Social Security and Elderly Living Arrangements

Evidence from the Social Security Notch

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ABSTRACT

Previous studies of the effect of Social Security on elderly living arrangements generally have relied on data from the distant past or differences in benefits across families or cohorts that potentially were correlated with other determinants of living arrangements. Using data from the 1980–99 Current Population Surveys, we attempt to isolate the causal effect of Social Security on living arrangements with an instrumental-variable approach that relies on the large shifts in benefits for cohorts born from 1910–21, the so-called Social Security notch. Over all elderly households, the estimated elasticity of living with others with respect to Social Security income is –0.4, with elasticities of –1.3 and –1.4 for the widowed and divorced, respectively; most of the effects on living arrangements appear to be concentrated among the lesser educated as well. Our estimated elasticities are substantially larger than those from previous studies and suggest that reductions in current benefits would alter living arrangements significantly.

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I. Introduction

Social Security is the largest and, in the view of many, the most successful social program in the United States. However, it has been well documented that, at the current level of payroll tax finance, the program cannot sustain the current generosity of benefits in the long run. This has ushered in a number of policy proposals for Social Security reform, some of which have advocated a reduction in benefits, and prompted concern that reforms might reverse the gains made by the program over the past 40 years in increasing the well-being of the elderly.

To assess the net effect on elderly well-being, one must incorporate the extent to which the elderly respond to benefit changes along a number of margins. When faced with a reduction in Social Security income, elderly can stay in the labor force longer, supply more post-retirement hours, reduce consumption, or substitute shared for independent living arrangements. While economists have given great attention to the effect of the program on labor force participation and saving behavior, there has been comparatively little attention on the effect on living arrangements, an important element of well-being.

There is a small existing literature on the sensitivity of elderly living arrangements to income. But this literature has produced a wide range of estimated elasticities of the likelihood of living in a shared arrangement with respect to income, from close to zero to −1. This may reflect the inherent difficulties in separating the impacts of income per se from the other factors that determine the desire of the elderly to live alone. The studies that most carefully have addressed this concern have focused on historical changes in the retirement income available to the elderly, but these changes occurred in a very different social and economic environment than today's, which may have implications for the sensitivity of living arrangements to incomes. Finally, the past literature almost exclusively has been focused on widows. This is understandable given their historically high poverty rate and prominence in policy debates. However, the share of the elderly who are widowed is falling over time, with particularly rapid growth in elderly divorcees. This suggests that the time is ripe for a broader look at the sensitivity of living situations to incomes for all groups of elderly.

Our paper makes two important contributions. First, we outline the econometric problems in the previous literature and propose an instrumental variable procedure to circumvent these difficulties. Specifically, we examine the effect on elderly living arrangements of the large changes in Social Security benefits that affected birth cohorts from 1910 through 1921. The early cohorts in this range saw enormous increases in their Social Security benefits due to double indexation of the system in the early 1970s. This double indexing was ended in the 1977 Amendments to the Social Security Act that generated the so-called "benefits notch." The 1977 law grandfathered all individuals born before January 1, 1917, under the old benefit rules, but those born in 1917–21 received benefit reductions that were as much as 20 percent lower than observationally equivalent individuals in the 1916 birth cohort. After 1921, benefits were roughly constant in real terms. It is this variation that was first identified by Krueger and Pischke (1992) as a fruitful means of identifying the behavioral effects of Social Security, in their case in the context of retirement decisions. We follow their methodology to define an instrumental variable for observed Social Security benefits. In a related paper, Snyder and Evans (2002) used the notch to examine the effect of income on mortality. Second, we focus in our analysis on the 1980–99 period, which is much more recent than other studies and, therefore, provides a better benchmark for thinking prospectively about policy changes.

We do so by using data on the living arrangements of the elderly from the Current Population Survey (CPS) from 1980–99. The large samples in this nationally representative survey allow us to use differences across birth cohorts to carefully identify the impact of legislated benefits changes on living arrangements and to separately assess the impacts on different groups of elderly persons.

We find that the likelihood of living with others is very sensitive to incomes for elderly widows and divorcees. For widows, we estimate an elasticity of living with others with respect to Social Security income on the order of −1.3, and for divorcees an elasticity of −1.4. The likelihood of living with others for those who were never married is related to benefit levels only modestly, with an insignificant elasticity of –0.4, and the decision of married couples to live with others is not sensitive to income levels. Averaging over all elderly, we obtain an elasticity of living with others with respect to benefits of –0.4. These elasticities are much larger than those found in previous studies, which may reflect our improved identification strategy, as well as more fluidity in living arrangements in more recent times. In addition, we find that the effect on living arrangements is concentrated among elderly with a high school education or less; there is little or no effect for those with greater education levels. Overall, the findings suggest that living arrangements are elastically demanded for nonmarried elderly, privacy is a normal good, and that reductions in Social Security benefits would alter significantly the living arrangements of the elderly. Assuming current and prospective elderly would respond in a manner similar to those affected by the notch, our estimates imply that a 10 percent cut in Social Security benefits would lead more than 600,000 independent elderly households to move into shared living arrangements.

The paper is organized as follows. The next section gives background on the Social Security system and the previous literature. Section III describes the CPS data and the construction of the instrumental variable. Section IV discusses the empirical results. The implications of our results for Social Security reform are discussed in Section V. There is a brief conclusion.

