correcting for rounding bias seem to systematically underestimate the true increase in the retirement rate. These estimates are also sensitive to different specifications, ranging from 13.0% to 26.4%.

5 Expenditure Changes at Retirement

We first investigate whether expenditures drop at all at the mandatory retirement age in China. To put things in context, Figure 2(a) presents the age profile of household income. As expected, household income drops sharply at the mandatory retirement age. It is then interesting to investigate how household expenditure and consumption change

[9]For the short 6-year window, the fourth-order polynomial obviously overfits the curve, while for the 15-year window, a quadratic seems to greatly underfit the curve. We omit results from those specifications.

in response to the expected income drop at the mandatory retirement age. Following the large body of the existing literature, we divide the total expenditure into four categories: 1) food, 2) clothing, 3) transportation, and 4) the remaining expenditures.

5.1 Food, Work-related and Other Expenditure

Figures 2(b), 2(c), 2(d), and 2(e) present the age profiles of these four categories of expenditure. Food expenditures have an obvious drop at the mandatory retirement age. In striking contrast to food expenditures, clothing expenditures decrease smoothly with age. Transportation expenditures also do not appear to change discontinuously. These preliminary findings suggest that our results may differ from the typical findings from the wealthy Western countries. Quite a few existing studies based on developed Western countries find that work-related expenditures, particularly clothing and transportation expenditures, drop significantly at retirement (see, e.g., Battistin et al. [2009] for RD evidence in Italy).

Note that these figures present preliminary visual evidence of possible changes at the mandatory retirement age. However, because of the discrete nature of age in years, these visual changes may not map into the true changes. We estimate the effects of retirement using the bias-corrected regressions described in the previous section. The outcome and the retirement equations are jointly estimated using GMM to maximize efficiency. Household consumption crucially depends on household size and other covariates, so in these regressions we control for household size, size squared, head's marital status, year fixed effects, province fixed effects, as well as year-province fixed effects. Later we show that household characteristics are smooth at the mandatory retirement age, so omitting these covariates does affect consistency. However, covariates help reduce the sampling variation of the outcome variable, and hence may provide more precise estimates. To facilitate comparison, we restrict the bandwidth to be the same (age range 45–75) for the large number of outcomes we examine. For each outcome, we choose the optimal order of polynomial based on commonly used goodness of fit measures.

Table 2 reports the estimated changes in the four categories of expenditure when

household heads retire.[10] We report estimates separately for college-educated household heads and non–college educated heads. We show that these two groups respond to retirement very differently, consistent with the documented heterogeneity in the existing literature. A small (4–5%) decline is found in food expenditures, particularly among the non–college education group. In contrast, among the college education group, food expenditures are shown to have a positive yet insignificant change.

Clothing and transportation expenditures, which arguably are "work-related," do not show any significant declines in China. This finding is consistent with the social norm in China. Work places in China typically do not impose dress codes, so workers do not purchase business attire specifically for work. In addition, the majority of workers in China either bike or rely on affordable public transportation to commute to work. Household transportation expenditures are very low. In our sample, the median transportation expenditure among the pre-retirement-age working households is only 266 Chinese Yuan (less than US $50) per year. Because of these social norms, cessation of work at retirement does not induce either category of expenditure to drop much. In addition, even if retirees can search for lower prices for clothing, one may not see a discrete change in clothing expenditures, since clothing is a semidurable.

5.2 Food at Home and Away from Home

Figures 3(a), 3(b), and 3(c) further show the age profiles of expenditure for food consumed at home and food consumed at restaurants. Spending on eating out shows dramatic declines at the mandatory retirement age, in contrast to a small decline in expenditures of food consumed at home.

Note that expenditures for food at home increase steadily with age before reaching the mandatory retirement age (roughly by 14% over 15 years of age), and then decline

[10]For food, clothing and other expenditures, we use logged values as our dependent variables. About 9% of the households reported zero spending on transportation in our sample. To avoid dropping observations with zero spending or transforming them differently, we use level instead of logged value of transportation expenditures as the dependent variable. This ensures comparability of means across age points. We then convert the estimated level changes into percentage changes to facilitate interpretation.

quickly. This may reflect the hump-shaped life cycle profile of household nondurable consumption (Fernandez-Villaverde and Krueger 2007). It is also possible that the curvature reflects time trend or cohort effects, given our cross-sectional data. We explicitly take the hump-shaped age profile into account by including polynomial functions of age. We also control for regional specific year fixed effects. Since the cohort is a perfect linear function of year and age, cohort effects are in part captured in our specifications by the flexible smooth function of age and year fixed effects. It is worth emphasizing that the standard RD model still correctly estimates the local average treatment effect of retirement even when any smooth cohort or time effects are omitted. This point is further discussed in Section 9.

Table 3 reports the estimated changes in spending on food at home and on food away from home. For the non–college education group, food at home shows small but insignificant declines, while food away from home shows significant declines: the total spending on eating out declines by about 20%, and of this amount, spending at non-workplace restaurants declines by about 27%. In contrast, the college education group does not experience any significant declines in either type of food expenditure.

6 Price versus Quantity Changes at Retirement

Eating out expenditures decline dramatically upon retirement among the non–college education group. If they substitute home meals for restaurant meals, one should see an increase in home production upon retirement, which may not be reflected in the expenditure data if prices change. This section disentangles consumption quantity changes from price changes at retirement.

We first look at quantities purchased and average prices paid for each category of food for some major food categories. One advantage of this aggregation is that it takes into account the substitutability of different types of food within a category. Another advantage is that it reduces heterogeneity in consumption across households, since the specific food consumed by each household varies greatly. The retirement effects on

quantities and prices can therefore be much more precisely estimated by looking at food categories. The disadvantage of focusing on food categories is that it masks any compositional changes within a category. To overcome this problem, we next look at prices and quantities of some commonly consumed foods.

6.1 Prices and Quantities of Major Food Categories

We look at five major food categories, including staples, vegetables, oils, meat and poultry, and fruit.[11] Figures 4 and 5 show how the quantity purchased and average price paid for each category of food change with household head's age using the full sample. Figures 6 and 7 present similar figures but for the non–college educated heads only, while figures 8 and 9 are for the college educated heads only.

In Figure 4, average prices for all food categories show big declines at the mandatory retirement age. In striking contrast, quantities do not appear to decline discretely in Figure 5. There even appear to be small increases, if any, in quantities at the mandatory retirement age. Figures 6 and 7 for the low-education sample largely mimic the quantity and price patterns in Figures 4 and 5 based on the full sample; whereas Figures 8 and 9 for the college education sample show no discrete changes in either prices or quantities at the mandatory retirement age, though the data is rather noisy because of a smaller sample size.

Tables 4 reports the estimated changes in the average price for each food category at retirement. Consistent with the visual evidence, prices paid by the non–college education group are estimated to decline significantly for all food categories except for staples. In contrast, prices paid by the college education group are estimated to be mostly positive but insignificant.

The estimated price declines for the low education group also vary a lot across categories. For example, meat and oil are estimated to have small declines of 2.8% and

[11] "Staples" include rice, flour, and other grain or grain products; "meat" includes pork, beef, lamb, and other meat or meat products; "poultry" includes chicken, duck, and other poultry or poultry products. All age profiles of quantities are quantities per household member.

4.1%, respectively, while vegetable and fruit are estimated to have declines of 7.1% and 10.6%, respectively, among the non–college education group. The varying sizes of the price declines are consistent with possible price variations in China. For example, staple, oil, and meat prices are largely regulated by the government and hence show little variation. Greater search efforts therefore may not lead to greater price reductions. In contrast, vegetables and fruits can be purchased through a variety of channels, including some local farmers' markets. Purchase prices can vary a lot, depending on when and where they are purchased.

Table 5 reports on the estimated changes in the quantities. Quantities for almost all categories, except for fruit, are estimated to increase significantly among the non–college education group, but no significant changes are found among the college education group.

Overall, the results suggest that average prices, not quantities, of food for each category decline. Quantities may even increase upon retirement, which to some extent offsets the significant decline in meals out. These results are consistent with a substitution of home meals for restaurant meals upon retirement.

6.2 Prices and Quantities of Specific Food

The average price for a food category is constructed by dividing the total expenditure by the total quantity, and therefore represents a unit value (see Deaton [1988] for discussions on unit values). In practice, prices within a category vary across different types of food. A decline in a unit value can therefore be driven either by real price declines within a category or by compositional changes—i.e., retirees may substitute relatively cheaper foods for more expensive foods within a category.

To investigate whether there are any real price declines or just compositional changes, we look at some commonly consumed foods, including rice, potato, pork, beef, and lamb. As shown in Figure 10, prices still show obvious declines at the mandatory retirement age.

Table 6 reports the estimated changes in prices of these specific types of food. The estimates are all negative among the non–college education group, though not precisely

estimated. Table 7 reports the estimated changes in the quantities of these specific types of food. No significant declines are found for quantities among the non–college education group.

7 Food Consumption Index and Predicted Expenditures

The analysis so far focuses on either major food categories or a few selected types of food. To aggregate various types of food a household consumes, we follow the approach in Aguiar and Hurst (2005) to construct a food consumption index. We also follow their approach to construct a predicted expenditure holding prices fixed (assuming that the retirement-age individuals pay the same prices as prime-age working individuals).

The food consumption index is constructed as a weighted average of various quantities of food purchased and prices paid by a household. As shown in Aguiar and Hurst (2005), household consumption (quantities and prices) has a significant forecasting power for permanent income; therefore, one can compute the implied permanent income of the retirement-age households based on their consumption baskets.

We obtain the required weights by projecting the logarithm of permanent income on quantities and expenditure of various foods consumed by working-age household heads. Including food expenditure takes into account price heterogeneity across households. The unit of the consumption index is therefore log permanent income. Also included in the projection is a vector of taste controls and a smooth polynomial function of household head's age. That is, given permanent income I^{perm}, we can estimate the following equation to get the weights used later in constructing the consumption index for the retirement-age household heads:

$$\ln(I^{perm}) = \eta_0 + \eta_{\mathbf{Q}}\mathbf{Q} + \eta_{\mathbf{E}}\mathbf{E} + \eta_{\boldsymbol{\pi}}\boldsymbol{\pi} + m(X) + \epsilon, \tag{5}$$

where $\mathbf{Q}$ and $\mathbf{E}$ are vectors of food quantities and expenditure, respectively, $\boldsymbol{\pi}$ is a vector

of taste parameters, and $m(X)$ is a low-order polynomial of household head's age.

