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ABSTRACT

Old-Age Pension and Intergenerational Living Arrangements[1]

China launched a pension program for rural residents in 2009, now covering more than 300 million Chinese. This program offers a unique setting for studying the ageing population, given the rapidity of China's population ageing, traditions of filial piety and co-residence, decreasing number of children, and dearth of formal social security, at a relatively low income level. This paper examines whether receipt of the old-age pension payment equips elderly parents and their adult children to live apart and whether parents substitute children's time involved in instrumental support to them with service consumption. Employing a regression discontinuity (hereafter RD) design to a primary longitudinal survey conducted in Guizhou province of China, this paper overcomes challenges in the literature that households eligible for pension payment might be systematically different from ineligible households and that it is difficult to separate the effect of pension from that of age or cohort heterogeneity. Around the pension eligibility age cut-off, results reveal large and significant reduction in intergenerational co-residence of the extended family and increase in service consumption among elderly parents.

JEL Classification: H55, I38, J14, J22

Keywords: rural pension, RD Design, living arrangement, service consumption

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1. Introduction

The impact of social pension income has far-reaching well-being implications for developing countries[2] in which families have been the primary informal institution for life-cycle savings and supports. Besides raising own income (Chen and Wang 2014) and reducing elderly labor supply (Ranchhod 2006; Filho 2008), social pensions have the potential to bring considerable changes to the extended families. For example, pensions reach prime-aged adults through promoting prime-aged labor migration (Bertrand et al. 2003; Posel et al. 2006; Ardington et al. 2009) and crowding out private transfers (Jensen 2004; Maitra and Ray 2003). Moreover, pensions affect children by discouraging child labor, promoting school enrollment (Edmonds 2006), improving nutrition and health status (Duflo 2000 and 2003; Case 2004), and pensioners' care of grandchildren (Case and Deaton 1998).

Establishing sustainable social pension systems in less developed countries is timely as the growth rate of senior population in these countries more than doubles that of developed countries (Edmonds et al. 2005). Similar to many other developing countries, China's population ages rapidly. By 2050, over 25.6 percent of China's population is expected to be over age 65.[3] Moreover, the implementation of stringent fertility policy in the last three decades has dramatically driven up the dependency ratio at a relatively low income level, which further intensifies the motivation among China's vast albeit shrinking working-age population to stay at home and take care of their elderly parents.

In this paper, we ask whether more income to elderly parents facilitate adult children to live apart and substitute service consumption for their time involved in instrumental support to parents. We answer this question through examining the pension payment out of a rural pension program. There has been mixed evidence on changes in family composition as a result of pension receipt in developing contexts. Edmonds et al. (2005) and Hamoudi and Thomas (2005) provide evidence on the selection of adults who co-reside with older adults when they become eligible for the pension. However, both Edmonds et al. (2005) and Jensen (2004) find no evidence that pension income leads to a significant increase in propensity to live independently or that the average size of the household respond to pension income.

The new rural pension scheme (NRPS) in China provides an opportunity to examine the pension with modest payment, compared to the well-documented generous pension program in South Africa that pays twice the average income (Lund 2007). The pension program in China may

[2] There are clusters of developing countries with large-scale social pensions: countries in the Southern Cone of Latin America; countries in southern Africa, including Botswana, Lesotho, Namibia, South Africa, and Swaziland; countries in south Asia, including Bangladesh, India, and Nepal. Outside these clusters, social pensions are scarce.
[3] Source: Population Division of the Department of Economic and Social Affairs of the United Nations Secretariat, *World Population Prospects: The 2010 Revision*, http://esa.un.org/unpd/wpp/index.htm.

demonstrate heterogeneous impacts considering the large disparities among age cohorts, between rural and urban areas and across regions in China. The elderly rural residents living in less developed regions are particularly disadvantaged. In 2010, the poverty rate for rural people aged 60 and above was as high as 22.3 percent, compared to 7.8 percent for rural population in China as a whole. Moreover, the sample under investigation comes from Guizhou, one of the less developed provinces in China, where the pension payment is about 25 percent of the average income.

This paper has two main advantages over the large literature that identifies the impact of pension. First, evidence suggests households eligible for pension payment are systematically different from ineligible households. Therefore, the estimations on pension impact using survey data likely suffer from omitted variable bias. Ardington et al. (2009) control for time-invariant individual fixed effects in panel data estimation, but fixed effect methods require very specific assumptions about the nature of unobserved factors and their persistence over time. Meanwhile, they cannot follow individuals who dropped out of the demographic surveillance between waves of the survey, which generates biases in longitudinal estimations. We draw on a rich survey to track each adult child's living arrangement and demographic information, regardless whether she is counted as a household member or living under the same roof with parents. Thus, we can generate datasets of pensioners' and non-pensioners' adult children without introducing complication of endogenous household formation.

Second, the co-residence decision of household members is largely affected by life-cycle pattern of the elderly parents and cohort heterogeneity. When pension is granted according to age, the prior research examining the linkage between pension and co-residence cannot separate the effect of pension from that of age or cohort heterogeneity. We apply a RD design in this paper to overcome this issue. Residents are eligible for the old-age pension starting from age 60. Most elderly remain active in agricultural production well into their 60s in rural China, so the rural pension program is not tightly linked to labor force participation. Such RD design has several advantages over other methods of identification and has been applied to multiple different settings (Edmonds et al. 2005; Card et al. 2009; Miller et al. 2013; Sun and Eggleston 2014). In our context, assuming that household structure is continuous with respect to the age of the elderly, our research design disentangles the effect of pensions from other age-related factors shaping living arrangements.

Our study is distinguished from the narrower literature that identifies the impact of pension on intergenerational living arrangements. In addition to the intent to treat (ITT) type of analyses that make use of discontinuity in age eligibility, such as the reduced form RD estimations in Edmonds et al. (2005), we take advantage of a discontinuity in actual pension receipt induced by age eligibility. Our fuzzy RD approach adjusts the attenuated measure of the true impact when take-up of the pension is incomplete or does not coincide exactly with age eligibility (Edmonds et al. 2005). Compared to studies using age eligibility as the instrumental variable, such as Hamoudi and Thomas (2005) and Posel et al. (2006), our RD design is more flexible to choose functional forms, optimal bandwidths, and a set of narrow age cohorts around the age eligibility cut-off, which reinforce that the results are robust. Compared to the diff-in-diff-in-diff (DDD) approach in Jensen (2004), our RD design does not assume that living arrangements between the groups receiving pension or not have to follow the same trend in the absence of the pension.

We focus on families with *adult* children as adult children lend major support to their parents and therefore are more sensitive to their parents' receiving pension payment. We focus on the subsample of adult son as their co-residence decisions are affected more by parents receiving pension given their higher rate of co-residence with parents. Our results utilizing a longitudinal household survey reveal several key findings. First, RD reveals large and significant reduction in adult sons' co-residence with parents around the cut-off age for pension. Second, service consumption among parents booms. Specifically, more service consumption is broadly reflected in three dimensions: pensioners are slightly more likely to pay for haircut and machine repair; pensioners start to purchase coal for heating and cooking that substitute for much time spent on collecting firewood in the mountainous region; pensioners seek formal medical treatment when recommended by a doctor, even though health insurance coverage shows no significant changes after the pension eligibility threshold.

Since half of these households are below the USD 1.25 international poverty line (Appendix Table 1), they are very likely to be credit constrained. The increase in the income of elderly parents may relax this constraint, making services more affordable to substitute for instrumental support directly provided by their adult children.

The RD design survives a series of key placebo tests and robustness checks, including verifying no jumps of outcomes at age cut-offs other than 60, no discrete changes in baseline characteristics and predetermined outcomes at the age 60 cut-off, no discontinuities of the density of the running variable, no confounded policy discontinuities at age 60, and that results are robust to various bandwidths, bandwidth selection procedures, and polynomial specifications.

Though our Guizhou survey is not nationally representative, we believe this study has methodological advantages and the results make it of broader interest. In particular, the unusual details of the survey enables us to link parents with adult children left home and shed light on intergenerational arrangements where people are so poor that the new pension program is most likely to generate salient impact through promoting income of elderly parents.

The rest of the paper is organized as follows. Section 2 introduces the new rural pension program in China, its implementation in Guizhou province, and its potential impact on intergenerational living arrangements. Section 3 describes the data and the identification strategy. Section 4 presents results and discusses interpretations. Section 5 conducts a series of key robust checks. Section 6 offers concluding remarks.

2. Institutional Background and Mechanisms

2.1 The New Rural Pension Reform in China and Guizhou Province

In September 2009, China launched a pension program for rural residents, the New Rural Pension Scheme (NRPS), now covering more than 300 million Chinese (Appendix Figure 1). The NRPS aims at universal coverage of formal and informal workers in rural China. By 2012, the central government and local government have invested more than 232 billion RMB and 30 billion RMB to the pension fund, respectively, and all 2853 counties in China have implemented the NRPS.[4] The number of contributors, beneficiaries, pension fund expenses and the balance have been rising sharply. Similarly, the rural pension system in Guizhou was vacant before the NRPS but has been rapidly established since 2010 (Appendix Figure 1).

The pension fund consists of three parts: an individual premium, a local government subsidy, and a central government subsidy. According to the guidance released by the State Council of China, there are five basic categories of individual premium: 100, 200, 300, 400, and 500 RMB per year per person. Local government may introduce more categories to adjust for increasing cost of living. Individuals are free to choose from this menu of options, while 100 RMB is the most popular choice (Lei et al. 2011). The local government subsidy must be no less than 30 RMB per person per year. The higher individual premium one chooses to pay, the more subsidies local government pays. The individual premium and the local government subsidy form the individual account. Besides, a basic pension benefit, 55 RMB per month per person in Guizhou but as much as 310 RMB per month per person in a few rich provinces / municipalities, is available to all residents financed by the central government. Pensioners receive payment from both the individual account and the basic pension benefit.