II. Background

The well-known rise in independent living by the elderly was a striking change in economic behavior in the twentieth century (Kobrin 1976; Kramarow 1995; Macunovich et al. 1995; McGarry and Schoeni 2000; Wolf 1995; Wolf and Soldo 1988). One factor often hypothesized to explain this trend was the increase in pension income of the elderly, particularly from the adoption in 1935 and expansion of Social Security. Probably the most important early study of the effect of Social Security on elderly living arrangements was by Michael, Fuchs, and Scott (1976). They analyzed cross-sectional data from states in 1970 and estimated that the elasticity of the state proportion of widows living with others with respect to state mean Social Security benefits ranged from –0.45 to −1.05, depending on the set of explanatory variables. Estimates from subsequent studies of the income elasticity of the proportion of elderly living in shared arrangements have varied substantially. On the low end, Börsch-Supan, et al. (1992) found that increases in income did not raise the probability that elderly lived in a shared arrangement, but decreased the chance of institutionalization. However, a number of studies have found higher estimates, with Costa (1999) estimating an elasticity of −1 from the state Old Age Assistance (OAA) program in the 1940s.

Although previous studies differed along a number of dimensions, including the type of household studied, the estimator, data source, level of aggregation, and the definition of the income variable, an important reason for the differences in estimated elasticities is due to differences in econometric identification. In particular, there are a number of potential econometric pitfalls when estimating the effect of Social Security income on elderly living arrangements. First, for those studies that used measures of income that were broader than Social Security (Börsch-Supan et al. 1992; Schwartz, Danziger, and Smolensky 1984; and Macunovich et al. 1995), some components of non-Social Security income may be endogenous (for example, post-retirement labor supply and asset spend-down). Second, estimates from studies that relied on actual Social Security income (other than Costa 1997; Costa 1999) may have been confounded by omitted variables correlated with observed Social Security income as well as with living arrangements. For example, Social Security benefits are primarily a function of average lifetime earnings, and higher lifetime earnings, separate from Social Security, should raise the demand for independent living if privacy is a normal good; this would tend to bias Ordinary Least Squares (OLS) estimated elasticities away from zero. More subtly, Costa (1997) and Costa (1998) argued that the prospects of increased living independence among the elderly may have made retirement more attractive. But earlier retirement implies a reduced average Social Security benefit level, for a given earnings history, so that there is a direct feedback from independent living to average benefit levels over time.

The studies of Costa (1997) and Costa (1999) did not suffer from these limitations, but were focused on very different time periods, either the early 20th century or the 1940s. The general changes in both the economy and society over the past century suggest that there may be quite different responsiveness to Social Security incomes in living arrangements now than there has been in the past. Increased social mobility, changes in the structure and availability of housing, rapidly rising female labor force participation, and changes in access to both home services and shopping all imply that the elderly make their decisions in a very different environment today than they did 50 or 100 years ago.

A particularly important change has been in the composition of the elderly. In 1960, 19 percent of elderly men and 53 percent of elderly women were widowed; fewer than 2 percent of the elderly were divorced. By 1995, the share of the elderly who were widowed had fallen, with a rapid rise in the share of the elderly who were divorced; the share who were married also has risen somewhat.[1] This suggests that it is important to examine how all groups of elderly, and not just widows, respond to benefit changes in their living arrangements.

1. These figures were taken from Table 40 of the 2000 *Statistical Abstract of the United States* (United States Department of Commerce 2000), Table 42 of the 1990 *Statistical Abstract of the United States* (United States Department of Commerce 1990), and Table 39 of the 1980 *Statistical Abstract of the United States* (United States Department of Commerce 1980).

III. Data

A. Sample Selection

This study uses data from the Current Population Surveys (CPS) of March 1980 through 1999. Each file is a cross-sectional, nationally representative sample of households. To construct our sample, we first assign families within the CPS. A family is defined as the household head, his or her spouse, and any children of the household head who are living in the household and under the age of 19. We assume any other member of the household is his/her own family for the purpose of our definition. These families serve as our observational unit. Note that there may be more than one "family" in a given CPS "household" (for example, if there are multiple nonmarried elderly living together).

To assign Social Security benefits to families, it is necessary to assign a "Social Security beneficiary." Our default is to assign this person to be the male in the family who is older than 65. If there is no such male, the Social Security beneficiary is assigned to be the oldest never-married female in the family. These two groups consist of people who are likely to have had Social Security benefits based on their own earnings history, rather than that of their spouse.

If there is neither a male nor a never-married female older than 65, we assign the Social Security beneficiary to be the divorced or widowed female who is older than age 62. We assume that her Social Security benefits are based on the earnings of her former or deceased spouse. We further assume that this spouse was three years older than she is, the median spousal age difference for widowed and divorced elderly tabulated from the 1982 Social Security New Beneficiary Survey, so that the "age of the Social Security beneficiary" is this woman's age plus three for the purposes of calculating our instrument (discussed below).

These restrictions lead to a sample consisting of any families that contain at least one male or never-married female over the age of 65, or that contain a widowed or divorced female over the age of 62. We select this age group because most people who are eligible to collect Social Security benefits begin doing so by age 65. The main sample is based on 230,045 family-year observations from the March 1980–99, CPS. Because the instrument varies primarily by year of birth, we aggregate these data into age-by-year-of-birth cells, producing either 473 or 494 cells depending on whether widows and divorcees are included in the analysis.[2] The average cell size was 466 families. We include both sexes in our data set.[3]

Finally, we create a variable to describe whether each family is living independently or with others. We consider the two individuals in a married couple residing together to be living independently. That is, even though one possible outcome of a change in Social Security income is a change in cohabitation of two individuals in the

2. One limitation of our analysis is that year of birth is not directly observable in the CPS. Instead, we measure it as CPS year minus age minus one. Because the CPS is fielded in early March, about 2/12ths of each year-of-birth cohort belongs to the next cohort. Unfortunately, there is no obvious way to circumvent this problem.