Permanent income is not directly observed for working-age individuals. In order to estimate Equation (5), we need to obtain $\ln(I^{perm})$ first. We first estimate a regression of log household income on household head's eduction, birth cohort, industry, and occupation controls, and the full set of occupation-industry interactions, using data on male household heads aged 25–45 who report working full time (16,772 individuals). We then take the fitted value as (estimated) permanent income and replace the unknown $\ln(I^{perm})$ in the above Equation (5) with the estimated value. $\mathbf{Q}$ and $\mathbf{E}$ consist of quantities and expenditures for 45 types of food, using the same sample. The vector of taste controls includes household size, household size squared, marital status, and province of residence. We next apply the estimated coefficients $\widehat{\eta}_{\mathbf{Q}}$ and $\widehat{\eta}_{\mathbf{E}}$ to quantities purchased and prices paid by the retirement-age household heads in our RD sample (19,887 non–college heads and 6,178 college heads).

We also obtain a predicted food expenditure for the retirement-age household heads using the estimated coefficients on food quantities $\eta_{\mathbf{Q}}$ in the above Equation (5). Recall that Equation (5) is estimated based on the sample of prime-age working household heads, so these weights can be taken as the average prices paid by these working households. Therefore, the predicted food expenditure assumes that the retirement-age households pay the same price as the prime-age working households.

Once we construct the food consumption index and predicted expenditure for our RD sample, we use those as our outcome variables in our RD analysis and test whether they decline significantly at retirement. We estimate both changes separately for the college and the non–college education groups. These estimates are given in Table 8.

The estimated changes in the predicated food expenditure are small and insignificant for both the college and the non–college education group. Therefore, had the retirement-age individuals paid the same prices for food as the prime-age working individuals, we would not see a decline in their food expenditure at retirement. In contrast, the consumption index is estimated to decline significantly by 2.2% among the non–college group, compared with the insignificant decline of 0.4% among the college education

group.[12] Since the only difference between the food consumption index and the predicted expenditure is that the latter holds prices fixed, the differential responses provide strong evidence that the observed decline in food expenditure is driven by price declines, not quantity declines. Therefore, consistent with what is documented in Aguiar and Hurst (2005), households, in response to forecastable income changes, smooth consumption but not necessarily expenditures, as predicted by the standard PIH augmented with home production.

8 Food Shopping Time Change at Retirement

Another caveat of the previous analysis is that price differences may reflect quality differences. Without detailed quality measures, we cannot really tell whether retirees pay lower prices because they buy a lower quality of the same goods or because they shop for bargains and thereby pay lower prices for the same quantity and quality of goods. For example, if retired households buy ordinary cuts of meat instead of premium cuts of meat, it would not be captured in our estimation.

To investigate whether retirees shop for bargains, we examine food shopping time changes upon retirement. We use data from the China Health and Nutrition Survey (CHNS), which contains information on food shopping time.[13] In particular, the CHNS asks, "During the past week, how much time (in minutes) did you spend per day, on average, to buy food for your household?"

The CHNS is an ongoing project that surveys households in nine provinces.[14] These nine provinces cover similarly geographically diverse areas as the UHS. So far, eight waves of data have been released for years 1989, 1991, 1993, 1997, 2000, 2004, 2006,

[12] Although we do not report them here, the estimated changes based on the full sample are insignificant, similar to the findings in Aguiar and Hurst (2005) based on the U.S. data. In particular, the estimated change in the consumption index is -0.005 with standard error 0.007, compared to -0.006 with standard error 0.02 as reported in Aguiar and Hurst (2005).

[13] CHNS is an open cohort, international collaborative project between the Carolina Population Center at the University of North Carolina at Chapel Hill and the National Institute of Nutrition and Food Safety at the Chinese Center for Disease Control and Prevention.

[14] These include Liaoning, Heilongjiang, Jiangsu, Shandong, Henan, Hubei, Hunan, Guangxi, and Guizhou.

and 2009, overlapping with our UHS sample period from 1997 to 2006. Ideally we would want to use data from the same years as our UHS data. However, because of the relatively small sample of the CHNS, we use all years of data. Here our main goal is to provide suggestive evidence on food shopping time changes at retirement among urban males rather than estimate how shopping time affects food prices causally. The CHNS contains households in both urban and rural areas. We limit our sample to the urban male household heads to make it comparable to our UHS data. The sample size is 4,742, including 4,240 non–college educated household heads and 502 college educated heads.

Figure 11(a) presents the age profile of the retirement rate in the CHNS sample. Similar to what we find in our UHS data, the retirement rate increases by about 30% at the mandatory retirement age, though at each age the retirement rate is slightly lower, which could be due to different provinces and time periods covered.

Figures 11(b) and 11(c) show the age profiles of food shopping time for the non–college and the college educated household heads, respectively. The average time spent on shopping for food clearly increases at the mandatory retirement age among the non–college education group, but not among their college-educated counterparts.

Table 9 reports the estimated changes in the probability of shopping for food at all last week and in the average time spent on shopping for food at retirement. Both are positive and significant among the non–college educated group. In particular, the average time spent on shopping for food is estimated to increase by about 22 minutes per day, and the probability of shopping for food last week is estimated to increase by 22.9% (The sample mean conditioning on shopping for food at all is 49.96 minutes per day and is 22.98 minutes per day unconditionally). No significant changes are found for the college education group. The probability and average time spent on shopping for food increase only among the non–college education group at retirement, consistent with the significant price decline among this group. Aguiar and Hurst (2007) show an inverse relationship between life cycle prices and shopping time. Therefore, the price decline observed among the low-education group is at least in part induced by retirees shopping for bargains.

One may wonder why college-educated individuals do not search more upon retirement, given that their opportunity cost of time also declines. One possible explanation may be that they value leisure more. The other may be that they do not need to search for lower prices, since they may have greater wealth.[15]

9 Time-Varying Effects of Retirement

We investigate immediate changes in average consumption at retirement, using the standard RD design. However, there may be variable delayed effects. That is, the long-run effect may be different from the short-run effect. Focusing entirely on immediate changes may under- or overestimate the full effects of retirement. For example, the age profiles of various consumption measures appear to show changes in slopes at the mandatory retirement age. Extrapolating the premandatory retirement age curve to postmandatory seems to suggest enlarged effects over time. To address this concern, this section extends the previous analysis to incorporating the time-varying effects of retirement. We first show that even if retirement effects change with time, the previous analysis using standard RD models still correctly identifies the immediate or short-run effect of retirement on consumption. We then discuss estimating the time-varying effects of retirement with our cross-sectional data (and hence without knowing at what age one is retired or how long one has been retired at the time of survey).

Let X^0 be a household head's retirement age minus 60, so, e.g., $X^0 = 1$ for someone who is retired at 61. Recall that X is one's true age at the time of survey minus 60, so $X - X^0$ measures how long one had been retired at the time of survey. Note that subtracting 60 is a free normalization.

Assume that household consumption depends on the household head's age, X, whether the head is retired or not, T, and other factors that change smoothly with age, υ. Further assume that retirement effects vary with one's retirement age, X^0, and how long one has been retired, $X - X^0$, where the latter captures time-varying retirement effects.

[15]There is no asset information in the UHS data set, so we could not directly investigate this.

The consumption model can be written as

$$Y = g(X) + h\left(X^0, X - X^0\right) T + \upsilon, \tag{6}$$

where $g(\cdot)$ and $h(\cdot)$ are smooth functions. For convenience of illustration, assuming a uniform kernel, we have the following local polynomial approximation of the above Equation (6):

$$Y = \sum\nolimits_{k=0}^{K} \lambda_k X^k + \sum\nolimits_{k=0}^{K} \sum\nolimits_{j=0}^{k} \tau_{kj} \left(X^0\right)^j \left(X - X^0\right)^{k-j} T + \varpi. \tag{7}$$

The average retirement effect at age 60 for those who retire at this mandatory retirement age (those with $X = X^0 = 0$) is τ_{00}.

It is easy to show that the immediate effect of retirement on consumption at age 60 is still identified by the standard RD estimator:

$$\tau_{00} = \frac{\lim_{x \to 0^+} E\left[Y \mid X = x\right] - \lim_{x \to 0^-} E\left[Y \mid X = x\right]}{\lim_{x \to 0^+} E\left[T \mid X = x\right] - \lim_{x \to 0^-} E\left[T \mid X = x\right]}. \tag{8}$$

Intuitively, $\sum_{k=1}^{K} \sum_{j=0}^{k} \tau_{kj} \left(X^0\right)^j \left(X - X^0\right)^{k-j}$, or, more generally, terms involving X_0 and $X - X_0$ in Equation (6), are smooth at the (normalized) mandatory retirement age $X = 0$, and so they drop in the difference in the numerator. Proof is provided in Appendix III.

Taking a local linear approximation (i.e., $K = 1$) of the true consumption model, we have

$$Y = \lambda_0 + \lambda_1 X + \tau_{00} T + \tau_{10} \left(X - X^0\right) T + \tau_{11} X^0 T + \varpi. \tag{9}$$

Assuming further that $\tau_{10} = \tau_1$, that is, the retirement effect heterogeneity in retirement age is assumed to be the same as the time-varying effect of retirement, then Equation (9) reduces to $Y = \lambda_0 + \lambda_1 X + \tau_{00} T + \tau_{10} XT + \varpi$. For example, the retirement effect at age 61 is the same for those who retire at age 61 and those who retire at 60 but are

now age 61.[16]

We estimate the immediate effect τ_{00} and the time-varying effect τ_{10} in Equation (9), using the mandatory retirement dummy D as an IV for retirement status T. The estimates are presented in Tables 10 and 11. We focus on the prices and quantities of major food categories to obtain relatively precise estimates. Incorporating the time-varying effect of retirement does not significantly change the estimated short-run effects. For both prices and quantities, the estimated coefficients of XT are small, so the retirement effects do not change much over time. For example, for food prices, the estimates among the low-education group range from 0 to -0.3%, so food prices decline slightly more the longer one has been retired. The estimates are mostly small and insignificant among the college education group. For food quantities, the estimates among the non–college education group are about -2%, compared with the initial increases of over 10% on average. For this group, the initial increase in the quantities of food purchased for home cooking has a small decline over time. Overall, the immediate changes in food prices and quantities upon retirement seem to persist over time.