All rural residents aged 16 or above who are not enrolled in urban basic pension program can enroll in NRPS voluntarily. Participants below age 45 are required to contribute at least 15 years to become eligible for pension payment upon reaching age 60. However, according to the State Council, there is no required minimum years of contribution for those aged between 45 and 60, who are only encouraged to contribute more to meet the shortfall in contributions to their own account over their remaining years before age 60. Both our Guizhou survey and two national representative samples in China, i.e. China Center for Agricultural Policy survey and China

[4] See the news report (in Chinese) http://www.eeo.com.cn/2013/0204/239758.shtml .

Family Panel Studies (CFPS), suggest that most people aged 55-60 wait until age 60 to receive free basic pension offered by the central government without any contribution to the individual account or follow younger cohorts to only pay for the lowest level premium (100RMB) (Chen and Wang 2014).

Pension payment is non-contributory to anyone aged 60 or above by the local rollout time of NRPS. In other words, they are not required to make any contributions to the NRPS system, have cumulative work histories, or have extended families to become eligible for the basic pension benefit.5 However, those aged 60 or above do not have individual accounts as the NRPS was only rolled out recently.

The NRPS may have heterogeneous impacts due to China's large regional disparities. The NRPS can be more attractive to the poor with little formal old-age support. While the benefit accounts for a smaller proportion of income in developed regions, it can be much larger in less developed regions. For example, in our surveyed area in Guizhou, pension payment represents one fourth of average income (Appendix Table 1). Considering that the elderly earn less than the young generation, the pension payment should exert a larger impact.[6]

2.2 Income and Living Arrangements: Potential Pathways

Pension payment directly raises pensioners' income. However, as people age, their work activities and income gradually decline. They may also reduce their labor supply in response to more pension income. An overall positive income shock is necessary to generate the observed pattern of living arrangements and service consumption. Consistent with the fact that receiving pension payment from the NRPS is not tied to individual employment decisions, tests using our Guizhou survey, available upon request, reassure that personal income of the elderly (excluding pension benefit) is continuous around the pension eligibility cut-off but discontinuously increases when pension income is accounted for.[7]

Half of the surveyed households are still below the USD 1.25 international poverty line (Appendix Table 1), pension payment may provide parents in these extended families more economic resources. Consequently, they are more economic independent and their adult children feel more relieved to live apart. On the demand side, more income enables parents to

[5] Some regions in China stipulate that those aged 60 or above when the NRPS rolled out in the local county are eligible for the basic pension benefit only when their children eligible for the NRPS also enroll in. However, this is not the case in our surveyed region in Guizhou.

[6] Though the pension benefit only accounts for a small proportion of household income, it means a 7-8 times higher income for the elderly population in rural China who sometimes had no income before the NRPS (Zhang et al. 2013).

[7] Using two national samples, China Health and Retirement Longitudinal Study (CHARLS) and China Family Panel Studies (CFPS), Zhang et al. (2013) and Chen and Wang (2014) find the same income pattern around the cut-off age.

consume more services. On the supply side, reduced co-residence leaves a vacuum of instrumental support originally provided by adult children. The gap between demand and supply of services are filled by those from the market (Hoerger et al. 1996).

If we consider living arrangements as a form of consumption (Costa 1997 and 1999), more economic resources to extended families may also explain a falling household public goods consumption (such as companionship) or even the decision to live apart. Half of the households in our sample are still in extreme poverty. With scarce economic resources, the economy of scale may determine that larger extended families are formed to share household public goods, and privacy is less attractive (Salcedo et al., 2012). Table 1 suggests that more than 80 percent of parents co-reside with adult children. However, the consumption of household public goods at lower cost ties to the optimal number of family members. The likely release of credit constraint motivates family members to live separately and therefore family size to fall (Zimmer and Kwong 2003), rendering public goods sharing less economical. Meanwhile, if the utility function displays more curvature in public than private goods, the increasing purchasing power as a result of pension payment may lead to a higher expenditure share on private goods. In both cases, the benefit from living together reduces, which promotes our observed co-residence pattern.

Due to the traditional patriarchal residency in rural China, adult sons are more likely to live with parents and play a more important role in taking care of their elderly parents than daughters do. Therefore, we hypothesize that the increase in income of pensioners enables them to substitute service consumption for care directly provided by adult children, especially adult sons, thereby freeing up them to live separately.

Since the current pension payment is still modest and a large proportion of elderly population in impoverished rural China are in poverty, their potential anticipation of becoming eligible for pension in the future and the ability to borrow from future, by saving less or borrowing more from others, can be restricted and may not offset much of our identified pension impact. In the following empirical analysis, these hypotheses will be indirectly tested by checking whether there is no discontinuity around pseudo cut-offs below age 60.

3. Empirical Strategy

3.1 Data

This paper makes use of a rich panel survey of households in Guizhou province. With the help of the International Food Policy Research Institute (IFPRI) and China Academy of Agricultural Sciences (CAAS), we attempted to keep track of all residents in 26 randomly selected villages in a county in Guizhou, one of the poorest provinces in China. With a total population of 402,000 people, per capita income in the county is above the provincial median but below the provincial mean, suggesting that its income profile is representative of Guizhou province as a whole. Albeit double-digit income growth, half of the residents in the 26 villages are still in poverty and experience worsening income inequality (Appendix Table 1).

Each wave of survey was conducted a few days before the Lunar New Year (January or February) when most migrants return to celebrate. We also managed to visit each household several times to make sure all family members at home were interviewed individually. In the latest 2011 wave 5189 individuals from 963 households were interviewed. Three waves of survey, i.e. 2004, 2006 and 2011, are utilized in the analysis.[8] Since the NRPS was introduced in Guizhou in 2010, the 2004 wave and the 2006 wave are used in placebo tests.[9]

Four subsamples are critical to our analysis, including the sample of adult son (N=1016), the sample of parents with adult son (N=1226), the sample of adult children (N=1741), and the sample of parents with adult children (N=1669). The survey collected rich information on demographics, income and consumption for all household members (including returning migrants and other non-coresident members), their relationship and living arrangements with the household head. At the beginning of each interview, every household was asked by the survey enumerators to show its official household registration record (a.k.a. the *Hukou* book) in which accurate official record for all members in the household, whether migrants or not, is displayed. Besides filling in the family roster based on information from the household registration book, the survey asked each household to recall previous members who moved out and therefore were removed from the household registration book.[10]

Living arrangements were identified from the family roster questionnaire, which records all co-resident and non-coresident members. Co-residence is defined as those living under the same

[8] Among all 4536 individuals surveyed in the 2004 wave, 4339 were resurveyed in the 2006 wave (attrition rate=4.4%), 4221 were resurveyed in the 2009 wave (attrition rate=7.0%), and 3954 were resurveyed in the 2011 wave (attrition rate=12.8%).

[9] The survey has four waves, i.e. 2004, 2006, 2009 and 2011. Each wave was implemented in January and/or February of the following year, i.e. 2005, 2007, 2010 and 2012. Since Guizhou province started the pilot NRPS in early 2010, the 2011 wave is in the post-NRPS period and the first two waves of 2004 and 2006 are in the pre-NRPS period. We do not use the 2009 wave data in placebo tests mainly due to two concerns. First, people may already know future implementation of the NRPS then. The resulting anticipation effect may confound the identification of the pension effect. Second, the 2009 wave overlaps with the recent world financial crisis, which possibly affects our main outcomes of interest, i.e. co-residence decisions, through unknown discontinuous changes around the age cut-off that contaminate the RD design.

[10] Due to the patriarchal residency tradition in rural China, most daughters move out of their parents' home after getting married. Parents, especially those in older ages, may underreport daughters who married and moved out a long time ago. This may explain the more skewed male-female sex ratios among (adult) children in our sample (1016/725) than that can be inferred solely from son preference. However, this potential bias should not affect our main results as adult sons are the main elderly care givers and are likely affected more by the pension given to elderly parents, while very few married daughters co-reside or stay in the same village with their parents across all age groups. Daughters left home many years ago should not be discontinuous around the age cut-off in 2011.

roof and dine together most of the time. Non-coresident family members include others living in the same village, same town, same county, same province or other provinces (Figure 1). Measures of intergenerational living arrangements were constructed from both adult children and parents perspectives. For adult children, we measure whether they co-reside with parents. For parents, we measure whether they live with any child or any son and the number of adult sons they live with.

3.2 The RD design

The empirical challenge in obtaining a consistent estimate of the causal effect of pension is that pension receipt is endogenous. We utilize an age discontinuity in the benefit structure of the social pension program. This discontinuity produces an abrupt shift in eligibility among otherwise similar individuals and households. Assuming other household and individual characteristics are smoothly distributed according to age, i.e., the individuals very close to pension eligibility (in their late 50s) and those who just became eligible (in their early 60s) are comparable, the abrupt shift in eligibility utilized by this research design allows us to isolate the causal link between pension receipt and any shifts in household or individual outcomes of interest. Specifically, two assumptions must hold. First, the outcome *y* for an age *a* depends on (observed and unobserved) characteristics of the household and all of its potential members, and these characteristics must have smooth effects on the outcome variables. Second, families cannot respond to the pension policies by altering the age of the elderly.

Social pension income depends on whether one passes age 60. Figure 2A plots the sample means of pension receipt by normalized age, where 0 represents the eligibility threshold of exact age 60. As expected, there is a sharp upward break in pension receipt at the age cut-off.