3. For widows, 84 percent of the observations are female so our results are very comparable to the previous literature; indeed, if we estimate models for females only, we obtain estimates almost identical to those presented below.

form of legal marriage, we assume that marriage is exogenous.[4] However, we do allow for the possibility that married couples share a living arrangement with another (third) individual (for example, a sibling of one of the spouses or a child). Any married couples that reside with another adult are considered to be in a shared living arrangement. For those who are not currently married, any elderly person who is living with others in their household is not considered to be living independently.[5] Table 1 shows sample means for selected variables, with standard deviations in parentheses. The mean proportion living independently ranges from about 0.59 for never-married individuals to 0.84 for married couples, and is 0.74 pooled over all families.

B. Construction of the Instrument

As highlighted earlier, the fundamental problem with previous studies of the impact of Social Security on living arrangements is that benefit levels are correlated with factors that might otherwise influence living arrangements. Partly this reflects differences across individuals, which is abstracted away in our cell-level analysis. But there are also important average differences across cohorts, such as differences in average lifetime earnings or tastes for independent living (that feedback to retirement decisions and, therefore, to Social Security benefit levels), which are correlated with living arrangements as well.

Our goal in this paper, therefore, is to construct an instrument for Social Security benefits that is independent of other factors that differ across year-of-birth cells; that is, we construct an instrument that is identified solely by legislative changes in benefits and not from differences in birth-cohort characteristics. We do so by exploiting the large changes in Social Security benefits documented in the introduction: the enormous run-up in benefits for birth cohorts from 1910 through 1916, followed by the striking decline for those birth cohorts from 1917 through 1921. Over this relatively short period, otherwise similar workers saw enormous unanticipated and permanent swings in their level of Social Security entitlement, allowing us to identify the effects of Social Security independently from individual or cohort characteristics.

Specifically, the 1977 law that created the notch generated variation in benefits along three dimensions: year of birth, retirement age, and earnings level. First, the formula for calculating the primary insurance amount (PIA) from Average Indexed Monthly Earnings (AIME) changed, which induced variation in benefits by year of birth. Second, for *each* of the different notch birth years, 1917–21, the amount of benefit reduction changed depending on the claiming age. That is, the law change induced retirement-age-by-year-of-birth variation in benefits. This is what Krueger and Pischke (1992) exploited in their study of the impact of the notch on labor supply. Third, the law raised the covered-earnings maximum used to calculate AIME so that after 1977 a greater fraction of annual earnings for high-wage workers entered

4. In a previous version of the paper, Engelhardt, Gruber, and Perry (2002), we tested this assumption with data from the June 1980 and 1985 CPS by regressing the proportion of individuals who were married on the instrument and demographic variables (discussed below) and found no evidence that benefits affected marriage.

5. We attempted to further decompose our data into those living with their own children versus those living with others. But, unfortunately, changes in the construction of the CPS family relationship variables halfway through our sample left us unable to draw any conclusions as to relative shifts across these groups.

Table 1
Sample Means for Selected Variables, with Standard Deviations in Parentheses, for Years of Birth from 1900–33, Using the 1980–99 March CPS

	(1)	(2)	(3) Sample	(4)	(5)
Variable	Pooled	Widowed	Divorced	Married	Never Married
Proportion living independently	0.735	0.710	0.698	0.836	0.586
	(0.042)	(0.047)	(0.114)	(0.063)	(0.118)
Proportion in shared arrangement	0.265	0.290	0.302	0.164	0.414
	(0.042)	(0.047)	(0.114)	(0.063)	(0.118)
Social Security income	5323	4402	3988	7838	3766
	(704)	(691)	(941)	(1217)	(654)
Proportion of heads with less than high school diploma	0.481	0.481	0.437	0.504	0.415
	(0.087)	(0.086)	(0.149)	(0.108)	(0.104)
Proportion of heads with high school diploma	0.261	0.315	0.301	0.163	0.277
	(0.040)	(0.067)	(0.123)	(0.050)	(0.094)
Proportion of heads with some college	0.130	0.118	0.149	0.146	0.108
	(0.039)	(0.036)	(0.090)	(0.057)	(0.067)
Proportion of heads with college degree	0.120	0.078	0.105	0.178	0.184
	(0.038)	(0.022)	(0.076)	(0.065)	(0.078)
Proportion of heads who are female	0.563	0.835	0.720	—	0.647
	(0.164)	(0.055)	(0.125)		(0.127)
Proportion of heads who are white	0.897	0.884	0.872	0.915	0.894
	(0.020)	(0.034)	(0.080)	(0.033)	(0.068)
Number of observations	494	494	494	473	473

Note: The table shows means calculated from the 494 age-by-year-of-birth cells based on the underlying sample of 230,045 family-year observations from the 1980–99 March CPS, as described in the text. Standard deviations are in parentheses. Social Security income is expressed in 1982–84 dollars.