10 Validity of the RD Design

The validity of the RD design we adopt relies crucially on the assumption that individuals in our sample do not systematically sort around the mandatory retirement age and that covariates do not change discontinuously in response to household heads' retirement. For example, if retirees systematically move in with adult children or if their spouses retire jointly with them, then the estimated retirement effects would be confounded.

Following the standard practice (see, e.g., McCrary [2008] and Lee and Lemieux [2010]), we examine the smoothness of the density of household head's age and the smoothness of predetermined covariates, including household head's education level, marital status, spouse's retirement status, and household size. Although we condition

[16]Without this restriction, in general with only cross-section data like ours, one cannot separate how the retirement effect changes with time for those "compliers" who retire at age 60 from retirement effect heterogeneity in the retirement age.

on household head's education, marital status, and household size in our analysis, a discontinuity in these variable means at the retirement age threshold would indicate incomparability of households just under and just above the retirement age and hence invalidity of our estimates.

Figure 12 presents the empirical density of household head's age, or the fraction of observations in each age cell and the age profile of covariate means. There are no obvious discontinuities or bunching around the mandatory retirement age of 60.

We formally test the smoothness of the density of age and these covariate means. We use the empirical density, or the fraction of observations at each age, as the dependent variable and then regress this density on a polynomial function of age and the full set of interactions between this polynomial function and the binary indicator for being 60 or older. The coefficient of this binary indicator represents the potential discontinuity in the density of age at the age threshold.

To test the smoothness of predetermined covariate means, we do parallel RD estimates using these covariates as dependent variables. False significant retirement effects on these predetermined covariates would indicate discontinuities in these covariate means at the RD threshold. Note that in all these estimates, we take into account the fact that we use rounded age in years instead of true age as the running variable, and hence we do similar bias corrections as those described previously. The test results are presented in Table 12. None of the estimates are statistically significant, so the density of household head's age and covariate means are smooth, which supports the validity of our RD analysis.

11 Conclusion

This study provides the first direct causal evidence on consumption changes upon retirement, in addition to expenditure changes. Mandatory retirement in China provides a unique quasi-experimental setting in which one can investigate consumption declines at retirement.

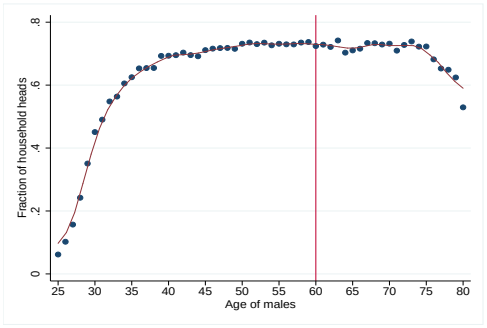
Based on detailed consumption diary information from the UHS, we show that food expenditure declines when household heads retire in China, particularly among household heads without a college education. We further show that the observed food expenditure decline is driven by reduced spending on eating out and by food price declines, instead of by food quantity declines.

One key issue is whether the lower prices paid for food by those households are due to the households shopping for bargains or for lower quality food. Although we cannot directly examine quality changes at retirement, we show that average time spent on shopping for food increases significantly at retirement among the non–college education group, suggesting that the price declines observed for this group are induced at least in part by shopping for bargains.

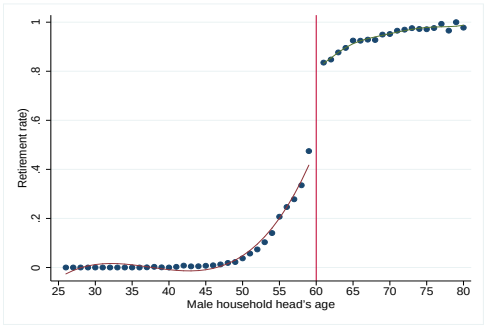
We additionally show that the one-off changes in food prices and quantities at retirement persist over time, and that there are no significant delayed effects. Overall, the evidence suggests that there is a change in time use rather than a real consumption decline among retirees. The observed food expenditure decline at retirement is largely consistent with a life-cycle model augmented with home production. As highlighted in Aguiar and Hurst (2005), given home production, expenditures on food are poor proxies for actual household consumption and mask the extent to which individuals smooth consumption in practice.

Food consumption consists of about 45% of the total consumption in our urban household sample. Over three-fourths of the household heads have less than a college education in urban China among the cohorts in our sample. Investigating the impact of retirement on food consumption, particularly among the relatively low-education group, has broad social and policy impacts in China.

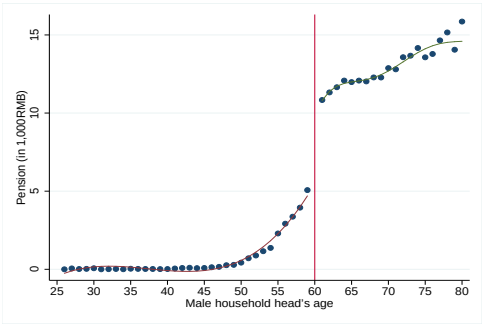
12 Appendix I



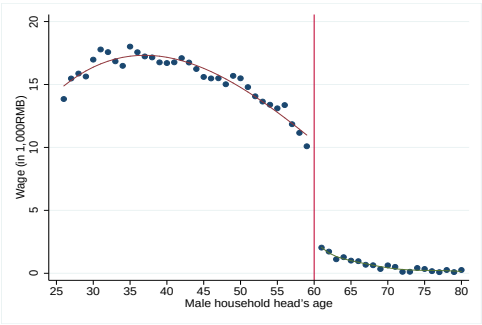
(a) Age distribution of household heads



(b) Retirement rate

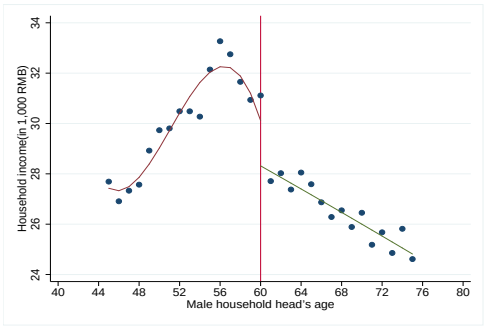


(c) Pension income

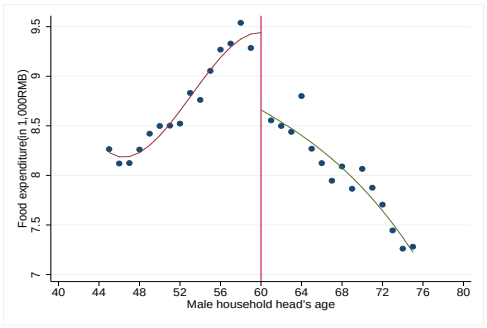


(d) Wage income

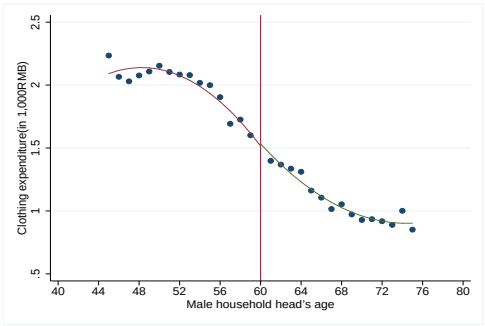
Figure 1: Age Profiles of Factions of Household Heads, Heads' Retirement Rate, Wage, and Pension Income, UHS 1997–2006



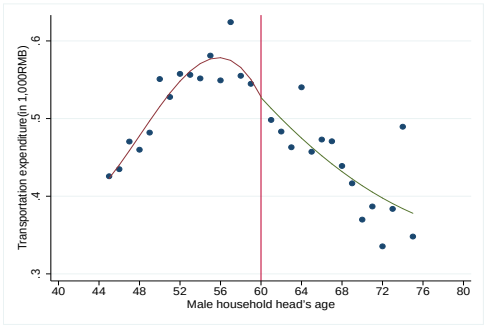
(a) Income



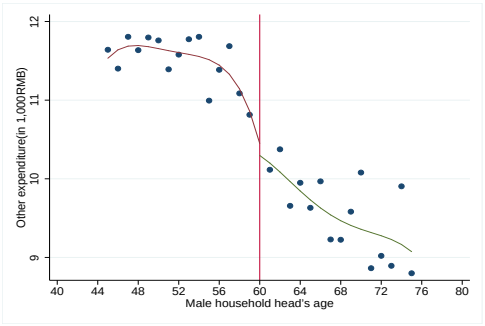
(b) Food



(c) Clothing

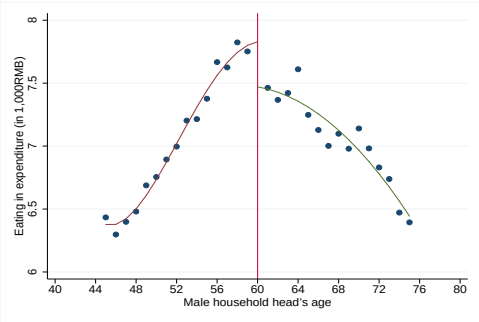


(d) Transportation

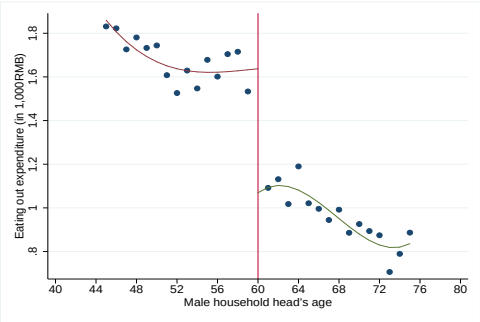


(e) Other

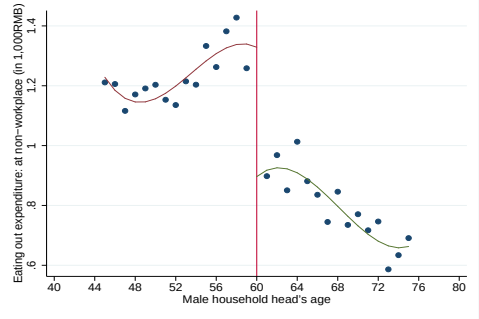
Figure 2: Age Profiles of Household Income and Expenditure: Male Household Heads, UHS 1996–2007