Therefore, we focus on discussing the local treatment effect of pension income. Since a few people below age 60 report receiving pensions, we adopt a Fuzzy RD design, which allows the jump in the probability of assignment to the treatment at the threshold to be smaller than 1. In this design, we interpret the ratio of the jump in the regression of the outcome on age to the jump in the regression of the treatment indicator on age as an average causal effect of the treatment (Imbens and Lemieux, 2008), which is analogous to an “intent to treat” parameter of a randomized controlled trial where the treatment is receiving pension income ($pension_i$), but because of “lack of compliance” a few people assigned to treatment by passing the eligible age ($1[age_i \geq 60]$) do not actually end up receiving pension. Formally, the estimand is

$$(1)\quad \tau_{FRD} = \frac{\lim_{a\downarrow 60} E[y \mid age_i = a] - \lim_{a\uparrow 60} E[y \mid age_i = a]}{\lim_{a\downarrow 60} E[pension_i \mid age_i = a] - \lim_{a\uparrow 60} E[pension_i \mid age_i = a]}$$

To interpret this result, we further require that assignment to treatment satisfy a monotonicity property. Specifically, we are interested in how pension income impacts subsequent household outcomes for compliers. Compliers are individuals who do not receive a pension if not eligible (below age 60) and receive a pension if eligible (above age 60). Formally,

(2) $\tau_{FRD} == E\left[y_i(1) - y_i(0) \mid individualiis\ a\ complier\ and\ age_i = a \right]$

Following Lee and Lemieux (2010) and estimating variants of the following equations:

(3) $Y_i = \alpha_1 + f_1(age_i - 60) + 1[age_i \geq 60] \cdot \pi_1 + X_i \delta_1 + \varepsilon_{1i}$

(4) $pension_i = \alpha_0 + f_2(age_i - 60) + 1[age_i \geq 60] \cdot \pi_0 + X_i \delta_0 + \varepsilon_{0i}$

where Y_i is the vector of outcome variables, $pension_i$ indicates whether a person actually receives pension, $f_1(\cdot)$ and $f_2(\cdot)$ are polynomials in the distance from the age 60 cutoff, X_i is a vector of demographic and socioeconomic covariates, and ε_i is the error term. As long as $f_1(\cdot)$ and $f_2(\cdot)$ are continuous at $age_i = 60$, π_1 and π_0 represent the magnitude of the discontinuities in the numerator and denominator, respectively, of (1). $\hat{\tau}_{FRD} = \hat{\pi}_1 / \hat{\pi}_0$ estimates the Local Average Treatment Effect (LATE) of pension receipt. The inclusion of covariates X_i, however, should not affect the estimation of π_1 and π_0 if X_i do not vary discontinuously at the age cutoff. In Appendix Figure 2, we draw the relationships between main baseline covariates and normalized age to check discontinuity.

To overcome the challenge of finding an appropriate estimator for $f_1(\cdot)$ and $f_2(\cdot)$, we specify a flexible parametric model for $f_1(\cdot)$ and $f_2(\cdot)$. The main results in this paper adopt the linear specification since a model selection algorithm using the Schwarz (1978) criterion suggests that in more than half of all cases the preferred specification is linear. In Appendix Table 2, we present the main results using the 2nd order and the 3rd order polynomial specifications. Specifically, when, for example, a 3rd order polynomial form is specified, $f_1(\cdot)$ and $f_2(\cdot)$ take the following forms

(5) $f_1(age_i - 60) = 1[\text{age}_i \geq 60]\square\sum_{p=1}^{3} \gamma_{1p}(age_i - 60)^p + (1 - 1[\text{age}_i \geq 60])\square\sum_{p=1}^{3} \gamma'_{1p}(age_i - 60)^p$

(6) $f_2(age_i - 60) = 1[\text{age}_i \geq 60]\square\sum_{p=1}^{3} \gamma_{0p}(age_i - 60)^p + (1 - 1[\text{age}_i \geq 60])\square\sum_{p=1}^{3} \gamma'_{0p}(age_i - 60)^p$

in which each term of the polynomial is interacted with the age eligibility dummy, allowing each term to have different parameters on either side of the age cutoff. γ_{1p}, γ_{0p} and $\gamma_{1p}^{'}$, $\gamma_{0p}^{'}$ are coefficients on the p polynomial terms for people eligible and ineligible for pension income, respectively, allowing the shape of the underlying conditional expectation to be different to the left and right of the threshold.

We first present the results of reduced form regressions that estimate the numerator and denominator in (2) respectively. To do so, we first use OLS regression for each of the estimations and then RD regression with different bandwidths. Next, we report the RD results of the treatment effect τ_{FRD}. As proved in Hahn et al. (2001), the IV estimator using $1[age_i \geq 60]$, the polynomial terms of ($age_i - 60$), and their interactions as instruments for pension produces the same τ_{FRD} as in equation (2). Therefore, in the empirical analysis on local treatment effect, we only present one estimator of τ_{FRD}.

All the RD estimates are bias-corrected with robust standard errors. The main results of this paper use rectangular kernel11 with the Calonico, Cattaneo, and Titiunik (2014, CCT thereafter) bandwidth selector. In Appendix Table 2, we also report results using the Cross-Validation (CV thereafter) bandwidth selector, proposed by Imbens and Lemieux (2008), and the Imbens and Kalyanaraman (2012, IK thereafter) bandwidth selector. We also report main outcomes with half of the optimal bandwidths to increases the weights on observations closer to the cutoff.

4. Empirical Results

4.1 Summary statistics

Figure 1 plots the distribution of living arrangements for adult sons and daughters. Adult sons are ten times more likely to co-reside with parents than adult daughters. However, among adult children who do not co-reside with their parents, daughters live much closer to them. Overall, this is consistent with the patriarchal residency tradition in rural China that sons are more likely to live with parents but daughters may play a supplementary role in taking care of parents.

Table 1 presents summary statistics for the key outcome measures. The first column in each panel reports the means and standard deviations for the whole sample. Columns (2) and (3) report means and standard deviations when confining the sample to households with a member between ages 55 and 65, i.e. 5 years around the age cut-off. Column (4) reports the result of t-test comparing the differences between those below and above the pension eligibility threshold. Note that these descriptive statistics could be very different from the causal effect of pension receipt or the estimands in expression (1) and (2) as the means and t-

[11] In general, with the rectangular kernel the RD estimation places more weight on observations closer to the cutoff relative to the triangular kernel.

tests may capture the impact of pension receipt as well as the trends regarding ageing and cohort heterogeneity. Fortunately, our RD design is able to distinguish the discrete changes associated with pension eligibility and the resulting effect of pension income from the life cycle pattern of extended household formation and other outcomes.

Panel A of Table 1 restricts the sample to adult sons (N=1016) as they generally co-reside with parents, we find up to 48.5% of them co-reside with parents. When limited to a five-year window around the pension threshold, the co-residence rate decreases from 42.4% with a nearly-eligible parent to 16.1% with an eligible parent. In Panel C, the intergenerational co-residence rate reduces when the sample is confined to adult children (N=1741) with a parent between age 55 and 65. Similarly, results using both the sample of parents with adult son (Panel B, N=1226) and the sample of parents with all adult children (Panel D, N=1669) demonstrate significant reduction in the number of adult sons to live with and the rate of co-residence with any child or any son.

These results are quite salient considering that cohort heterogeneity and life cycle factors may work in the opposite direction that prevents us from observing this pattern. For example, as parents becoming older, they may demand more instrumental support, and their adult children tend to co-reside to take care of them.

Panel E of Table 1 summarizes service consumption for parents with adult son. For those between age 55 and 65, the eligible are 42.7% more likely to pay for haircut and machine repair, slightly more likely to buy coal for heating and cooking, and less able to be assisted by adult children on a daily basis.

However, we should note that the t-tests on the outcomes by pension eligibility cannot identify the effect of pension receipt from various contaminating channels. Next we use a RD design to isolate the causal impact of pension income.

4.2 Empirical Estimations

A. The first stage: Pension receipt

Figure 2A plots the average rate of pension receipt against the cut-off age 60 for the whole sample of 5189 respondents. Using a semi-parametric RD estimator, the fitted probability curve of receiving pension in Figure 2B suggests an almost 80 percentage jump in pension receipt immediately above the cut-off age. We regress the dummy indicator of receiving pension (1 if receiving pension, 0 otherwise), allowing a change in pension eligibility beginning from age 60. Panel A of Table 2 shows that at the eligibility threshold, around 86 percentage more people receive pension payment.

To make sure the discontinuity of pension receipt holds for all subsamples we analyze, Panel B through Panel F run OLS regressions and first stage RD estimates for each subsample. Results suggest discontinuous jump across all subsamples with increasing rates of pension receipt from 50 to 90 percentages.

B. *Living arrangements*

Living arrangements between adult children and their parents in a post-NRPS year are plotted in Figure 3 column 1, while living arrangements for the same group of adult children are performed in a pre-NRPS year in columns 2-3 as placebo tests. Specifically, we compare living arrangements around age 60 in 2011 to living arrangements in 2004 and 2006 around age 60 in 2011. If there is no discrete change at the age 60 cut-off in 2011 using the 2004 and 2006 waves (the pre-NRPS period), the discontinuity in 2011 using the 2011 wave (the post-NRPS period) suggests that our results should not be driven by any pre-existing discrete change at age 60.

China has experienced labor shortage since 2004. As a result, real wage in urban China has increased dramatically, attracting a large flow of migrants (over 200 million) to leave their home villages. While there is no discrete change at the cut-off in panel B and D, the rate of co-residence for sons greatly decreases from 2004-2006 to 2011, which is consistent with the declining trend of co-residence in rural China (Peng, 2011). The rise in co-residency with parents above age 60 mainly reflects the tradition of instrumental care directly provided by adult sons when their parents get older.[12] Daughters' living arrangement, however, is mainly dependent on marital status under the patriarchal residency tradition, so we do not see significant jump in daughters living arrangement at the cut-off.

Table 3 reports the estimation results for OLS, reduced form RD and RD LATE. Our RD design enables us to separate pension's effect from the common time trend. We confirm that separation happens only on adult sons. In Panel A post-program period, it shows that the adult sons are 26.6 percent less likely to co-reside with parents once a parent receives pension. The living arrangements of adult daughters are less likely to be driven by declining opportunity cost upon policy changes after marriage.