AIME. As we document below, this induced earnings-level-by-year-of-birth variation in benefits.[6]

Our strategy for constructing the instrumental variable is to create a measure of Social Security benefits entitlement that is identical for each birth cohort *except for changes in the benefits law*. To create such an instrument, we first assign an earnings history to the 1916 birth cohort. The *Annual Statistical Supplement* produced by the Social Security Administration contains the median Social Security earnings by gender for five-year age groups on a yearly basis for the current year as well as years past. We used median male earnings from these tables. We assign median earnings at age 22 (from the median earnings for ages 20–24 in 1938), age 27 (from median earnings for ages 25–29 in 1943), and so forth, in five-year intervals. We then assume a linear trend in earnings in between these five-year intervals. We use this method through age 60 and then assume earnings grow with the Consumer Price Index (CPI) rate of inflation for ages 60–65. We do not use median earnings for workers over 60 because many of these workers have entered "bridge" jobs, so that the median worker's earnings at these ages may not be representative of workers who have remained in their lifetime jobs through age 65. This generates an earnings history for a median male earner in the cohort born in 1916. We use the same earnings profile even when assigning benefits to never-married females, because we assume that their earnings profile would more closely resemble that of a male worker than that of the median female worker.[7]

Importantly, we want our instrument to vary only with changes in Social Security benefit rules and do not want to capture changes in earnings profiles due to human-capital and productivity changes in cohorts over time. Therefore, we use the earnings history that we constructed for the 1916 cohort for *all* birth cohorts and simply use the CPI to adjust this profile for inflation for earlier and later cohorts. Thus, all birth cohorts have the same real earnings trajectory over time. By holding lifetime earnings constant by construction, this ensures that all of the variation in the instrument comes from variation in the benefit formula due to the law change.

Our next step is to input the constructed earnings histories into the Social Security Administration's ANYPIA program. This program calculates the monthly benefit (PIA) at retirement, given a date of birth, date of retirement, and earnings history. We assign birthdays of June 2 in the particular year of birth and assume that people retire and claim benefits in June of each possible year they could retire. That is, we do not incorporate any variation across cohorts in average *actual* retirement ages (which might be correlated with tastes for independent living). Instead, for each year of birth, we calculate benefits for each *possible* retirement age and then weight the retirement-age-specific benefits by the distribution of claiming ages from the 1985 *Annual Statistical Supplement* to yield a retirement-age-weighted PIA for that year of birth. The same claiming-age weights are applied to each year of birth. Married couples are assigned 150 percent of this PIA.

6. The law also introduced real-wage indexation, changed the lifetime-earnings' concept to AIME, and had other minor changes that affected benefits calculation, primarily minimum earnings for a quarter of coverage and changes in the earnings test, which do not apply in this context (Kollman 1996).

7. In fact, in separate tabulations in the CPS, the median earnings of never-married females are significantly more highly correlated with male earnings than with the earnings of all females.

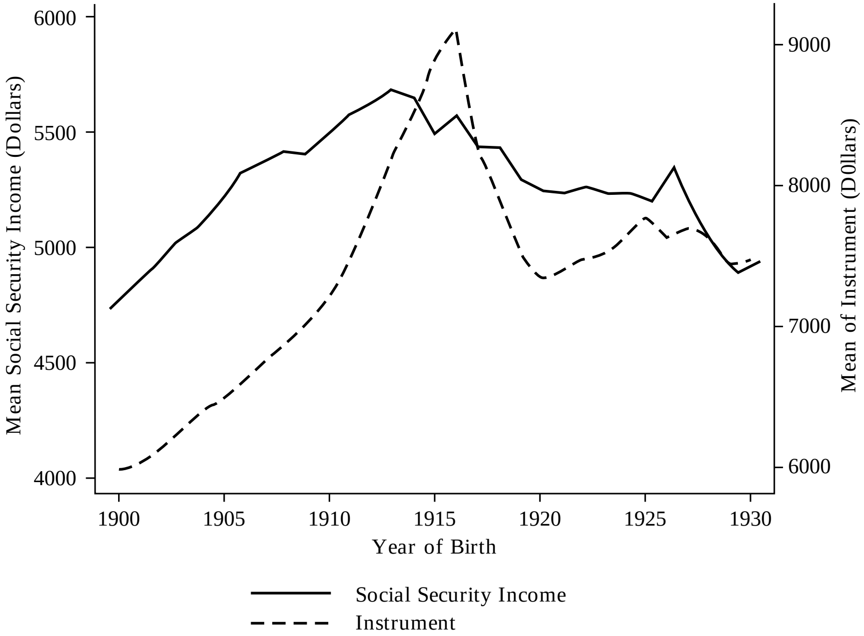

Figure 1
Social Security Income and the Instrument by Year of Birth, 1900–30

The Social Security Administration periodically increases nominal benefits to adjust for inflation. To obtain a value for the predicted benefit for a given age and year-of-birth cohort, we need to account for all "cost of living adjustments" (COLAs) until the date of interview. We calculate the median month in which a given age and year-of-birth cell was interviewed and administer all COLA adjustments from the time that the person would have retired through this date.[8] This produces a predicted, COLA-adjusted Social Security monthly benefit for each age and year-of-birth cell. We then multiply by 12 to get the predicted annual benefit.

Figure 1 shows the plot of cell mean annual Social Security income versus the instrument by year of birth.[9] Unless otherwise indicated, we use years of birth from 1900–33 in the empirical analysis. The variation in benefits, even conditional on constant earnings histories, is readily apparent in the graph of the instrument: benefits rise steadily until the 1910, ramp up quickly for the 1910–16, fall precipitously for the

8. We assume that they claim in June because some cost-of-living (COLA) adjustments were administered in June of a given year, rather than December of a given year. We assume that the beneficiary claims in June so that he will receive any COLA in that year. This prevents variation across years of birth based simply on the timing of the COLA.
9. Although our data run through year of birth of 1933, we censor these figures at 1930 because small sample sizes in the last few years lead to highly variable patterns in the data.