(a) Food at home

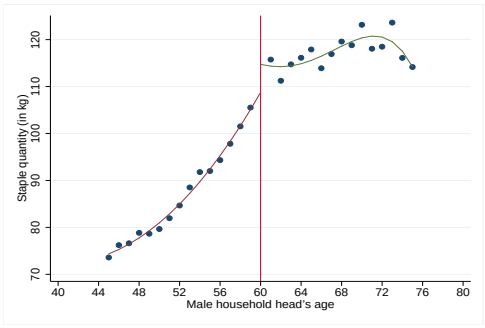


(b) Eating out

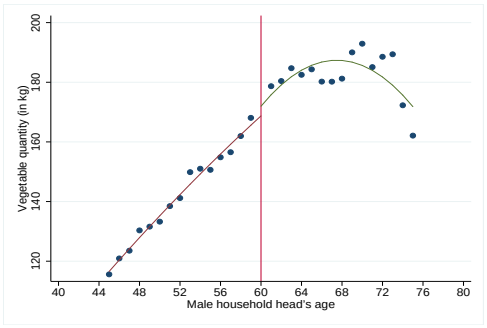


(c) Eating out: nonworkplace restaurants

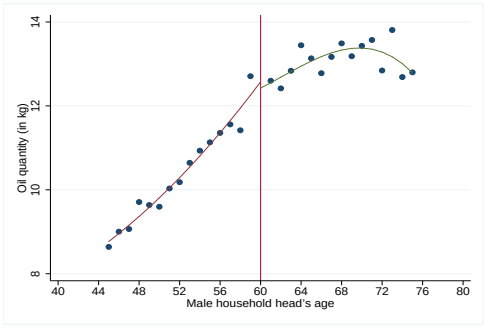
Figure 3: Age Profiles of Food Expenditure: Male Household Heads, UHS 1997–2006



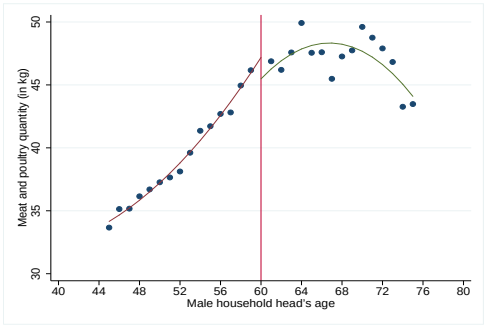
(a) Staple quantity



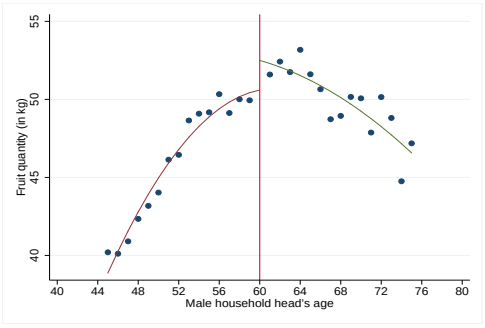
(b) Vegetable quantity



(c) Oil and fat quantity

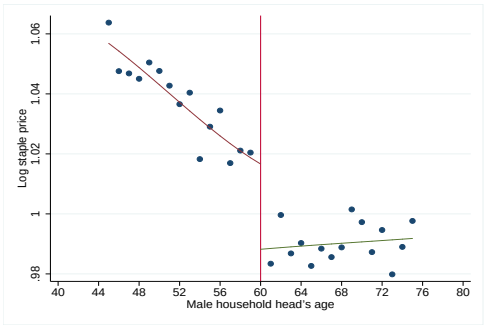


(d) Meat and poultry quantity

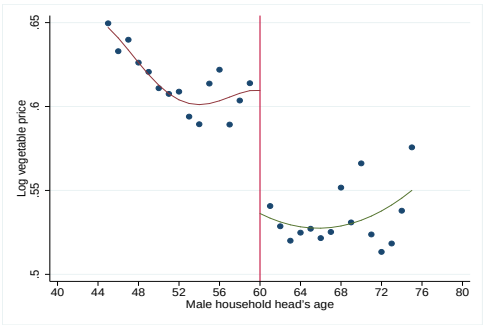


(e) Fruit quantity

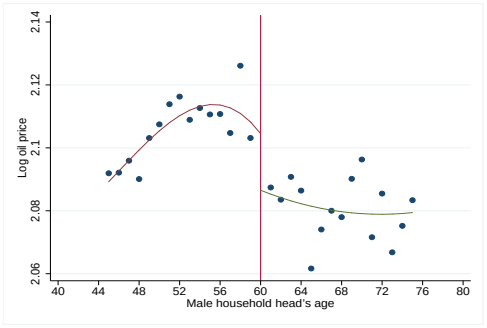
Figure 4: Age Profiles of Food Quantities Purchased in Different Categories: Male Household Heads, UHS 1997–2006



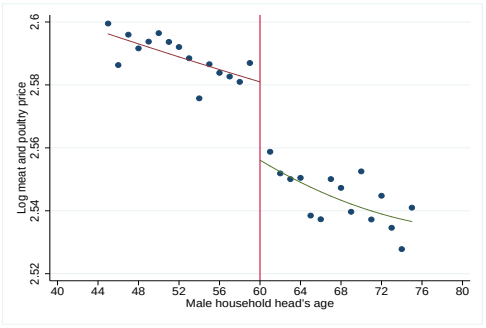
(a) Staple price



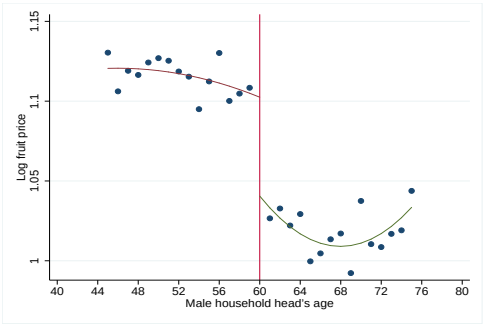
(b) Vegetable price



(c) Oil and fat price

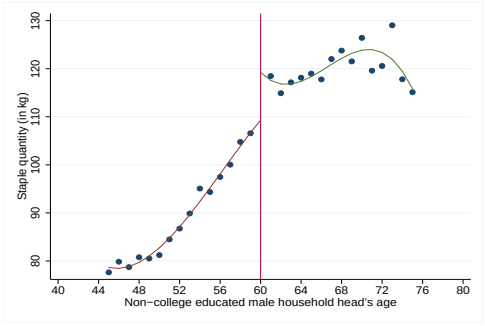


(d) Meat and poultry price

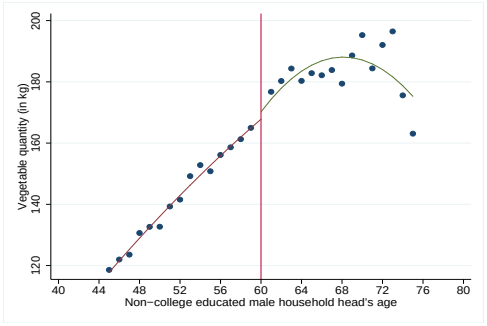


(e) Fruit price

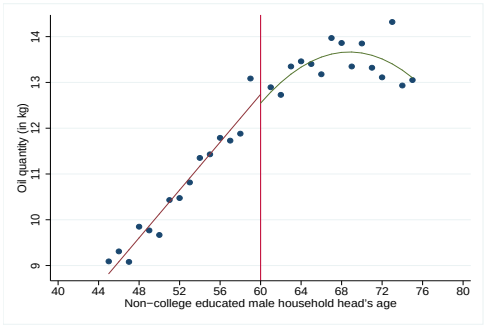
Figure 5: Age Profiles of Prices Paid for Different Categories of Food: Male Household Heads, UHS 1997–2006



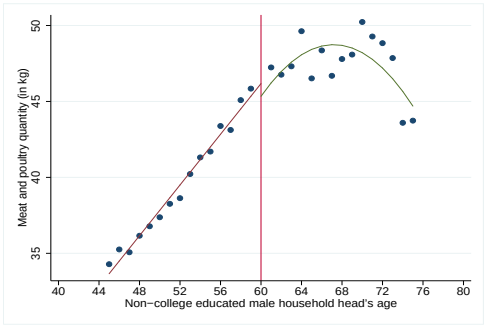
(a) Staple quantity



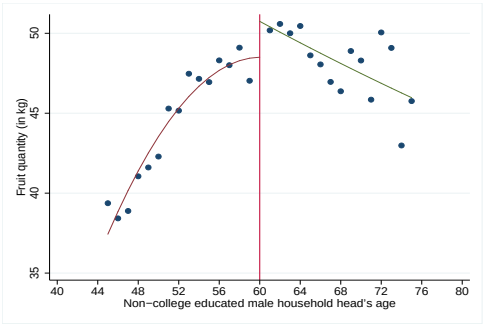
(b) Vegetable quantity



(c) Oil and fat quantity

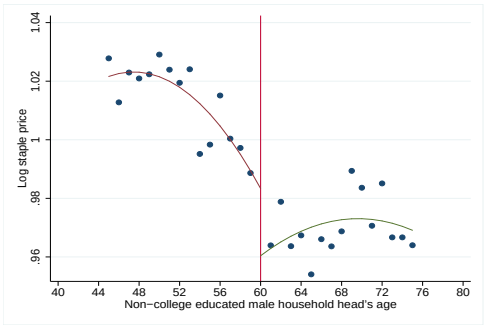


(d) Meat and poultry quantity

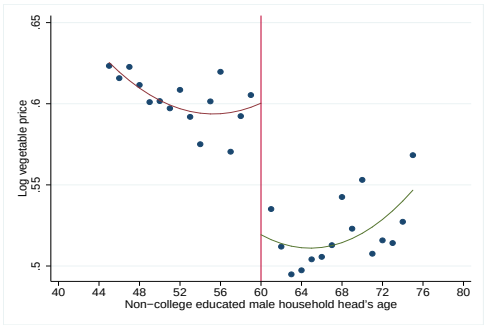


(e) Fruit quantity

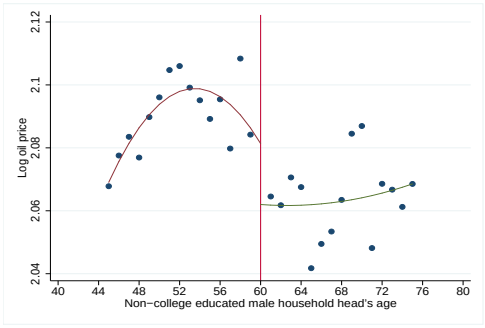
Figure 6: Age Profiles of Quantities Purchased of Different Categories of Food: Non–college Educated Male Household Heads, UHS 1997–2006