We reexamine the same idea using the sample of adult parents. As adult daughters' living arrangement under the patriarchal residency tradition is much more responsive to their marital status than to the arrival of pension income, we confine the sample to those with at least one son. Within this smaller sample, at the eligible age the number of sons with whom the adult parents live with decreases by .500. This is consistent with Panel A that adult sons are less likely

[12] The average life expectancy for residents in rural Guizhou, our surveyed province, is below 70. Though adult sons are strongly motivated to come back and take care of their elderly parents, Figure 3A suggests that it takes almost 10 years for the average rate of co-residency returns to that for parents before age 60.

to co-reside with parents. In the right part of Table 3, none of the co-residence indicators using the 2004 and 2006 waves around age 60 in 2011 show a significant jump at the placebo age threshold, supporting that the effects for 2011 identify the true pension impact. We plot the number of co-residing sons for parents with at least one son in Figure 4.

In an environment with a large number of biological sons work part-time in the home county or migrate out to work, with the tradition of male children taking financial responsibility for elderly care, and with the patriarchal residency tradition in rural society, daughters-in-law may play an important role in taking care of their parents-in-law and even substitute for the care provided by biological daughters. Our estimations, available upon request, show a smaller but statistically insignificant decline in co-residence between parents and daughters-in-law as a result of pension receipt, which are robust to various bandwidths, bandwidth selection procedures, and polynomial orders around the age cut-off. Some couples move out together, while some daughters-in-law stay to provide instrumental support to elderly parents when their husbands leave home. This may explain our finding that biological sons respond more to parental pension receipt than daughters-in-law do.[13]

C. Consumption

As only 4.8 percent of adult daughters co-reside with parents and they often play a secondary role in caring for parents compared to adult sons, pension income is expected to mainly affect families with adult son. Therefore, our analysis on service consumption is restricted to parents with adult sons.

Appendix Table 3 reports the effect of pension income on whether paying for haircut or machine repair in the past year. We group the two items as a dummy indicator. The fuzzy RD estimand shows a jump of 33.2% in receiving any of the two services among parents, and it is on the margin of significance at the 10 percent level. The survey also collects information on coal purchase. More than half of the families in the survey collect firewood to heat and cook, which has been time costly in the mountainous region where our survey was implemented. The pension payment triggers a 40.2 percent discrete increase in the purchase of coal, which saves labor. All these changes may largely due to stronger purchasing power as a result of pension. The reduced intergenerational co-residence and assistance may contribute as well.

In rural China, financial difficulties often prevent patients from receiving necessary treatment. Perceived access to medical services when recommended by a doctor is coded as a dummy

[13] Future surveys with rich information on the time devoted to taking care of parents among adult children (biological and in-law) may enable us to more carefully investigate this important hypothesis.

variable. The RD estimation finds that the propensity of not receiving formal medical treatment when recommended reduces 46.3% at the cut-off age.

Many elderly parents who are seriously ill receive assistance from adult children. However, this number inevitably reduces when the two generations no longer co-reside. The RD estimation finds that the chance of being assisted by adult children on a daily basis reduces 43.9%. All these patterns are consistent with pension income enabling the pensioner to consume more services to replace the care provided by children.

To establish the causal link, we need to rule out potential mechanisms behind the increase in the perceived access to needed medical services. For example, it is possible that on the demand side of health care pensioners' health status decreases significantly at age 60 or on the supply side doctors are more likely to prescribe a treatment once they know an elderly patient is above age 60 or coverage of the New Cooperative Medical Scheme (NCMS) is higher for individuals above 60. We test the continuity of health and the take-up of NCMS using the information collected in the survey. As shown in Appendix Table 3, there is no discontinuity of self-reported health status or take-up rate of NCMS for the elderly at age 60.

5. Robustness

Our results suggest that after pension payment adult children are less likely to co-reside with their parents. The parents tend to consume more services. To the extent that parents around the age cut-off are similar in other characteristics that determine these outcomes, these results imply that pension receipt equips pensioners' extended families with more purchasing power. This section presents several pieces of evidence that this key identification condition is met.

First, we test possible discontinuities at other age cut-offs, which sheds light on whether the jumps we observe are driven by other age-related changes rather than the arrival of pension income. An individual barely below pension eligible age should be very similar to those just turn 60 except that the former receives no pension payment. Appendix Table 4 reports the placebo test results on pension receipt and main outcome variables using age 59 as the cut-off. The first stage estimation of pension receipt using the sample of parents with adult children finds no significant change around the placebo age cut-off (Panel A). First stage placebo tests using other subsamples, available upon request, are quite similar. In the second stage, Panel B employs the sample of parents with adult son, while Panel C utilizes adult children data. None of the fuzzy RD estimands is significant for any outcome variable. The same tests are conducted for other pseudo age cut-offs nearby, such as 59.5, 60.5 and 61, no significant discontinuities are found. Therefore, we are more confident that the discrete change at age 60 is driven by pension income.

The result of no discontinuity at the pseudo cut-offs below age 60 also suggests that both the anticipation effect and people's ability to borrow from future are limited. If borrowing from future is a viable option, people may do so, for example, by saving less or borrowing more from others in anticipation of becoming eligible for pension income in the future. The anticipation effect, if exists, may offset the impact of pension, rendering our identified pension impact a lower bound of the true effect of pension on co-residence and service consumption.

Second, we test for discontinuities in baseline characteristics that should not be affected by the treatment. If baseline demographic or socioeconomic characteristics show discontinuous jumps, we would be concerned about the similarity between people barely below and barely above the age cut-off. While we can never be certain that the unobservable characteristics of the elderly satisfy this condition, the validity of this assumption can be tested by ensuring that the conditional expectations of the observable characteristics do not vary discontinuously in the neighborhood of the cut-off age. Appendix Figure 2 shows the graphic equivalent of the test on main baseline demographic and socioeconomic characteristics of parents (Appendix Figure 2b) and their adult children (Appendix Figure 2a). All the horizontal axes represent parents' normalized age relative to 60. The vertical axes are ethnicity, cadre and party membership, whether one finishes nine-year mandatory education, marital status, and religious beliefs, respectively. Appendix Figure 2a also tests adult child age (while Appendix Figure 2b is not able to include parents' age since it is the horizontal axis). None of these characteristics shows discontinuity around the age 60 cut-off. The statistical equivalent of these Figures is available upon request.

Third, another standard check in a RD design is to verify that no other predetermined outcomes display discontinuities around the cut-off apart from the treatment. As already reported in Table 3, outcome variables before the pension program, drawn from our 2004 and 2006 waves of survey, find no discrete change at the age cut-off, which lends us credibility that our results should not be driven by these predetermined jump at the cut-off.

Fourth, we test for continuity of the density. Assignment to pension program around the age 60 threshold should ideally be randomized. However, if richer families are more able to manipulate age, they are more likely to receive pensions and be able to afford living separately. This could lead to biased estimates of the effect of pension payment. In the present case this seems unlikely. The age profiles on the registration card were issued far ahead of the pension reform. Moreover, China's strict *hukou* system mitigates the potential cases of misreporting age. The normalized age as a continuous running variable further eliminates the heaping issue as a result of irregular rounding up (or down) of age profile.

Formally, we conduct two verification tests. First, we examine the income and other characteristics of the very few cases in our sample who report receiving pensions before age

60. No significant demographic and socioeconomic characteristics are found for this small group of pensioners. Second, we check whether there is a discontinuity of age profiles at the age cut-off – in this case whether "missing" ages just below the cut-off are followed by "hump" above. The histogram of age in Appendix Figure 3 shows little discrete change in the density around the age 60 cut-off. We adopt the McCrary (2008) test for discontinuity in the age density function. We also fit a linear term in the normalized age, a dummy equal to one if one is eligible for pension income, and their interaction to the log of the fraction of observations with each baseline age using weighted least squares regression. The test confirms that no statistically significant discontinuities are evident in the (log of the) age density.

Fifth, to my best knowledge, there exists no other policy discontinuity at age 60 that may confound the identification of pension receipt. For example, there is no other social security program stipulating that people above age 60 are eligible for it. Meanwhile, there is no subsidized nursing home regulating that age 60 is the eligible age cut-off to live in.

Sixth, besides the CCT bandwidth selector, our results are robust to other bandwidth selection procedures, such as the CV bandwidth selector and the IK bandwidth selector (Appendix Table 2). Results are also robust to bandwidth selections. Main outcomes with half of the optimal bandwidths and with twice the optimal bandwidths generate quite similar results as with the optimal bandwidths (Appendix Table 2). Meanwhile, results are robust to polynomial selections, such as in linear, quadratic and cubic specifications (Appendix Table 2).

Finally, additional checks make better use of the longitudinal feature of the survey data and reassure us that the observed changes in intergenerational living arrangements are not driven by changes in the composition of our sample across waves of survey. Results using the 2006 and 2011 waves are presented in Appendix Table 5. In our RD framework, the 2006 wave is only used for placebo tests simply because there was no policy induced age discontinuity. However, estimations using the diff-in-diff (DD) strategy take advantage of both before-after and treatment-control differences in the 2006 and 2011 waves. Results are similar between our current RD design and the DD strategy, though the former estimates a LATE while the latter estimates an ITT.

6. Conclusions

This paper provides early evidence on the impact of the newly implemented old-age pension program, the NRPS, in China. Utilizing our primary longitudinal household survey from rural western China, we show that adult children, who by virtue of having parents reach age 60 and receive pension, are less likely to co-reside with their parents. Their parents substitute service consumption for instrumental support provided by them.

We contribute to the growing body of research indicating the large impact of social pension on pensioners and its spillover effects on extended families, especially providing the first piece of evidence for China where the new pension payment is not as generous. More importantly, this paper is among the first that apply RD in the study of social pension, which overcomes two main challenges in the literature that pension eligible households can be quite different from ineligible households and that age or cohort related heterogeneity may confound the pension impact. However, we infer but do not directly test whether rural households are credit constrained.