1917–21 cohorts, respectively, and then rise more slowly for cohorts thereafter. The graph of actual Social Security incomes by cohort tracks this pattern fairly well, with the benefits notch apparent in the data. So there is a good first-stage relationship here: our legislative-variation instrument clearly predicts actual Social Security incomes. This instrument uses only variation in benefits by retirement age and year of birth.

The relationship between this instrument and the share of elderly living with others is shown in Figure 2. There is a negative correspondence between these two series: when legislative generosity rises in the early part of the sample, the share living with others falls, then both reverse at a similar time, and flatten out in the later years. The correlation between these series is –0.18.

C. Regression Specification

To examine the effect of Social Security on living arrangements, we estimate the following basic specification,

(1) $P_{ij} = \delta'\mathbf{X}_{ij} + \theta SSIncome_{ij} + u_{ij}$,

where i and j index year of birth and age, respectively. P is the proportion of families in a shared living arrangement, *SSIncome* is the cell mean reported annual Social

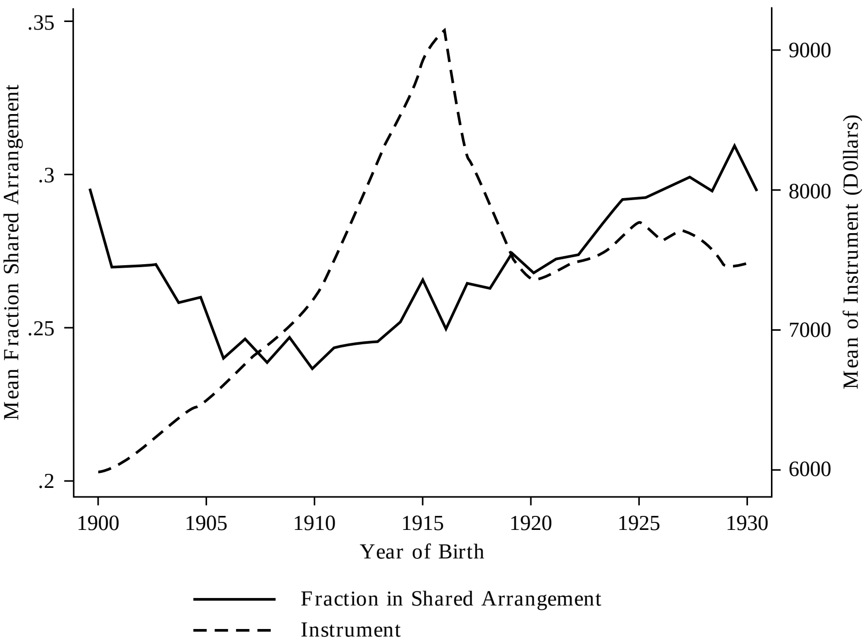

Figure 2
Fraction of Elderly in a Shared Living Arrangement and the Instrument by Year of Birth, 1900–30

Security income, and u is a disturbance term.[10] The parameter θ indicates the change in the proportion of elderly in shared living arrangements for a change in Social Security income. **X** is a vector of all other explanatory variables. We specify $\delta'\mathbf{X}$ as

$$(2) \quad \delta'\mathbf{X}_{ij} = \beta' x_{ij} + \sum_{j=65}^{90} \gamma_j D_{ij}^{Age\,j} + \sum_{t=1980}^{1998} \alpha_t D_{ij}^{Year\,t} + \sum_{r=1}^{8} \phi_r D_{ij}^{Region\,r},$$

where x is a vector of demographic variables that includes controls for cell means of educational attainment (high school diploma, some college, and college or advanced degree) of the head and spouse (if present), age of the spouse (if present), marital status (married, widowed, and divorced in the pooled sample) white, and female. By controlling for these cell characteristics, we control for any other trends in cohort characteristics that might be correlated with both the legislative changes in benefits determination and living arrangements. Following Krueger and Pischke (1992), we also include in Equation 2 a full set of dummies for the age of the head, $D^{Age\,j}$, calendar-year dummies, $D^{Year\,t}$, and Census region of residence share in each cell, $D^{Region\,r}$.[11] The age dummies control for differences across age groups in their propensity to live alone; the year dummies control for any general time trends in living arrangements. Thus, after controlling for age and calendar year, the variation in *SSIncome* is based only upon year of birth. When we then instrument with the variable described above, which we denote as Z, our model is identified solely by legislative variation in benefits generosity across birth cohorts, and not any differences in their earnings history. The means of the dependent variable and primary explanatory variables are shown in Table 1 for each sample.[12]

IV. Results

Panel A of Table 2 gives the grouped OLS estimate of θ for samples based on marital status. All models use cell sizes as sample weights. Standard errors are shown in parentheses. All coefficients are multiplied by 1,000 for ease of interpretation; so, the coefficient shows the impact of a real \$1,000 rise in annual Social Security benefits on living arrangements.[13]

For the pooled sample in Column 1, the OLS estimate is –0.0088, and it is marginally statistically significant. This says that for each \$1,000 of annual Social Security income, the likelihood that the typical elderly person lives with others falls by 0.9 percentage points. Thus, across all elderly, privacy is clearly a normal good. The implied elasticity of living with others with respect to Social Security income

10. We use Social Security income as the regressor of interest for consistency with the previous literature. If we use total family income instead, the coefficients are almost identical, as there is little crowd-out of other income when Social Security income is moved by our instrument. The elasticities are larger, however, because Social Security income is only a fraction of total income.
11. The excluded group consists of families with heads' age over 90, observed in calendar year 1999, residing in the ninth (Pacific) Census region.
12. Descriptive statistics for all variables and samples are available in an appendix from the authors.
13. Based on the standard deviations of annual Social Security income by marital status in Table 1, a \$1,000 increase in annual benefits represents between a 0.8 (for married) and 1.5 (for never married) standard deviation change in annual income.