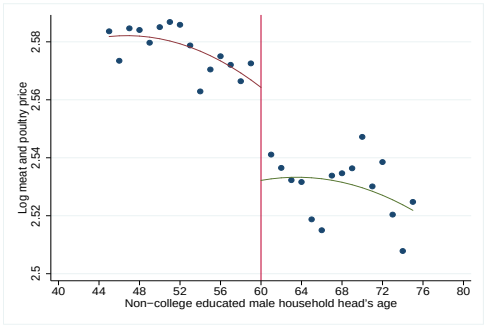
(a) Staple price



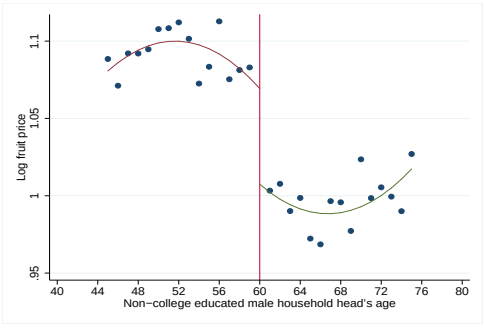
(b) Vegetable price



(c) Oil and fat price

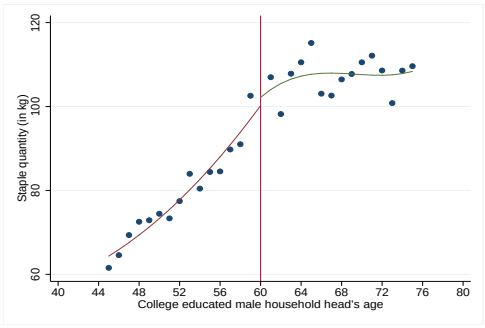


(d) Meat and poultry price

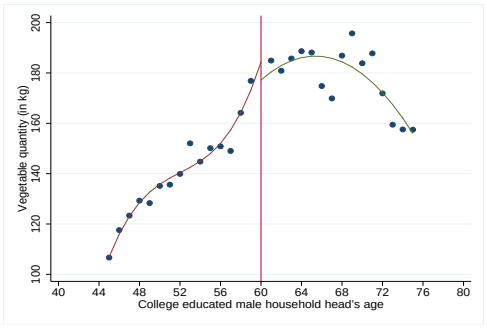


(e) Fruit price

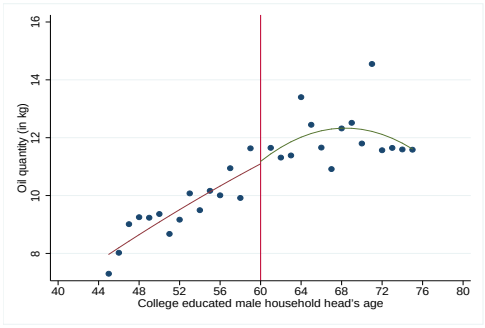
Figure 7: Age Profiles of Prices Paid for Different Categories of Food: Non–college Educated Male Household Heads, UHS 1997–2006



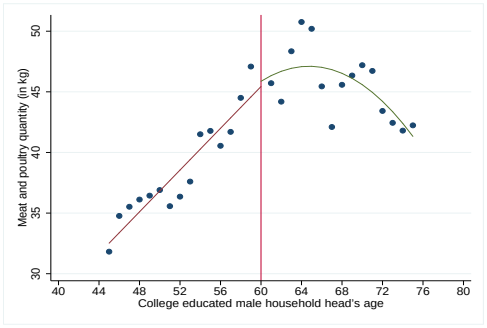
(a) Staple quantity



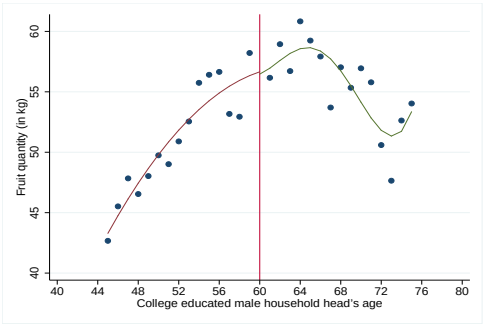
(b) Vegetable quantity



(c) Oil and fat quantity

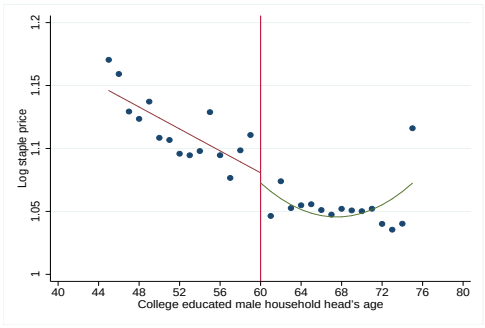


(d) Meat and poultry quantity

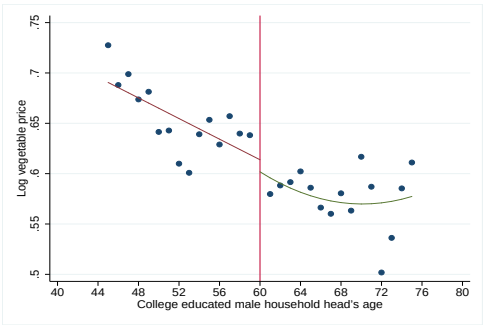


(e) Fruit quantity

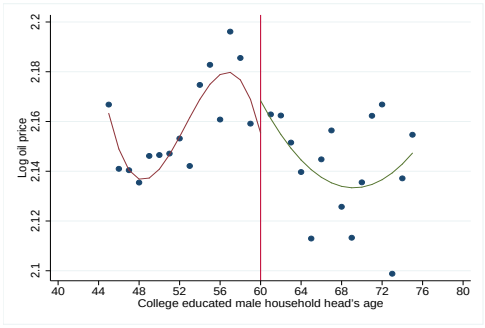
Figure 8: Age Profiles of Food Quantities Purchased in Different Categories: College Educated Male Household Heads, UHS 1997–2006



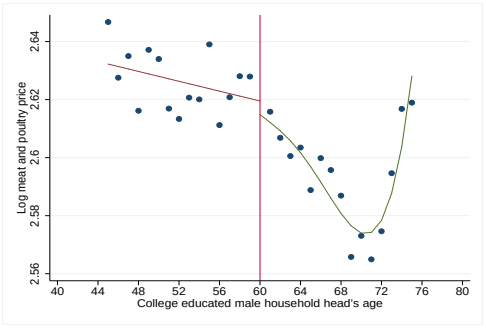
(a) Staple price



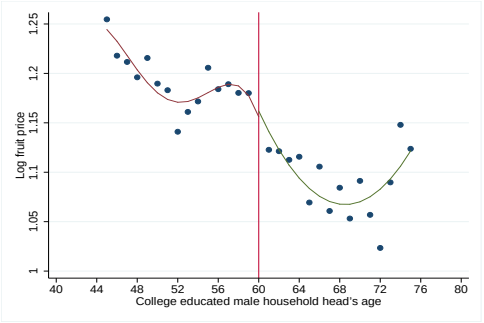
(b) Vegetable price



(c) Oil price

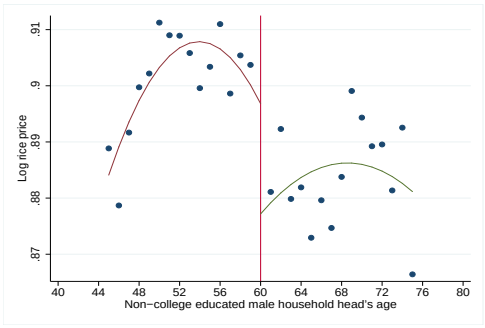


(d) Meat and poultry price

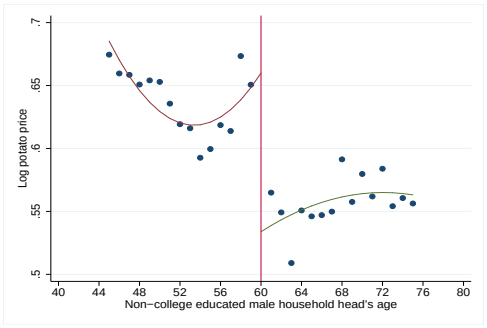


(e) Fruit price

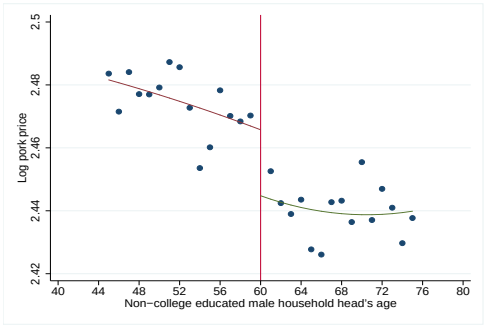
Figure 9: Age Profiles of Prices Paid for Different Categories of Food: College Educated Male Household Heads, UHS 1997–2006