Our empirical findings hold important implications and general lessons for household theory. First, social pensions directed at the elderly could benefit other generations. Second, pension programs that transfer income to elderly population should be more cautious as substantive changes in living arrangement may offset some benefits through reduced instrumental support to parents directly provided by children. Third, the changes in living arrangements as a result of pension payment may have implications for other aspects of family life that deserve more research.

Future research will highlight the implication of reduced intergenerational co-residence on career choices of the younger generation, such as off-farm activities and migration. A second avenue for future research is to examine whether gender identity of the pension recipient affects its use. Pension payment provides an income source that is attributed to a specific family member and can be used to assess how household behavior and well-being are affected differently due to different members gaining control of this income source. Future studies will distinguish potential different impacts of income shock to fathers versus mothers. When parents have the same utility function, pension affects co-residence decisions of extended families and pensioners' service consumption only through increase in household income. However, co-residence and consumption decisions may be affected by individual income shock when parents have different utility functions and intra-household bargaining power (or the weight for a household member in the household utility maximization) changes. Finally, apart from gender identity, the number of pensioners in a household provides an under-utilized and plausibly exogenous source of income variation for our future study.

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Figure 1. Distribution of living arrangement for adult children

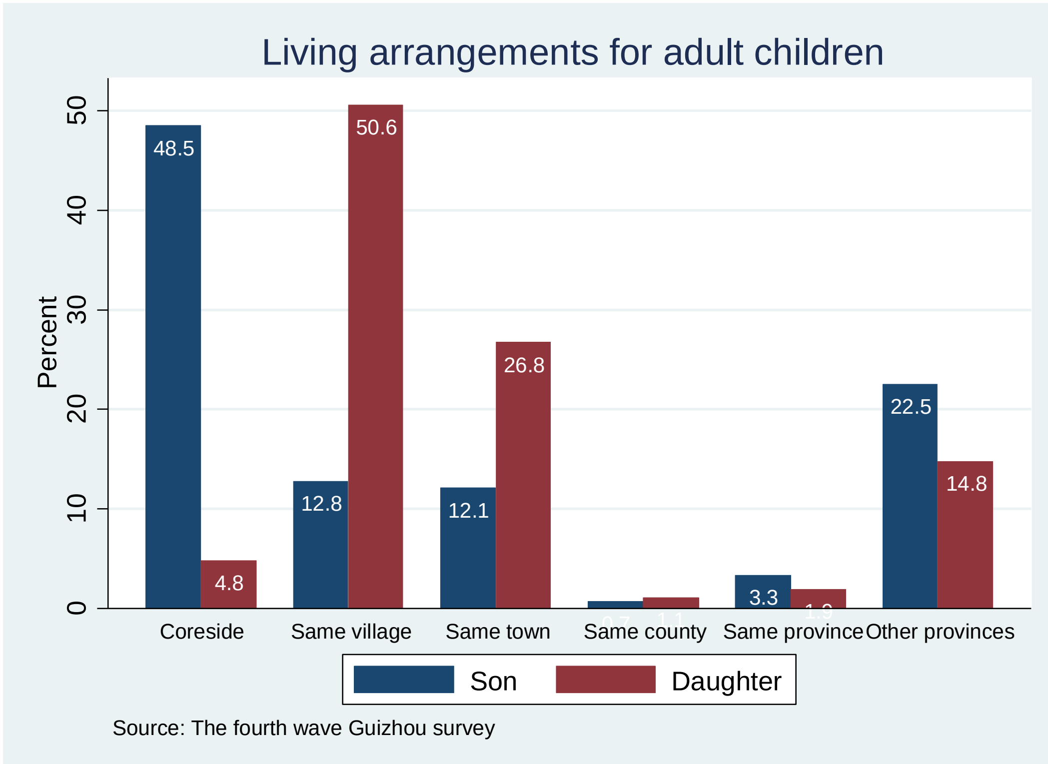
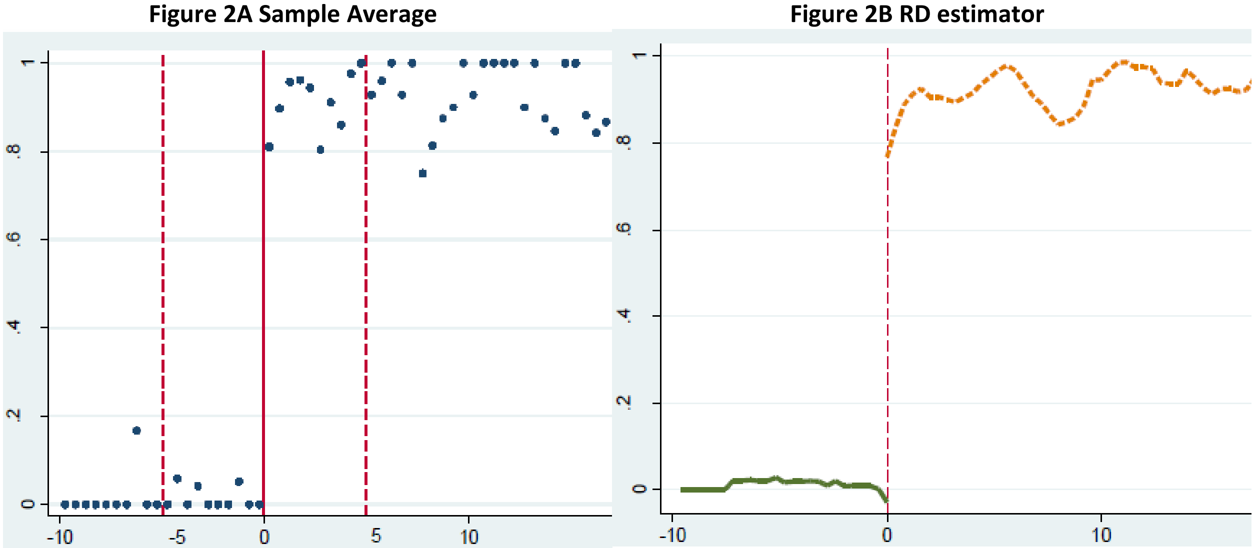


Figure 2. Pension receipt according to normalized age (0=eligibility threshold at age 60)

Figure 2A Sample Average **Figure 2B RD estimator**

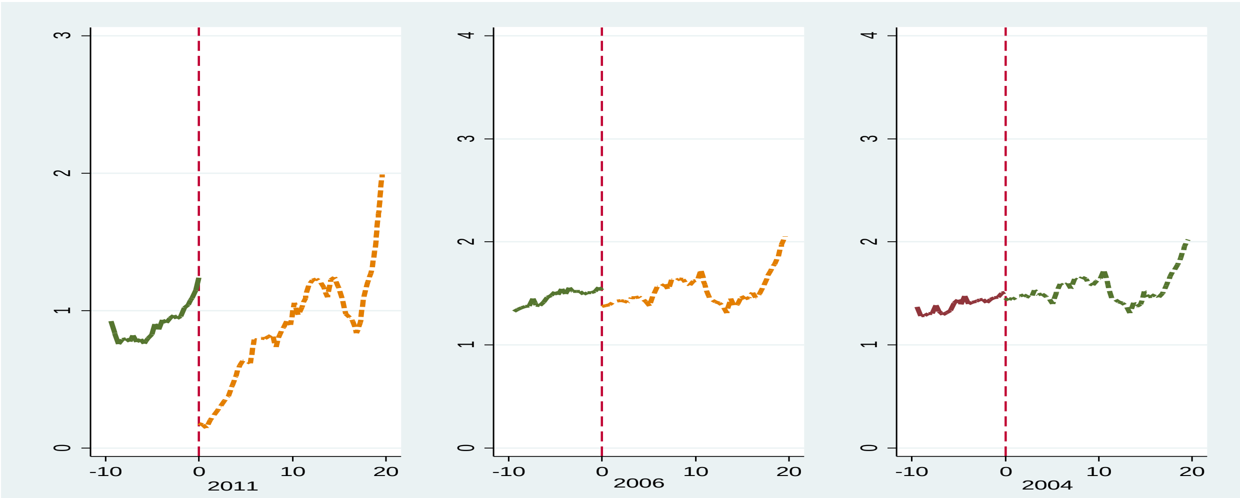


Source: The fourth wave (2011) Guizhou survey.
Notes: This figure shows the relationship between pension take-up rate and respondents' normalized age. Age is normalized based on the day, month, and year information from date of birth.
General graphing notes: 0=eligibility threshold at age 60. The lines are nonlinear fit using rectangular weights on either side using the micro data. The dots represent averages of bins centered at .5 year bins (approximately 180 days). Graphs with different window and bin widths are available upon request.

Figure 3. The proportion of adult sons/daughters co-residing with elderly parents, for treatment (2011, *post-NRPS*) and placebo (2004 and 2006, *pre-NRPS*) years

Source: Three waves of Guizhou survey (2004, 2006, and 2011).
Notes: General notes from Figure 2 apply. This figure shows the relationship between rates of adult sons/daughters co-residing with parents and normalized age of the older parent. We compare the likelihood that adult sons/daughters co-reside with parents around age 60 in 2011 (the 1st column of sub-figures) to that in 2004 (the 3rd column of sub-figures) and 2006 (the 2nd column of sub-figures) around age 60 in 2011.

Figure 4. The number of adult sons with whom the elderly parents co-reside, for treatment (2011, *post-NRPS*) and placebo (2004 and 2006, *pre-NRPS*) years



Source: Three waves of Guizhou survey (2004, 2006, and 2011).
Notes: General notes from Figure 2 apply. This figure shows the relationship between number of adult sons co-residing with elderly parents and normalized age of the parents. We compare the number of adult sons with whom parents co-reside around age 60 in 2011 to that in 2004 and 2006 around age 60 in 2011.