Table 2

Parameter Estimates of the Effect of Social Security on the Proportion in a Shared Living Arrangement, Elasticity in Brackets, for Years of Birth from 1900–33, Using the 1980–99 March CPS (Standard Errors in Parentheses)

	(1)	(2)	(3) Sample	(4)	(5)
Explanatory Variable	Pooled	Widowed	Divorced	Married	Never Married
A. OLS Estimates					
Social Security income	–0.0088	–0.0366	–0.0290	0.0006	–0.0226
	(0.0051)	(0.0078)	(0.0091)	(0.0036)	(0.0095)
	{–0.18}	{–0.55}	{–0.38}	{0.03}	{–0.21}
B. First-stage estimates					
Instrument	0.1862	0.1017	0.1073	0.3757	0.1729
	(0.0135)	(0.0121)	(0.0271)	(0.0261)	(0.0322)
C. Instrumental variable estimates					
Social Security income	–0.0205	–0.0863	–0.1080	–0.0030	–0.0486
	(0.0092)	(0.0218)	(0.0528)	(0.0062)	(0.0375)
	{–0.41}	{–1.30}	{–1.42}	{–0.14}	{–0.44}
Proportion of heads who are White	–0.0958	–0.0405	0.0164	–0.1750	–0.1460
	(0.0800)	(0.0815)	(0.0717)	(0.0648)	(0.0905)
Proportion of heads with high school diploma	–0.0345	–0.0167	0.1449	–0.0090	–0.0578
	(0.0419)	(0.0559)	(0.0604)	(0.0460)	(0.0710)
Proportion of heads with some college	–0.0657	0.0014	0.0089	0.0643	–0.0243
	(0.0580)	(0.0677)	(0.0819)	(0.0550)	(0.1000)
Proportion of heads with college degree	0.0389	0.1579	–0.0050	0.0850	–0.1582
	(0.0667)	(0.0886)	(0.0728)	(0.0514)	(0.0770)
Proportion married	–0.2016	—	—	—	—
	(0.1052)				
Proportion widowed	–0.0827	—	—	—	—
	(0.0792)				
Proportion divorced	–0.0655	—	—	—	—
	(0.0963)				
Number of observations	494	494	494	473	473

Note: The dependent variable is the proportion of households in the cell in a shared living arrangement. The table shows the parameter estimate of the effect of Social Security income on the proportion in a shared living arrangement. All models use weighted-instrumental-variable estimation with cell sizes as sample weights. Standard errors are in parentheses. Income is measured in thousands of 1982–84 dollars. The specifications also include controls for dummy variables for single years of age from 65 to 90 for the head, calendar years 1980–98, cell share for the age and educational attainment of the spouse (if present), cell share in eight Census regions, and cell share female. The elasticity of the proportion in shared living arrangements with respect to Social Security income is shown in curly brackets. It was calculated based on the parameter estimate shown in the table and sample means of the dependent variable and Social Security income shown in Table 1.

is –0.18. This is at the lower end of the previous literature, but that is not really a sensible comparison, as we are pooling all elderly and not examining singles only.

Unfortunately, the OLS estimates might be biased and inconsistent due to endogenous and omitted variables, as outlined in Section II. Panels B and C of Table 2 show the grouped first-stage and instrumental variable (IV) estimates, respectively. Overall, there is a very good first-stage fit. For the pooled sample, the IV coefficient rises to –0.02, and the elasticity more than doubles to –0.41. This is a sizeable effect for the entire pool of elderly.

The pooled sample combines households of different marital types, some of which might be expected to display quite different responsiveness of Social Security to living arrangements. For example, because most married couples live independently (of other adults) and have many potential sources of income with which to support themselves, they may be expected to have relatively low sensitivity of shared living arrangements to Social Security a priori. On the other hand, widowed individuals may be heavily reliant on Social Security as an income source, and, therefore, be expected to have a much more elastic response. In addition, never-married elderly are less likely to have had children, and therefore have fewer options for shared living arrangements.[14] Therefore, Columns 2–5 in Table 2 show estimation results for four different subsamples split out by marital status.

The first subsample is the sample of most interest from the previous literature, widows. Our OLS estimate for this population is that each $1,000 in annual benefits leads to 3.66 percentage points fewer widows living in a shared arrangement. The implied elasticity is –0.55, which is in the center of the previous literature. When we instrument, however, the effect more than doubles, so that each $1,000 in Social Security income leads to 8.63 percentage points fewer widows living in a shared arrangement, for an implied elasticity of living with others of −1.3. This is well above even the largest estimates from the previous literature and suggests that identification problems or different timing have biased downward estimates of the responsiveness of widows to income in their residential decisions.[15]

The remaining columns of the table show the results for the other marital categories. In Column 3, we find that divorcees are roughly as income sensitive as widows, with an instrumental variable elasticity of −1.42, significantly larger than those of Costa (1999).[16] The estimates for married households in Column 4 indicate that the

14. One important omitted determinant was the availability of kin, and children in particular, with which to share living arrangements (Macunovich et al. 1995; Wolf 1994). We cannot control directly for this because the March CPS did not gather data on fertility histories, but in Engelhardt, Gruber, and Perry (2002) we used the fertility histories for women from the same cohorts from the June 1980 and 1985 CPS, regressed the number of children on the instrument and other explanatory variables, and found no evidence that benefits were correlated with the number of children ever had.