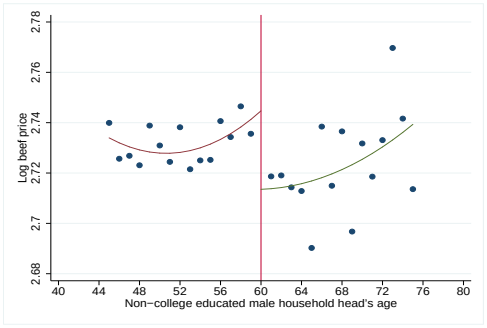
(a) Rice price



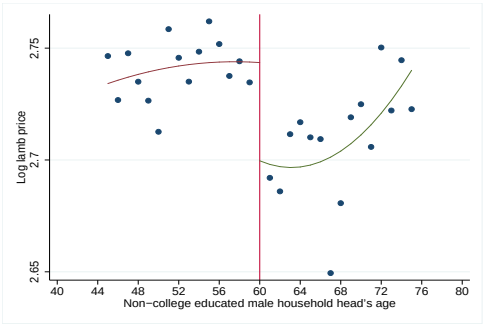
(b) Potato price



(c) Pork price

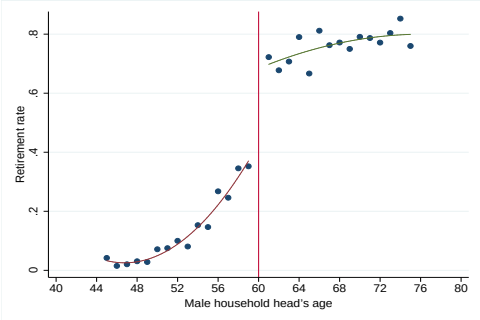


(d) Beef price

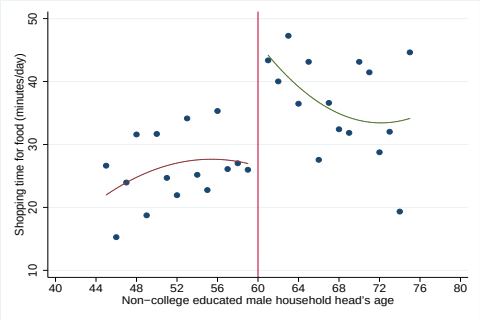


(e) Lamb price

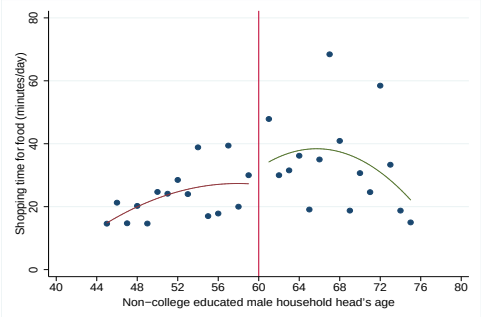
Figure 10: Age Profiles of Prices Paid for Food: Non–college Educated Male Household Heads, UHS 1997–2006



(a) Retirement rate in the CHNS sample

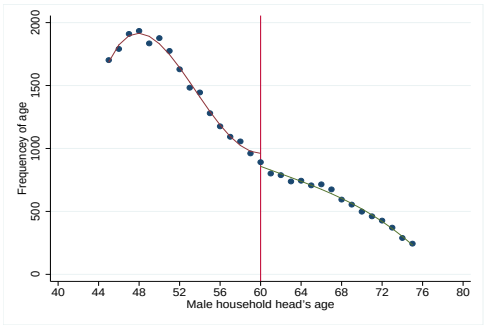


(b) Average daily shopping time for food: Non-college educated household heads

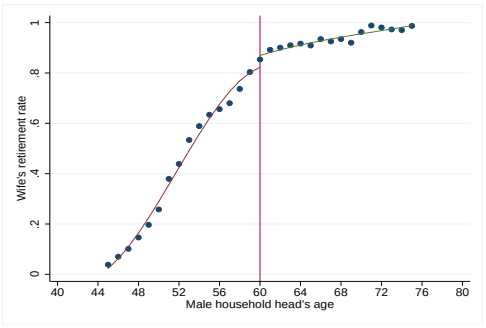


(c) Average daily shopping time for food: College educated household heads

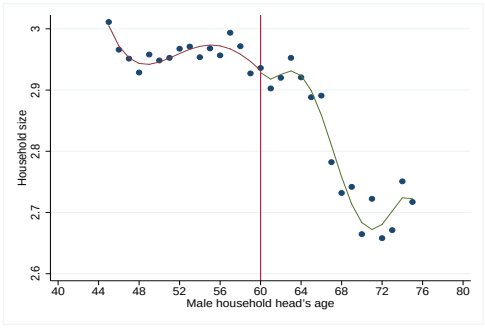
Figure 11: Age Profiles of Retirement Rate, Shopping Time for Food: Male Household Heads, CHNS 1989–2009



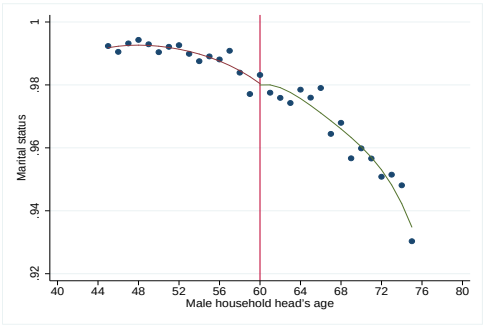
(a) Frequency of age



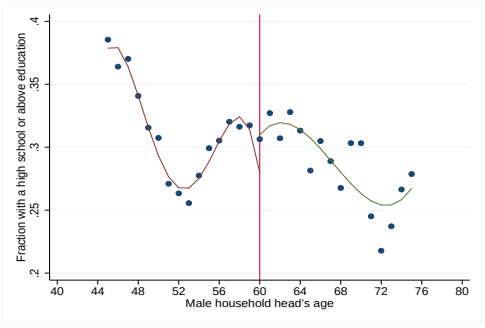
(b) Wife's retirement rate



(c) Household size



(d) Marital status



(e) Household head's education

Figure 12: Age Profiles of the Empirical Density of Age and Covariate Means: Male Household Heads, UHS 1997–2006

Table 1a The Retirement Rate Increase at the Mandatory Retirement Age of 60 for Male Household Heads

2nd order polynomial	0.290	0.288	0.280	0.307	0.301	0.294			
	(0.037)***	(0.036)***	(0.036)***	(0.020)***	(0.020)***	(0.020)***			
3rd order polynomial	0.285	0.294	0.295	0.277	0.275	0.269	0.295	0.292	0.285
	(0.085)**	(0.084)**	(0.084)**	(0.036)***	(0.035)***	(0.035)***	(0.020)***	(0.020)***	(0.020)***
4th order polynomial				0.301	0.306	0.307	0.270	0.269	0.254
				(0.068)***	(0.068)***	(0.067)***	(0.033)***	(0.033)***	(0.033)***
Year fixed effects	N	Y	Y	N	Y	Y	N	Y	Y
Province fixed effects	N	Y	Y	N	Y	Y	N	Y	Y
Year-province fixed effects	N	N	Y	N	N	Y	N	N	Y
Demographic controls	N	N	Y	N	N	Y	N	N	Y
Bandwidth	6	6	6	10	10	10	15	15	15

NOTE: Male household heads, UHS 1997–2006; demographic controls include household head's education and marital status, family size, and family size squared; robust standard errors are in parentheses; * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 1b The Retirement Rate Increase at the Mandatory Retirement Age of 60 for Male Household Heads, without Correcting for Discrete Age Bias

2nd order polynomial	0.251	0.249	0.241	0.278	0.271	0.264			
	(0.033)***	(0.033)***	(0.033)***	(0.019)***	(0.019)***	(0.018)***			
3rd order polynomial	0.133	0.137	0.130	0.241	0.238	0.232	0.262	0.258	0.252
	(0.067)**	(0.066)**	(0.066)**	(0.031)***	(0.031)***	(0.030)***	(0.018)***	(0.018)***	(0.018)***
4th order polynomial				0.188	0.195	0.191	0.236	0.233	0.228
				(0.052)***	(0.052)***	(0.052)***	(0.028)***	(0.027)***	(0.027)***
Year fixed effects	N	Y	Y	N	Y	Y	N	Y	Y
Province fixed effects	N	Y	Y	N	Y	Y	N	Y	Y
Year-province fixed effects	N	N	Y	N	N	Y	N	N	Y
Demographic controls	N	N	Y	N	N	Y	N	N	Y
Bandwidth	6	6	6	10	10	10	15	15	15

NOTE: Male household heads, UHS 1997–2006; demographic controls include household head's education and marital status, family size, and family size squared; robust standard errors are in parentheses; * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 2 Retirement Effects on Different Categories of Expenditure

	Noncollege group			College group		
Food	-0.041	-0.053	-0.047	0.031	0.045	0.052
	(0.027)	(0.025)**	(0.025)**	(0.039)	(0.038)	(0.038)
Clothes	0.222	0.191	0.198	0.028	0.072	0.079
	(0.182)	(0.176)	(0.178)	(0.189)	(0.192)	(0.192)
Transport	0.021	0.023	0.060	0.158	0.159	0.192
	(0.210)	(0.212)	(0.217)	(0.258)	(0.258)	(0.258)
Other	0.115	0.214	0.276	0.053	0.103	0.129
	(0.233)	(0.228)	(0.228)	(0.258)	(0.255)	(0.247)
Year fixed effects	Y	Y	Y	Y	Y	Y
Province fixed effects	Y	Y	Y	Y	Y	Y
Demographic controls	N	Y	Y	N	Y	Y
Year-province fixed effects	N	N	Y	N	N	Y

NOTE: Male household heads, UHS 1997–2006; demographic controls include household head's education and marital status, household size, and household size squared; robust standard errors are in parentheses; * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 3 Retirement Effects on Different Categories of Food Expenditure

	Noncollege group			College group		
Food at home	-0.026	-0.033	-0.031	0.025	0.029	0.039
	(0.025)	(0.024)	(0.024)	(0.036)	(0.035)	(0.035)
Food out (total)	-0.229	-0.250	-0.232	-0.003	0.020	0.047
	(0.070)***	(0.068)***	(0.069)***	(0.116)	(0.118)	(0.117)
Food out (nonworkplace	-0.276	-0.300	-0.289	0.092	0.101	0.125
restaurants)	(0.069)***	(0.067)***	(0.068)***	(0.128)	(0.129)	(0.130)
Year fixed effects	Y	Y	Y	Y	Y	Y
Province fixed effects	Y	Y	Y	Y	Y	Y
Demographic controls	N	Y	Y	N	Y	Y
Year-province fixed effects	N	N	Y	N	N	Y

NOTE: Male household heads, UHS 1997–2006; demographic controls include household head's education and marital status, household size, and household size squared; food-out expenditure has a non-negligible fraction of zeros and so is not logged in estimation; percentage changes are reported here by converting the estimated average level changes; robust standard errors are in parentheses; * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 4 Retirement Effects on Average Prices for Different Categories of Food