Table 1. Sample summary statistics of the outcome variables

Dependent variable	All	[55,60yrs)	[60,65yrs]	Diff (3)-(2)	All	[55,60yrs)	[60,65yrs]	Diff (3)-(2)
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
Co-residence	**Panel A. Adult son**				**Panel B. Parents with adult son**			
N	1,016	208	217		1,226	190	113	
How many adult sons to live with					.618	.963	.298	-.665***
					(.785)	(.814)	(.561)	(.055)
Whether co-reside w. the elderly	.485	.424	.161	-.262***				
	(.500)	(.495)	(.369)	(.042)				
Whether co-reside w. any child					.794	.700	.637	-.063
					(.405)	(.460)	(.483)	(.056)
Whether co-reside w. any son					.505	.663	.540	-.123**
					(.500)	(.474)	(.501)	(.058)
Co-residence	**Panel C. Adult Children**				**Panel D. Parents with adult children**			
N	1,741	324	325		1,669	237	153	
How many adult sons to live with					.463	.807	.577	-.230***
					(.691)	(.777)	(.640)	(.083)
Whether co-reside w. the elderly	.303	.298	.114	-.184***				
	(.460)	(.458)	(.318)	(.031)				
Whether co-reside w. any child					.814	.681	.585	-.096*
					(.389)	(.467)	(.495)	(.054)
Whether co-reside w. any son					.375	.609	.496	-.113**
					(.479)	(.489)	(.502)	(.056)
Service consumption					**Panel E. Parents with adult son**			
N					1,226	190	113	
Paying for haircut or machine repair					.356	.016	.443	.427***
					(.479)	(.125)	(.499)	(.038)
Purchasing coal for heating and cooking					.482	.589	.619	.030
					(.500)	(.493)	(.488)	(.058)
Being assisted on the daily life by adult children due to illness					.052	.060	.027	-.033
					(.222)	(.238)	(.161)	(.028)

Source: The fourth wave (2011) Guizhou survey.

Notes: [-5yrs,0) and [0,5yrs] mean parents' age relative to the 60yrs cutoff. N denotes sample size.

Table 2. First Stage: Pension Receipt

Dependent variable	N	OLS	RD estimate
	(1)	(2)	(3)
Panel A. Sample of all respondents (*N=5189*)			
Indicator of Pension Receipt	5189	.907***	.861***
		(.014)	(.067)
Panel B. Sample of parents with adult children (*N=1669*)			
Indicator of Pension Receipt	1669	.904***	.650***
		(.023)	(.137)
Panel C. Sample of parents with adult son (*N=1226*)			
Indicator of Pension Receipt	1226	.906***	.765***
		(.024)	(.126)
Panel D. Sample of adult children (*N=1741*)			
Indicator of Pension Receipt	1741	.934***	.769***
		(.014)	(.126)
Panel E. Sample of adult sons (*N=1016*)			
Indicator of Pension Receipt	1016	.928***	.873***
		(.018)	(.051)
Panel F. Sample of adult daughters (*N=725*)			
Indicator of Pension Receipt	725	.900***	.504***
		(.021)	(.132)

Source: The fourth wave (2011) Guizhou survey.

Notes:

[1] The optimal bandwidths are calculated implementing the CCT bandwidth selector.

[2] The nearest-neighbors match procedure is implemented to compute the variance-covariance matrix estimator.

[3] The RD regression is weighted using rectangular weights. The RD estimates are bias-corrected with robust standard errors.

[4] The regressions control for ethnicity, cadre and party membership, education, marital status, religious beliefs, adult child age, different types of family assets, year of birth and village fixed effects.

*** Significant at the 1 percent level.

** Significant at the 5 percent level.

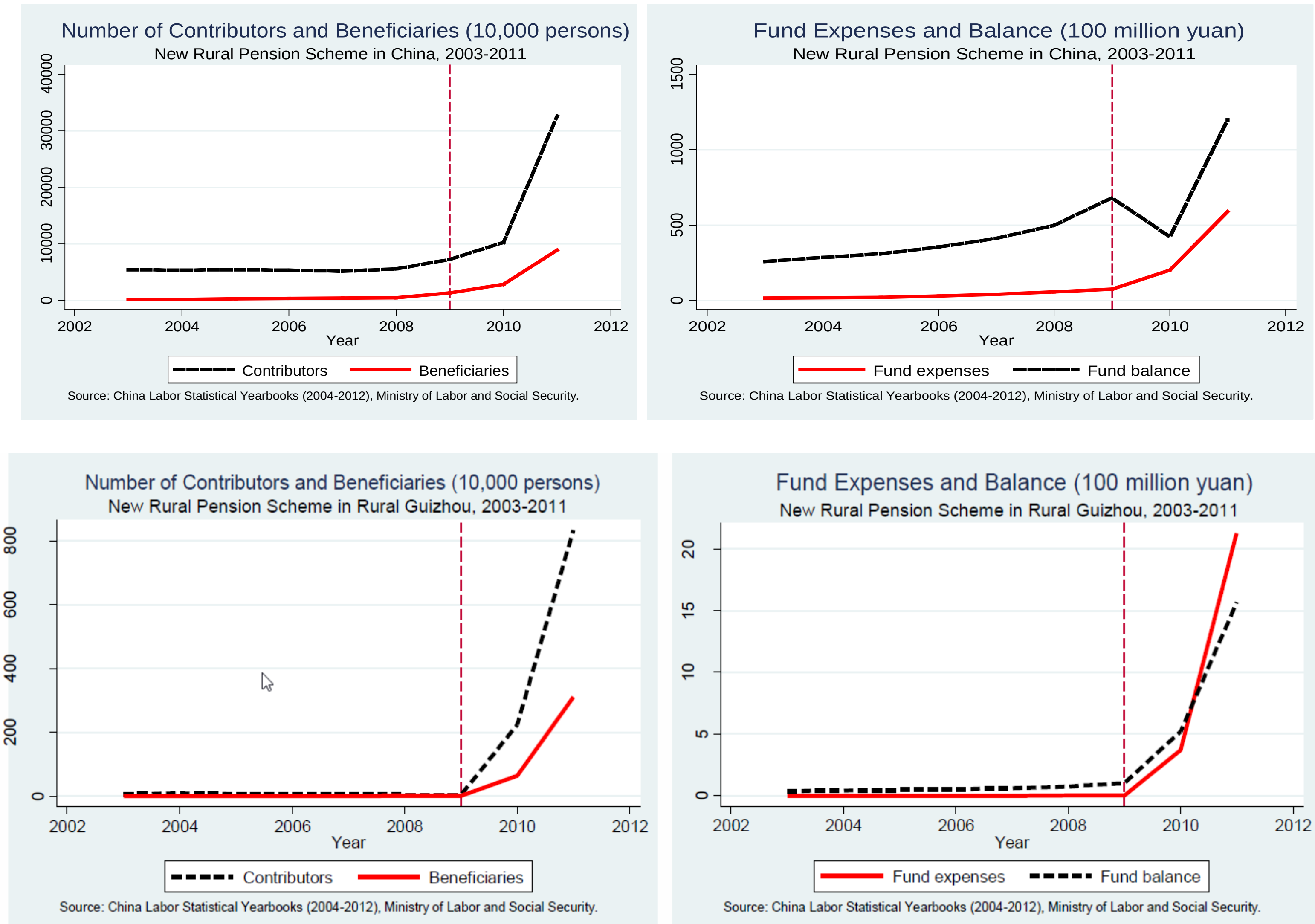
* Significant at the 10 percent level.

Table 3. Main results: Living arrangement of extended household members

	Post implementation (2011)				Before implementation (2004, 2006)			
Dependent variable	**N**	**OLS**	**RD reduced form**	**RD LATE**	**N**	**OLS**	**RD reduced form**	**RD LATE**
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A. Sample of adult sons (*N=1016*) and daughters (*N=725*) respectively								
Co residing with adult parent(s)								
Sons	1016	-.262***	-.389***	-.266*	1208	.021	-.033	-.041
		(.042)	(.081)	(.140)		(.036)	(.127)	(.219)
Daughters	725	-.071**	-.059	.063	1005	-.027	-.035	-.048
		(.030)	(.060)	(.059)		(.064)	(.074)	(.099)
Panel B. Sample of parents with adult son (*N=1226*)								
#adult sons living together	1226	-.244***	-.374*	-.500*	2357	-.132	-.054	-.073
		(.087)	(.205)	(.279)		(.129)	(.112)	(.194)
Whether co-reside w. any child	1226	-.060	-.147	-.322	2357	-.094	-.072	-.083
		(.056)	(.112)	(.250)		(.144)	(.193)	(.151)
Whether co-reside w. any son	1226	-.121**	-.263***	-.311***	2357	-.072	-.047	-.057
		(.058)	(.097)	(.118)		(.114)	(.124)	(.161)

Source: Three waves of Guizhou survey (2004, 2006, 2011).
Notes: We compare living arrangement indicators around age 60 in 2011 (left table) to those in 2004 and 2006 around age 60 in 2011 (right table). The observations for the right table are from 2004 and 2006. General notes from Table 2 apply.