15. A widow whose husband dies at a relatively young age will receive less than a widow whose spouse dies at an older age, due to a longer earnings history for the deceased spouse. To address this, we constructed an alternative instrument: a weighted-average expected Social Security benefit were the spouse to stop working and die at each possible age from 40–65 based on mortality probabilities for the 1916 cohort. We then reestimated the models, and the results did not change.

16. Some divorced women may have reported they were widowed upon death of the ex-spouse, generating some classification error in reported marital status between widowed and divorced. As a robustness check, we pooled widowed and divorced together as a group and reestimated the specification; our results did not change.

effect of additional Social Security income on the proportion in shared arrangements is small and not statistically different than zero. The final column examines the impact on never-married individuals. Here there is a sizeable negative effect, but it is not statistically significant. It implies that each $1,000 of Social Security income lowers the likelihood of living with others by 4.86 percentage points, for an implied elasticity of –0.44. These elasticities are similar to the IV estimates in Costa (1999).[17]

Social Security is a more important source of retirement income for low-relative to high-income elderly, and, hence, changes in benefits might have had a larger impact on living arrangements for low-income elderly. To examine this, rather than focus on elderly income, which is potentially endogenous because of crowd out, we split our age-by-year-of-birth cells by educational attainment group: high school diploma or less and greater than high school diploma.[18] Then we remade our instrument. In doing so, we imposed within a single year of birth the same real earnings trajectory described above but allowed the *level* of earnings to vary by education group according to a scaling factor: the quotient of education-group earnings to median earnings for workers born in 1916 in the March 1962 CPS. The education-group earnings-scaling factor was applied to the earnings profile for all years of birth. This implies, for example, that, across years of birth, all individuals with a high school diploma or less had the same real earnings histories for the purposes of calculating our instrument and similarly for the other group.

Figure 3 shows the plots of the instrument by year of birth for the two education groups. The variation in benefits within a year of birth across education groups is readily apparent. In particular, benefits fall precipitously in the 1917–21 period for both groups, but for those with greater than a high school education almost had returned to the prenotch levels by the time the 1930 cohort reached retirement, whereas benefits for those with less than a high school degree never really recovered and remained just above the post-notch level. As described above, this occurs because the 1977 law raised the covered-earnings maximum used to calculate AIME. This meant that after 1977 a greater fraction of annual earnings for high-wage (that is, high-education) workers entered AIME, whereas essentially all of the earnings of low-wage workers were below the maximum even before the law. This was magnified with time as high-education individuals in younger birth cohorts had a greater proportion of their lifetime earnings exposed to the higher covered-earnings maxima.

Table 3 shows the IV parameter estimates of the effect of Social Security on living arrangements by education group and marital status. The parameters in Equation 1 were estimated separately for these two groups. The estimates indicate that the impact of Social Security on living arrangements is economically and statistically more important for elderly with a high school education or less both for the pooled sample in Column 1 and in Columns 2–5 for the four different subsamples, split out by marital status. Social Security generally has an economically smaller and imprecisely

17. Table 2 also shows estimates for a selected group of demographic variables from the IV specifications. The complete set of parameter estimates is available from the authors. There is a large literature in economics, demography, and sociology that documents an inverse correlation between income and mortality. Unfortunately, we have no direct way to address the impact of differential mortality. However, Snyder and Evans (2002) examined the effect of income on elderly mortality using the Social Security notch as a source of variation in income and found that the notch led to longer, rather than shorter, lives.

18. Cell sizes were too small to break the sample into finer educational-attainment groups.

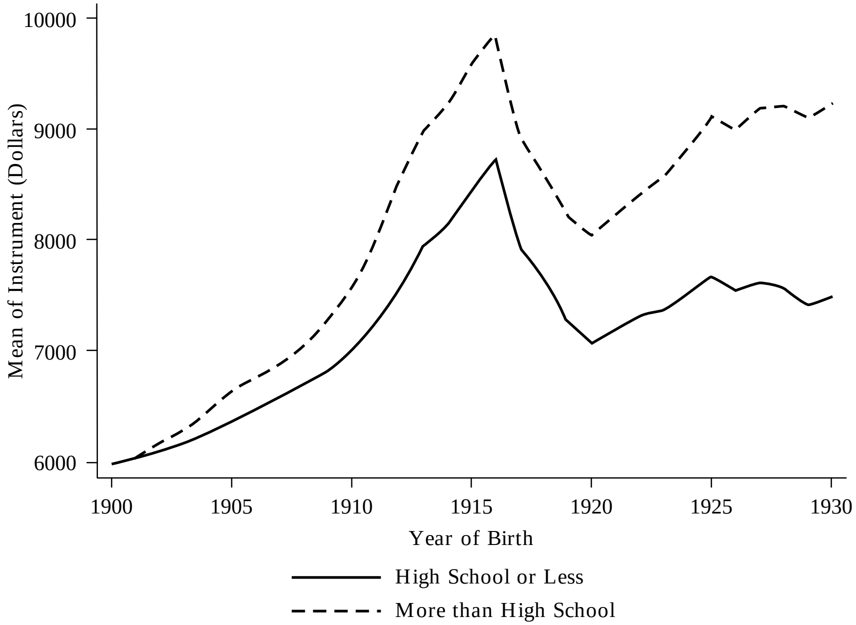

Figure 3
Instrument by Year of Birth and Education Group, 1900–30

estimated impact for elderly with greater than a high school education, although given the estimation precision, the results across the education groups are not statistically distinguishable from each other.