	Noncollege group		College group	
Staple	-0.013	(0.011)	-0.013	(0.020)
Vegetable	-0.075	(0.016)***	0.006	(0.024)
Oil	-0.041	(0.012)***	0.007	(0.023)
Meat	−0.028	(0.009)***	0.007	(0.014)
Meat and poultry	-0.025	(0.009)***	0.004	(0.014)
Fruit	-0.106	(0.018)***	-0.022	(0.026)

NOTE: Male household heads, UHS 1997–2006; all estimates control for year fixed effect, province fixed effect, year-province fixed effects, a low-order polynomial of household head's age, head's education and marital status, household size, and household size squared; robust standard errors are in parentheses; * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 5 Retirement Effects on Quantities Purchased of Different Categories of Food

	Noncollege group		College group	
Staple	0.151	(0.078)**	0.055	(0.044)
Vegetable	0.101	(0.026)***	-0.045	(0.138)
Oil	0.109	(0.039)***	0.062	(0.067)
Meat	0.089	(0.029)***	0.053	(0.043)
Meat and poultry	0.081	(0.028)***	0.033	(0.041)
Fruit	-0.008	(0.035)	-0.008	(0.049)

NOTE: Male household heads, UHS 1997–2006; all estimates control for year fixed effect, province fixed effect, year-province fixed effects, a low-order polynomial of household head age, household head's education and marital status, household size, and household size squared; robust standard errors are in parentheses; * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 6 Retirement Effects on Prices of Food

	Noncollege group		College group	
Rice	-0.015	(0.022)	-0.016	(0.015)
Potato	-0.125	(0.061)**	-0.044	(0.034)
Pork	-0.007	0.022	0.017	(0.012)
Beef	-0.038	0.046	0.000	(0.048)
Lamb	-0.096	(0.083)	0.028	(0.032)

NOTE: Male household heads, UHS 1997–2006; all estimates control for year fixed effect, province fixed effect, year-province fixed effects, a low-order polynomial of household head's age, household head's education and marital status, household size, and household size squared; robust standard errors are in parentheses; * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 7 Retirement Effects on Quantities of Food

	Noncollege group		College group	
Rice	0.102	(0.129)	0.096	(0.086)
Potato	0.181	(0.057)***	0.085	(0.084)
Pork	0.090	(0.038)**	0.014	(0.059)
Beef	-0.032	(0.080)	0.051	(0.404)
Lamb	-0.060	(0.080)	-0.186	(0.440)

NOTE: Male household heads, UHS 1997–2006; all estimates control for year fixed effect, province fixed effect, year-province fixed effects, a low-order polynomial of household head's age, household head's education and marital status, household size, and household size squared; robust standard errors are in parentheses; * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 8 Retirement Effects on Food Consumption Index and Predicted Food Expenditure

	Noncollege group		College group	
log of food consumption index	-0.022	(0.008)***	-0.004	(0.011)
log of predicted food expenditure	0.005	(0.011)	0.008	(0.014)

NOTE: Male household heads, UHS 1997–2006; food consumption index is in the unit of permanent income, taking into account both quantities and prices of various foods a household consumed, while predicated food expenditure holds food prices fixed; detailed construction of both are in the main text; robust standard errors are in parentheses; * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 9 Retirement Effects on Time Spent Shopping for Food

	Noncollege group		College group	
Time spent on shopping for food	22.04	(6.944)***	8.267	(15.96)
Whether shopping for food last week (0/1)	0.229	(0.077)***	0.031	(0.169)

NOTE: Male household heads, CHNS 1989–2009; all estimates control for province, year and province-year fixed effects; robust standard errors are in parentheses; the average shopping time for the noncollege sample is 22.98 minutes per day and is 49.96 minutes per day among those with positive time, while the average shopping time for the college sample is 26.66 minutes per day and is 45.53 minutes per day among those with positive time. * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 10 Time-Varying Effects of Retirement on Food Prices

	Noncollege group		College group	
	Retire	Retire*(Age-60)	Retire	Retire*(Age-60)
Staple	-0.016	0.002	-0.013	0.004
	(0.011)	(0.001)***	(0.020)	(0.001)***
Vegetable	-0.075	0.000	0.006	0.001
	(0.016)***	(0.001)	(0.024)	(0.002)
Oil	-0.040	-0.001	0.005	-0.002
	(0.012)***	(0.001)*	(0.023)	(0.002)
Meat	-0.026	-0.001	0.006	-0.001
	(0.009)***	(0.000)***	(0.013)	(0.001)
Meat & poultry	-0.023	-0.001	0.004	-0.001
	(0.009)***	(0.000)***	(0.014)	(0.001)
Fruit	-0.102	-0.003	-0.022	0.002
	(0.017)***	(0.001)***	(0.026)	(0.002)

NOTE: Male household heads, UHS 1997–2006; all estimates control for year fixed effect, province fixed effect, year-province fixed effects, a low-order polynomial of household heads' age, household head's education and marital status, household size, and household size squared; Robust standard errors are in parentheses; * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 11 Time-Varying Effects of Retirement on Food Quantities

	Noncollege group		College group	
	Retire	Retire*(Age-60)	Retire	Retire*(Age-60)
Staple	0.204	-0.031	0.052	−0.019
	(0.079)**	(0.006)***	(0.044)	(0.003)***
Vegetable	0.124	-0.017	-0.037	-0.065
	(0.026)***	(0.001)***	(0.144)	(0.042)
Oil	0.132	-0.017	0.060	-0.013
	(0.040)***	(0.002)***	(0.067)	(0.005)***
Meat	0.110	-0.016	0.050	-0.017
	(0.030)***	(0.001)***	(0.043)	(0.003)***
Meat & poultry	0.101	-0.015	0.030	−0.017
	(0.029)***	(0.001)***	(0.041)	(0.003)***
Fruit	0.022	-0.023	-0.011	-0.023
	0.036	(0.002)***	(0.049)	(0.003)***

NOTE: Male household heads, UHS 1997–2006; all estimates control for year fixed effect, province fixed effect, year-province fixed effects, a low-order polynomial of household head age, head's education, marital status, household size, and household size squared; robust standard errors are in parentheses; * significant at the 10% level, ** significant at the 5% level, *** significant at the 1% level.

Table 12 Estimated Discontinuities in the Density of Age and Retirement Effects on Covariate Means

Density of age	-0.005	(0.003)	Married	0.003	(0.008)
College education	-0.055	(0.057)	Wife retired	-0.106	(0.221)
High school education	-0.012	(0.063)	Household size	0.026	(0.056)

NOTE: Male household heads, UHS 1997–2006; the regression for the density of age controls for a smooth polynomial age function, while all other GMM IV estimates additionally control for year fixed effects, province fixed effects, and province-year fixed effects. robust standard errors are in parentheses.

13 Appendix II

This appendix illustrates the bias incurred when using age in years as a running variable and describes a bias correction procedure. For simplicity, consider the following quadratic regression for retirement

$$T = \sum_{j=0}^{J} a_j X^j + \sum_{j=0}^{J} b_j X^j D + v. \tag{10}$$

Recall that X is the (unobserved) true continuous age. Let $\widetilde{X}$ be the reported age in years minus 60, which is the exact age rounded down to the nearest integer, so that $X = \widetilde{X} + e$, where e is the difference between the true age and the rounded age in years, or the rounding error. Assuming that one's birth-date e is independent of his integer age X, we have

$$\begin{aligned} T &= \textstyle\sum_{j=0}^{2} a_j (\widetilde{X} + e)^j + \sum_{j=0}^{2} b_j (\widetilde{X} + e)^j D + v \\ &= (a_0 + a_1\mu_1 + a_2\mu_2) + (a_1 + 2a_2\mu_1)\,\widetilde{X} + a_2\widetilde{X}^2 \\ &+ (b_0 + b_1\mu_1 + b_2\mu_2)\,D + (b_1 + 2b_2\mu_1)\,\widetilde{X}D + b_2\widetilde{X}^2 D + w \\ &= \textstyle\sum_{j=0}^{2} \alpha_j \widetilde{X}^j + \sum_{j=0}^{2} \beta_j \widetilde{X}^j D + w, \end{aligned} \tag{11}$$

where $\mu_j = E(e^j)$ for $j = 0, 1, 2$ is the jth raw moments of the birth-date distribution within a year, $\alpha_j \equiv \sum_{k=j}^{2} \binom{k}{j} a_k \mu_{k-j}$ and $\beta_j \equiv \sum_{k=j}^{2} \binom{k}{j} b_k \mu_{k-j}$ for $j = 0, 1, 2$, and $w = T - E(T \mid \widetilde{X})$. Assuming that birth dates within a year are uniformly distributed, so that the rounding error e has a uniform distribution between 0 and 1, then the jth moment is $\mu_j = 1/(j+1)$.[17]

If one estimates a polynomial regression of T on the rounded age in years $\widetilde{X}$, or Equation (11), then the discontinuity in the retirement rate is $\beta_0 \equiv b_0 + b_1\mu_1 + b_2\mu_2$, which in general would not equal the true change b_0, unless b_1 and b_2 are both zero,

[17]There exists evidence of small but statistically significant seasonal departures from uniformity in the distribution of births within a year. However, this seasonal variation appears to have very little impact on the lower-order moments. Alternatively, one could estimate those moments using a second source of data where one observes age in days. However, we are not aware of any comparable data sets that have age in days.

given that $\mu_1 \neq 0$ and $\mu_2 \neq 0$.

To obtain a consistent estimate of the true change in the mean consumption, one can use β_j for $j = 0, 1, 2$ in the discrete data regression (11) and the rounding error moments, μ_j, to back out the true coefficients, b_j. Plug in these moments and then solve the system of equations $\beta_j = \sum_{k=j}^{2} \binom{k}{j} b_k \mu_{k-j}$ for b_j, where $j = 0, 1, 2$. Then the true consumption change at the mandatory retirement age is $b_0 = \beta_0 - 1/2\beta_1$. The standard errors can be obtained by the Delta method. A general formula that works for any order of polynomial can be found in Dong (2015).

14 Appendix III

This appendix shows that the standard RD estimation still provides correct estimates of the immediate effects of retirement on outcomes, even when there exist variable delayed effects of retirement, and when the retirement effects are heterogeneous in retirement age.