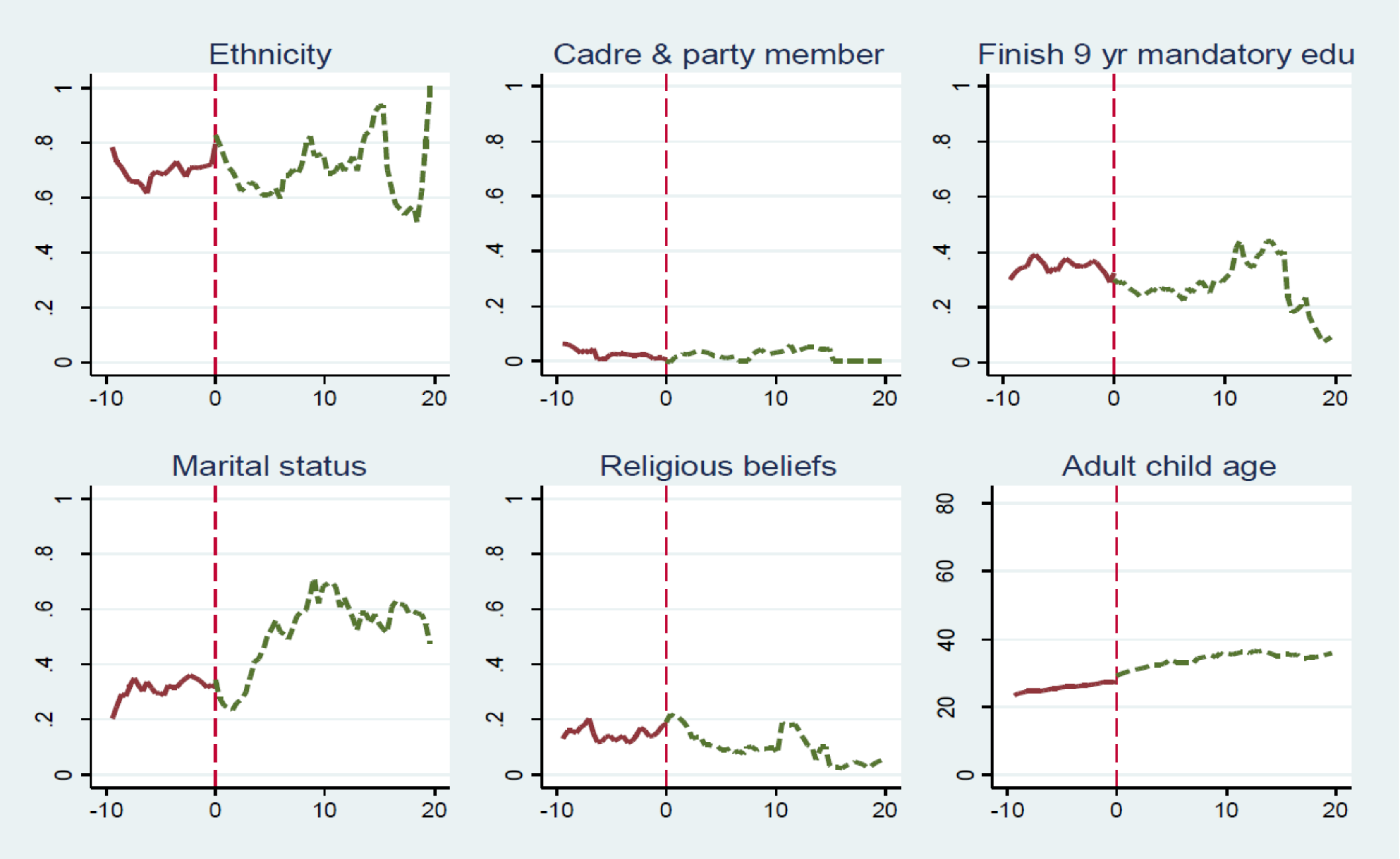
Online Appendix Figure 1 – Statistics on the Rural Pension System in China and Guizhou Province (2003-2011)



Source: China Labor Statistical Yearbooks (2004-2012), Ministry of Labor and Social Security.

Notes: The NRPS in China initiated at the end of 2009. The non-zero figures before 2009 represent the old pension policy that covered a tiny proportion of rural residents, mainly in developed regions in China. Almost no social pension existed in rural Guizhou before 2009.

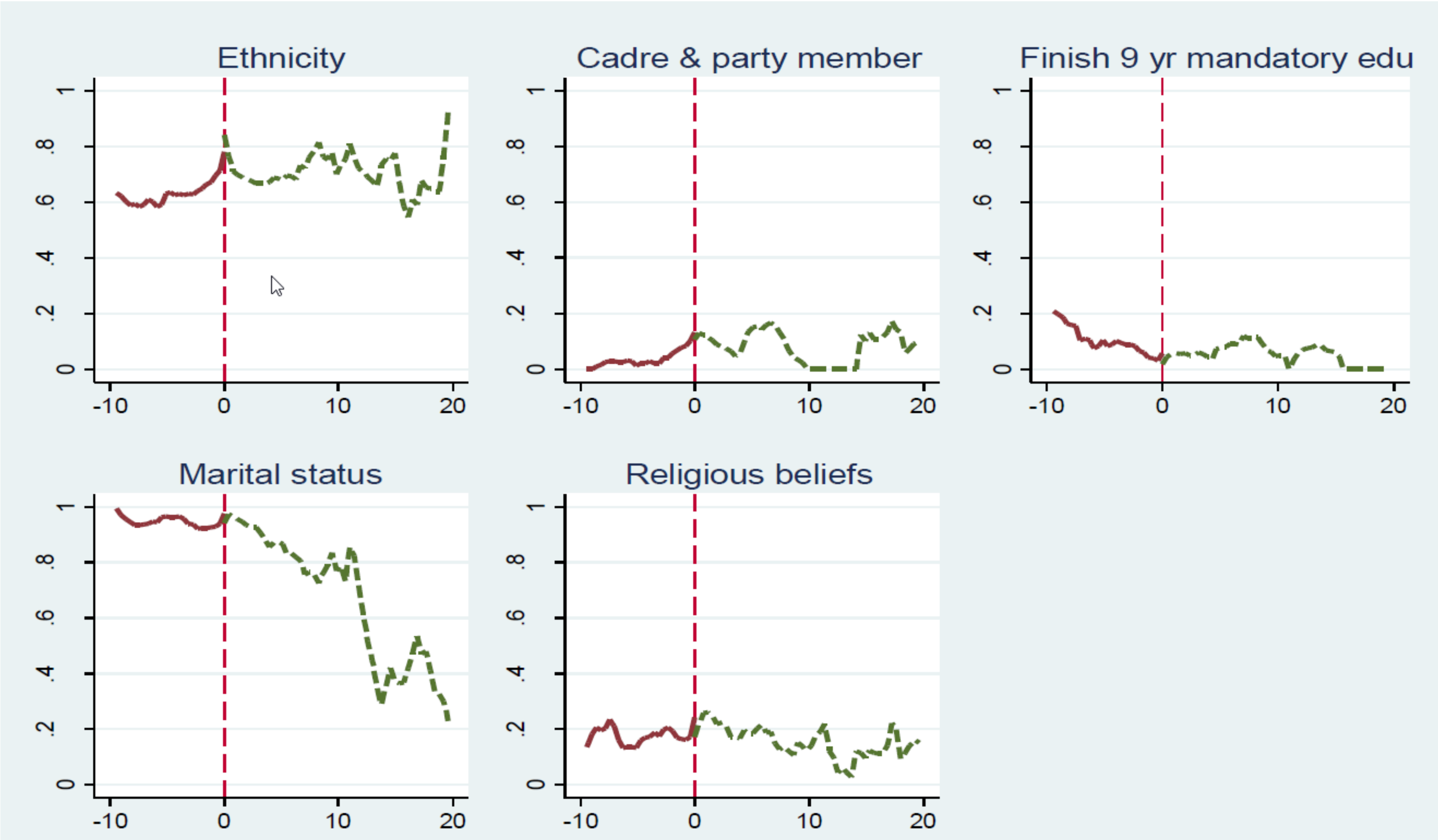
Online Appendix Figure 2a Baseline demographic and socioeconomic characteristics of adult children



Source: The second wave (2006) Guizhou survey.

Notes: This figure shows main baseline demographic and socioeconomic characteristics of adult children and normalized age of the older parent. All these characteristics (except adult child age) are dummy variables, indicating whether an adult child is from the main ethnic group, holds a village leader position or communist party membership, finishes nine-year mandatory education, gets married, has religious beliefs, respectively. To save space, figures for other covariates, i.e., different types of family assets, controlled in the analysis are available upon request. General notes from Figure 2 apply.

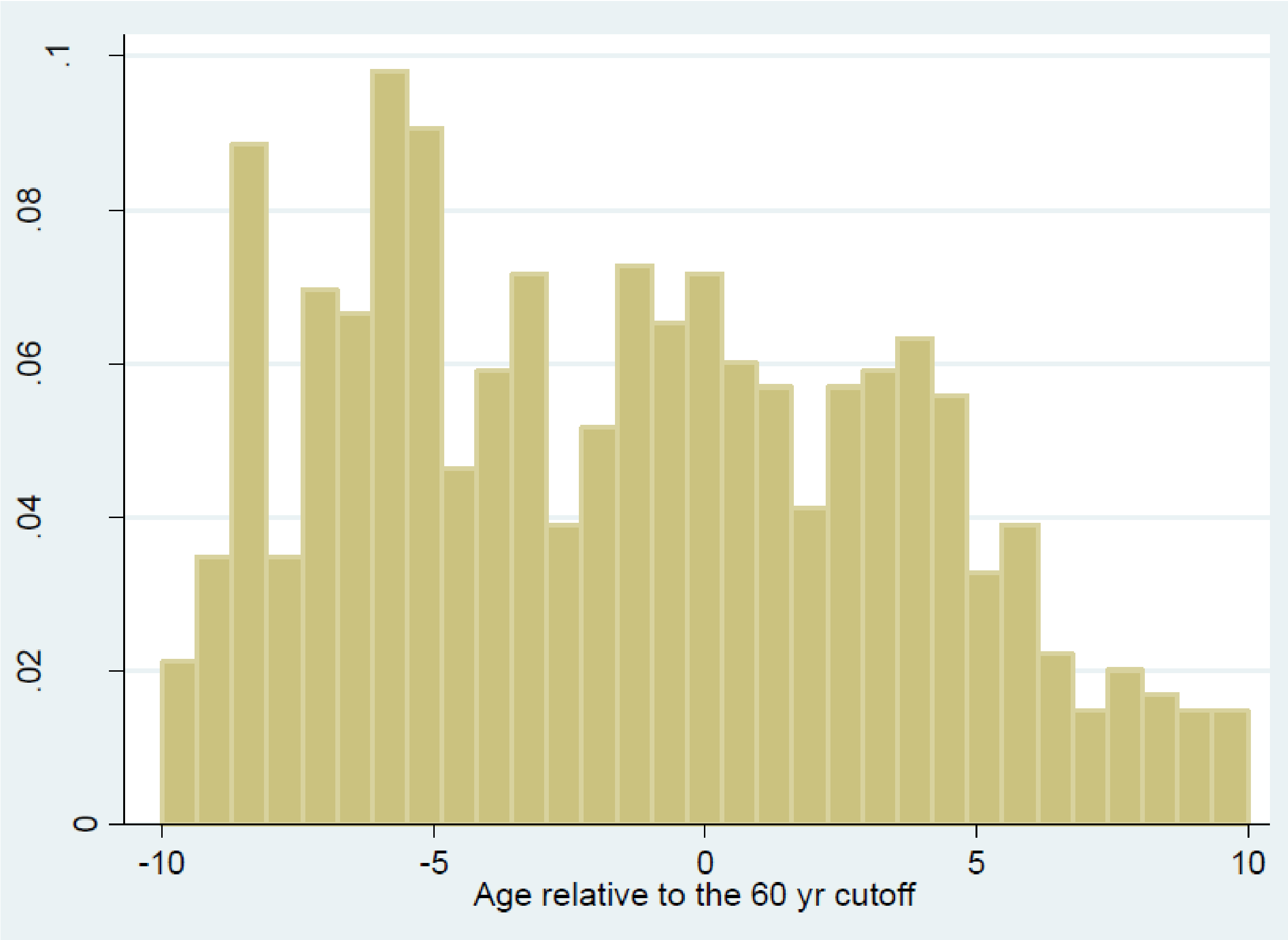
Online Appendix Figure 2b Baseline demographic and socioeconomic characteristics of parents