V. Implications for Benefit Reduction

To give some sense of the implied policy effects from our estimates, Table 4 shows the impact of a 10 percent cut in benefits on living arrangements, all else held equal, in 1999, the last year in our sample. The first column shows the number of households 65 and older who currently live with others by marital status. In 1999, over 5.5 million elderly households lived in shared arrangements, and almost two-thirds of these households (about 3.8 million) were either divorced or widowed individuals. The second column shows the mean annual Social Security benefit in each group. The third column shows the additional number of elderly who would live in shared arrangements if benefits were cut by 10 percent. The results are striking. In total, more than 600,000 elderly households would move into a shared arrangement if benefits were cut, and more than 430,000 widows would do so. Overall, almost all the elderly affected would be either widowed or divorced individuals. This is because those groups comprise the great majority of the elderly (as shown in Column 1) and

Table 3
Instrumental Variable Parameter Estimates of the Effect of Social Security on the Proportion in a Shared Living Arrangement by Education Group, Elasticity in Brackets, for Years of Birth from 1900–33, Using the 1980–99 March CPS (Standard Errors in Parentheses)

	(1)	(2)	(3) Sample	(4)	(5)
Education Group	Pooled	Widowed	Divorced	Married	Never Married
High school or less	–0.0408	–0.1184	–0.0895	–0.0041	–0.0443
	(0.0090)	(0.0191)	(0.0325)	(0.0081)	(0.0186)
	{–0.76}	{–1.71}	{–1.12}	{–0.18}	{–0.37}
	[494]	[494]	[483]	[464]	[470]
More than high school	–0.0131	–0.0181	–0.0388	0.0018	–0.1096
	(0.0116)	(0.0339)	(0.3059)	(0.0097)	(0.1153)
	{–0.37}	{–0.41}	{–0.68}	{0.12}	{–1.36}
	[494]	[494]	[427]	[451]	[454]

Note: The dependent variable is the proportion of households in the cell in a shared living arrangement. The table shows the instrumental variable parameter estimate of the effect of Social Security income on the proportion in a shared living arrangement. All models use weighted-instrumental-variable estimation with cell sizes as sample weights. Standard errors are in parentheses. Income is measured in thousands of 1982–84 dollars. The specifications also include controls for dummy variables for single years of age from 65 to 90 for the head, calendar years 1980–98, cell share for the age and educational attainment of the spouse (if present), cell share in eight Census regions, and cell share female. The estimates in this table differ from those in Table 2 in that the specifications were estimated separately by education group, so that all of the model parameters were allowed to vary by education group. The elasticity of the proportion in shared living arrangements with respect to Social Security income is shown in curly brackets. It was calculated based on the parameter estimate shown in the table and the education-group means of the dependent variable and Social Security income. The sample size is shown in square brackets.

Table 4
Estimated Effect of a 10 percent Social Security Benefit Cut on the Number of Households 65 and Older in Shared Living Arrangements in 1999

	(1)	(2)	(3)
Marital Status	Number of Households in Shared Arrangements	Mean Social Security Benefit	Additional Households in Shared Arrangement Due to 10% Benefit Cut
Married	1,324,008	$8,843	30,084
Never married	416,304	$4,242	25,491
Divorced	743,562	$4,302	118,556
Widowed	3,042,040	$4,885	434,225
Total	5,525,914	—	608,356

Note: Authors' calculations using the sample mean proportion in shared arrangements, Social Security income, and cell sizes in the four samples (married, never married, divorced, and widowed) in 1999, the weighted-instrumental-variable parameter estimates for Social Security income in Columns 2–5 from Panel C of Table 2, and the CPS population weights for 1999.

had the most elastic response of living arrangements to Social Security (in Panel C of Table 2). We emphasize that these simulations assume that the responsiveness of current and prospective elderly to benefit cuts is the same as those elderly affected by the notch, so that these implications are to some degree speculative. In particular, the elderly today are increasingly living longer and are in better health, and the size of the elderly cohort is rising rapidly relative to the size of the nonelderly relatives with whom they may live, both of which may lead the elderly to crave more independent living.

VI. Conclusion

Our findings raise important questions about the welfare implications of shifting living arrangements. The fact that living arrangements are so income sensitive, particularly for widows and divorcees, implies that privacy is a valued good. If there is rational, forward-looking decision making by the elderly and their families/others who share their households, and if utility is jointly maximized over the household unit, then this implies that welfare is reduced (along this dimension) when benefits are cut and the elderly are forced to live with others.

However, these assumptions may not hold in reality. For example, the elderly may crave independence in the short run, but underestimate the long-run costs of living alone, either due to information failures or to time inconsistency in discounting the future. There have been numerous studies in the demography, medical, and gerontology literatures that suggest there are significant costs and risks to living alone for the elderly. One pathway is through physical and health risks; another is through attenuated social interaction. Whether the elderly fully anticipate these costs and risks and rationally trade them off for the benefits of independence is unclear. Even if the elderly make their decisions rationally, if it is the others in the household who control the decision on living arrangements, and if they are not jointly maximizing the well-being of the elderly and themselves, then there may be suboptimal allocation of living arrangements. For example, if children want to "get rid" of their parents, so long as the parents have some minimum level of income on which to live independently, then rising benefits could lead to more independence but lower welfare. Thus, the welfare implications of these findings are unclear. Exploring these dimensions is well beyond the scope of this paper, but clearly worthy of future research.

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