Let the potential treatment status be T_d if $D = d$ for $d = 0, 1$. Define four types of individuals as events in a common probability space $(\Omega, \mathcal{F}, P)$, as in Angrist, Imbens and Rubin (1996): always takers (denoted as A) are individuals with $T_1 = T_0 = 1$; never takers (N) are individuals with $T_1 = T_0 = 0$; compliers (C) are individuals with $T_1 = 1$ and $T_0 = 0$; and defiers (E) are individuals with $T_1 = 0$ and $T_0 = 1$. Assume that there are no defiers.

Given $Y = \sum_{k=0}^{K} \lambda_k X^k + \tau_{00} T + \sum_{k=1}^{K} \sum_{j=0}^{k} \tau_{kj} \left(X^0\right)^j \left(X - X^0\right)^{k-j} T + \varpi$, the following shows that $\sum_{k=1}^{K} \sum_{j=0}^{k} \tau_{kj} \left(X^0\right)^j \left(X - X^0\right)^{k-j}$ is smooth and hence will drop in

the mean difference in Y at $X = 0$.

$$\begin{aligned}
&\lim_{x\to 0^+} E\left[\sum_{k=1}^{K}\sum_{j=0}^{k}\lambda_{kj}\left(X^0\right)^j\left(X-X^0\right)^{k-j}T \mid X=x\right] \\
=\ &\lim_{x\to 0^+} E\left[\sum_{k=1}^{K}\sum_{j=0}^{k}\lambda_{kj}\left(X^0\right)^j\left(X-X^0\right)^{k-j} \mid X=x, T=1\right]\Pr\left(T=1 \mid X=x\right) \\
=\ &E\left[\sum_{k=1}^{K}\sum_{j=0}^{k}\lambda_{kj}\left(X^0\right)^j\left(-X^0\right)^{k-j} \mid X=0, A\right]\Pr\left(A \mid X=0\right) \\
&+E\left[\sum_{k=1}^{K}\sum_{j=0}^{k}\lambda_{kj}\left(X^0\right)^j\left(-X^0\right)^{k-j} \mid X=0, C\right]\Pr\left(C \mid X=0\right),
\end{aligned}$$

and similarly

$$\begin{aligned}
&\lim_{x\to 0^-} E\left[\sum_{k=1}^{K}\sum_{j=0}^{k}\tau_{kj}\left(X^0\right)^j\left(X-X^0\right)^{k-j}T \mid X=x\right] \\
=\ &\lim_{x\to 0^-} E\left[\sum_{k=1}^{K}\sum_{j=0}^{k}\tau_{kj}\left(X^0\right)^j\left(X-X^0\right)^{k-j} \mid X=x, T=1\right]\Pr\left(T=1 \mid X=x\right) \\
=\ &E\left[\sum_{k=1}^{K}\sum_{j=0}^{k}\tau_{kj}\left(X^0\right)^j\left(-X^0\right)^{k-j} \mid X=0, A\right]\Pr\left(A \mid X=0\right).
\end{aligned}$$

Then we have

$$\begin{aligned}
&\lim_{x\to 0^+} E\left[\sum_{k=1}^{K}\sum_{j=0}^{k}\lambda_{kj}\left(X^0\right)^j\left(X-X^0\right)^{k-j}T \mid X=x\right] \\
&-\lim_{x\to 0^-} E\left[\sum_{k=1}^{K}\sum_{j=0}^{k}\lambda_{kj}\left(X^0\right)^j\left(X-X^0\right)^{k-j}T \mid X=x\right] \\
=\ &E\left[\sum_{k=1}^{K}\sum_{j=0}^{k}\tau_{kj}\left(X^0\right)^j\left(-X^0\right)^{k-j} \mid X=0, C\right]\Pr\left(C \mid X=0\right) \\
=\ &0\cdot\Pr\left(C \mid X=0\right)=0,
\end{aligned}$$

where the second equality follows from the fact that compliers retire at age 60, and so have $X^0 = 0$.

Due to smoothness of $E\left[\varpi \mid X = x\right]$ by assumption, continuity of $E\left[\sum_{k=0}^{K}\lambda_k X^k \mid X = x\right]$ and continuity of $E\left[\sum_{k=1}^{K}\sum_{j=0}^{k}\tau_{kj}\left(X^0\right)^j\left(X-X^0\right)^{k-j}T \mid X=x\right]$ at $x = 0$ as shown above, we have

$$\lim_{x\to 0^+} E\left[Y \mid X=x\right]-\lim_{x\to 0^-} E\left[Y \mid X=x\right]=\tau_{00}\left(\lim_{x\to 0^+} E\left[T \mid X=x\right]-\lim_{x\to 0^-} E\left[T \mid X=x\right]\right).$$

It follows immediately that

$$\tau_{00} = \frac{\lim_{x \to 0^+} E\left[Y \mid X = x\right] - \lim_{x \to 0^-} E\left[Y \mid X = x\right]}{\lim_{x \to 0^+} E\left[T \mid X = x\right] - \lim_{x \to 0^-} E\left[T \mid X = x\right]}.$$

References

[1] Aguiar, M., and E. Hurst (2005), "Consumption versus Expenditure," Journal of Political Economy 113(5), 919-948.

[2] Aguiar, M., and E. Hurst (2007), "Life Cycle Prices and Production," American Economic Review 97(5), 1533-1559.

[3] Aguila, E., O. Attanasio, and C. Meghir (2011), "Changes in Consumption at Retirement: Evidence from Panel Data," The Review of Economics and Statistics 93(3), 1094-1099.

[4] Ameriks, J., A. Caplin, and J. Leahy (2007), "Retirement Consumption: Insights from a Survey," The Review of Economics and Statistics 89(2), 265-274.

[5] Angrist, J., G. Imbens, and D. Rubin (1996), "Identification of Causal Effects Using Instrumental Variables," Journal of the American Statistical Association 91(434), 444455.

[6] Banks, J., R. Blundell, and S. Tanner (1998), "Is There a Retirement-Savings Puzzle?" American Economic Review 88(4), 769-788.

[7] Battistin, Erich, Agar Brugiavini, Enrico Rettore, and Guglielmo Weber (2009), "The Retirement Consumption Puzzle: Evidence from a Regression Discontinuity Approach," American Economic Review 99(5), 2209-2226.

[8] Bernheim, D., J. Skinner, and S.Weinberg (2001), "What Accounts for the Variation in Retirement Wealth Among U.S. Households?" American Economic Review 91(4), 832-857.

[9] Blau, D. (2008), "Retirement and Consumption in a Life Cycle Model," Journal of Labor Economics 26(1), 35-71.

[10] Borella, M., F. C. Moscarola, and M. Rossi (2011), "(Un)expected Retirement and the Consumption Puzzle," unpublished manuscript.

[11] Calonico, S., M. D. Cattaneo, and R. Titiunik (2014), "Robust Nonparametric Confidence Intervals for Regression-discontinuity Designs," Econometrica 82(6), 2295-2326.

[12] Deaton, A. (1988), "Quality, Quantity, and Spatial Variation of Price," The American Economic Review 78(3), 418-430.

[13] Dong, Y. (2015), "Regression Discontinuity Applications with Rounding Errors in the Running Variable," Journal of Applied Econometrics, 30(3), 422-446.

[14] Fernandez-Villaverde, J., and D. Krueger (2007), "Consumption over the Life Cycle: Facts from Consumer Expenditure Survey Data," The Review of Economics and Statistics 89(3), 552-565.

[15] French, E. (2005), "The Effect of Health, Wealth, and Wages on Labor Supply and Retirement Behavior," Review of Economic Studies 72(2), 395-427.

[16] Friedman, M. (1956), "A Theory of the Consumption Function," Princeton, NJ: Princeton University Press.

[17] Gustman, A. L., and Steinmeier, T. L. (2001), "How Effective is Redistribution Under the Social Security Benefit Formula," Journal of Public Economics, 82, 1-28.

[18] Haider, Steven J., and M. Stephens (2007), "Is There a Retirement-Consumption Puzzle? Evidence Using Subjective Retirement Expectations," The Review of Economics and Statistics 89(2), 247-264.

[19] Hahn, J., P. E. Todd, and W. van der Klaauw (2001), "Identification and Estimation of Treatment Effects with a Regression-Discontinuity Design," Econometrica 69(1), 201-209.

[20] Hall, E. Robert (1978), "Stochastic Implications of the Life Cycle–Permanent Income Hypothesis: Theory and Evidence," Journal of Political Economy 86(6). pp. 971-987.

[21] Hurd, M. D., and S. Rohwedder (2013), "Consumption Smoothing during the Financial Crisis: The Effect of Unemployment on Household Spending," RAND Working Paper.

[22] Hurst (2008), "The Retirement of a Consumption Puzzle," NBER Working Paper No. 13789.

[23] Lee, D. S., and T. Lemieux (2010), "Regression Discontinuity Designs in Economics," Journal of Economic Literature 48(2), 281-355.

[24] Lusardi, A. (1999), "Information, Expectations, and Savings for Retirement," in Henry Aaron (ed.), Behavioral Dimensions of Retirement Economics, Washington, D.C.: Brookings Institution and Russell Sage Foundation: 81-115.

[25] McCrary, J. (2008), "Manipulation of the running variable in the regression discontinuity design: A density test," Journal of Econometrics 142 (2), 698-714.

[26] Miniaci, R., C. Monfardini, and G. Weber (2003), "Is There a Retirement Consumption Puzzle in Italy?" Institute for Fiscal Studies, Working Paper W03/14.

[27] Robb, A. L., and J. B. Burbridge (1989), "Consumption, Income, and Retirement," Canadian Journal of Economics 22(3), 522-542.

[28] Schwerdt, G. (2005), "Why Does Consumption Fall at Retirement? Evidence from Germany," Economics Letters 89(3), 300-305.

[29] Smith, S. (2006), "The Retirement-Consumption Puzzle and Involuntary Early Retirement: Evidence from the British Household Panel Survey," The Economic Journal 116, 130-148.

[30] Li, H., X. Shi, and B. Wu (2015), "The Retirement Consumption Puzzle Revisited: Evidence from the Mandatory Retirement Policy in China," Journal of Comparative Economics, forthcoming.

[31] Wakabayashi, M. (2008), "The Retirement Consumption Puzzle in Japan," Journal of Population Economics 21, 983-1005.