Source: The second wave (2006) Guizhou survey.

Notes: This figure shows main baseline demographic and socioeconomic characteristics of parents and normalized age of them. All these characteristics are dummy variables, indicating whether a parent is from the main ethnic group, holds a village leader position or communist party membership, finishes nine-year mandatory education, gets married, has religious beliefs, respectively. To save space, figures for other covariates, i.e., different types of family assets, controlled in the analysis are available upon request. General notes from Figure 2 apply.

Online Appendix Figure 3 Histogram of age and the density discontinuity test



Density Discontinuity Test

Log discontinuity estimate	Standard error
.085	.091

Source: The fourth wave (2011) Guizhou survey.

Notes: This table presents results testing the density discontinuity of age distribution. The test statistic is based on McCrary (2008).

Online Appendix Table 1 - Income and inequality indicators of Guizhou household survey in 2004, 2006, 2009 and 2011

	2004	2006	2009	2011
Per capita real annual income (in RMB)	1404	1817	2855	2911
Income below poverty line of US $1.25 per day using 2005 PPP (%) (P0)	71.3	64.1	52.7	47.6
Income inequality (Gini)	43.1	48.2	55.2	58.6

Source: The four wave Guizhou survey.

Notes:

[1] RMB = yuan renminbi. PPP = purchasing power parity. P0 denotes the standard Foster-Greer-Thorbecke poverty measure of headcount ratio.

[2] The 2005 PPP exchange rate is at the "China-rural" level. See http://iresearch.worldbank.org/PovcalNet/jsp/index.jsp. The Poverty lines for 2004-2011 are adjusted according to the published annual inflation rate in various issues of *China Statistic Year Book,* published by China's National Bureau of Statistics.

Online Appendix Table 2 – Effect on main outcomes selection procedures

Dependent variables	With different selection procedures			With different bandwidths			With different polynomial orders		
	CCT	IK	CV	Optimal bandwidth	½ optimal bandwidth	2 optimal bandwidth	Linear	Quadratic	Cubic
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Panel A. Sample of parents with adult children (*N=1669*)									
Pension receipt	.650***	.674***	.626***	.650***	.517***	.737***	.650***	.567***	.587***
	(.137)	(.083)	(.107)	(.137)	(.146)	(.067)	(.137)	(.177)	(.085)
Panel B. Sample of parents with adult son (*N=1226*)									
#sons living together	-.500*	-.531***	-.503***	-.500*	-1.168***	-.412*	-.500*	-.817**	-.472**
	(.279)	(.192)	(.186)	(.279)	(.293)	(.228)	(.279)	(.419)	(.221)
Whether co-reside w. any child	-.322	-.282**	-.181	-.322	-1.442***	-.277	-.322	-.396	-.199
	(.250)	(.123)	(.130)	(.250)	(.493)	(.186)	(.250)	(.280)	(.170)
Whether co-reside w. any son	-.311***	-.296***	-.240***	-.311***	-.820***	-.494***	-.311***	-.362**	-.467**
	(.118)	(.074)	(.062)	(.118)	(.251)	(.128)	(.118)	(.148)	(.194)
Panel C. Sample of adult children (*N=1741*)									
Co-residing with adult parent(s)	-.232**	-.249***	-.302***	-.232**	-.404**	-.275***	-.232**	-.294*	-.267**
	(.118)	(.083)	(.071)	(.118)	(.193)	(.082)	(.118)	(.177)	(.107)

Source: The fourth wave (2011) Guizhou survey.

Notes:

[1] CCT, CV and IK are three bandwidth selectors introduced in Section 3.2.

[2] The optimal bandwidths are calculated using the CCT bandwidth selector.

[3] Each term of the polynomial is interacted with the age eligibility dummy, allowing each term to have different parameters on either side of the age cutoff.

[4] General notes from Table 2 apply.

Online Appendix Table 3 – Additional results: Service consumption by parents with adult son (N=1226)

Dependent Variable	**N**	**OLS**	**RD reduced form**	**RD LATE**
	(1)	**(2)**	**(3)**	**(4)**
Paying for haircut or machine repair	1226	.435***	.237*	.332*
		(.038)	(.133)	(.171)
Purchasing coal for heating and cooking	1223	.019	.291**	.402**
		(.059)	(.124)	(.184)
Not getting formal medical treatment in the past 12 months when recommended	917	-.298***	-.429**	-.463**
		(.072)	(.205)	(.219)
Receiving in-family medical services	1223	-.041	.179	.236
		(.059)	(.153)	(.198)
Receiving long-term assistance in daily life by adult children when ill	1024	-.031	-.316**	-.439**
		(.028)	(.158)	(.222)
New Cooperative Medical Scheme (NCMS) enrollment	1226	-.026	.122	.162
		(.017)	(.074)	(.100)

Source: The fourth wave (2011) Guizhou survey.

Notes:

[1] For some estimations the number of observations are smaller because they are conditional on being recommended for medical treatment or being ill.

[2] We group paying for haircut or machine repair in the past 12 months as a dummy indicator.

[3] The survey also collects information on coal purchase. More than half of the families still collect firewood as heating and cooking fuel. Stronger purchasing power and absence of family labor as a result of family division may increase coal purchase.

[4] In rural China, financial difficulties often prevent patients from receiving necessary treatment. The survey asks the perceived access to needed medical services when recommended by a doctor. We define responses to this question by a dummy indicator that equals 1 if not getting formal medical treatment in the past 12 months when recommended and 0 otherwise.

[5] General notes from Table 2 apply.

Online Appendix Table 4 – Effect on main outcomes allowing changes at age 59

Dependent variables	N	OLS	RD reduced form/ first stage	RD LATE
	(1)	**(2)**	**(3)**	**(4)**
Panel A. Sample of parents with adult children (*N=1669*)				
Pension receipt	1669	.626***	-.152	-
		(.038)	(.126)	
Panel B. Sample of parents with adult son (*N=1226*)				
#sons living together	1226	-.659***	.353	-.679
		(.071)	(.250)	(.559)
Whether co-reside w. any child	1226	.017	-.060	-4.084
		(.032)	(.104)	(2.751)
Whether co-reside w. any son	1226	-.052	.152	-.616
		(.056)	(.161)	(.679)
Panel C. Sample of adult children (*N=1741*)				
Co-residing with adult parent(s)	1741	-.312***	.141	-.286
		(.035)	(.101)	(.250)

Source: The fourth wave (2011) Guizhou survey.
Notes: General notes from Table 2 apply.

Online Appendix Table 5 – Comparing main RD results on intergenerational living arrangement with results using the diff-in-diff (DD) strategy

Dependent variable	N (wave 2011)	RD LATE	N (waves 2006, 2011)	DID reduced form
	(1)	(2)	(3)	(4)
Panel A. Sample of adult sons and daughters, respectively				
Co residing with adult parent(s)				
Sons	1016	-.266*	1638	-.242**
		(.140)		(.097)
Daughters	725	.063	1236	-.057
		(.059)		(.072)
Panel B. Sample of parents with adult sons				
#adult sons living together	1226	-.500*	2452	-.432*
		(.279)		(.246)
Whether co-reside w. any child	1226	-.322	2452	-.297
		(.250)		(.231)
Whether co-reside w. any son	1226	-.311***	2452	-.274**
		(.118)		(.126)

Source: The second wave (2006) and the fourth wave (2011) Guizhou survey.

Notes: The right section of this table estimates DD estimator for the impact of pension eligibility on intergenerational living arrangements. Different from main RD estimations (in the left part of this table) that are only able to use the 2011 wave (the year with discontinuity in pension policy), the DD estimations pool together the 2006 and the 2011 waves of survey. Parents who were eligible to receive pension in the 2011 wave of survey are assigned to the treatment group (TREAT=1). AFTER=1 if a respondent is in the 2011 wave and AFTER=0 if in the 2006 wave. The reported point estimates and standard errors correspond to the DD estimator, i.e. the interaction term TREAT*AFTER. General notes from Table 2 apply.