The Impact of Employment on Parental Co-Residence

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Abstract

Parents may use housing and shared living arrangements as a form of risk-sharing in the face of labor-market uncertainty for their adult children. We test this by estimating the effect of employment on co-residence between parents and adult children using detailed data on children and parents in the Health and Retirement Study (HRS) for 1998-2012. Our estimates indicate that on average a young man moving from full-time to non-employment raises the likelihood of co-residing with a parent by 1.7 percentage points. Similarly, on average moving from full-time employment to being part-time employed raises the likelihood of co-residing with a parent by 2.1 percentage points. The implied elasticity of parental co-residence with respect to the son's income is -1.3. We find more muted responses for daughters: the implied elasticity of parental co-residence with respect to the daughter's income is -0.3. For sons, the impact of employment changes on co-residence depends on the quality and quantity of housing services supplied by the parent: co-residence is substantially less likely with parents who rent, have smaller, and low-quality residences.

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1. Introduction

There has been recent interest in the extent to which parents use housing and shared living arrangements as a form of risk-sharing in the face of labor-market uncertainty for their adult children (Kaplan, 2009, 2012; Wiemers, 2014; Lee and Painter, 2013). In particular, co-residence may be a form of non-employment insurance that allows children to smooth non-housing consumption over the business cycle. This may result in household formation, a key determinant of homeownership and housing demand, being more elastic in business cycle troughs than peaks. In this paper, we present estimates of the effect of transitions away from full-time employment on co-residence between parents and adult children.

Our analysis has a number of key features. First, we focus on so-called "boomerang kids." These are adult children who are living independently, but then may choose to return home and co-reside with a parent. To isolate the impact of transitions away from full-time employment, we limit our analysis to adult children who are initially observed to be full-time employed. Second, we use detailed data from 1998 through 2012 on children of respondents in the Health and Retirement Study (HRS) to construct a panel dataset of young adults and their parents. Existing studies have focused on men between the ages of 24 and 44, who often experience the most difficulty in adjusting to unemployment and whose changes in employment status are more plausibly attributable to adverse changes in labor demand, rather than labor supply. Women have received far less attention. We estimate impacts for both sexes. Third, a key empirical challenge is that there may be unobserved heterogeneity in the taste for leisure of adult children that may be correlated with the demand for shared living arrangements with parents. If young adults with a high taste for leisure also prefer to live with their parents, then standard estimates of the effect of employment loss on parental co-residence would be biased upward. Similarly, parental demand for co-residence due to functional limitations associated with declining health may also bias estimates. The rich array of information collected on parents' own health, functional status, marital status, income, employment, housing, and wealth trajectories, as well as the panel nature of the HRS, allows us to control for a large number of independent influences on residential decisions and employ a variety of fixed-effect estimation techniques.

Our best estimates indicate that, on average, a young man moving from full-time employment to being not employed (i.e., either unemployed or out of the labor force) raises the likelihood of co-residing with a parent by 1.7 percentage points. Similarly, on average moving from full-time employment to being part-time employed raises the likelihood of co-residing with a parent by 2.1 percentage points. The implied elasticity of parental co-residence with respect to the son's income is -1.3. These effects are economically large, statistically significant, and suggest that the ability to move in with parents is an important way that sons adjust during economic downturns. We find more muted responses for daughters: the implied elasticity of parental co-residence with respect to the daughter's income is -0.3.

Unlike earlier studies, we are able to exploit the tremendous detail on the housing and health status of parents in the HRS to test two key hypotheses about the role of parental characteristics in the co-residence decision. In principle, for a loss of full-time employment, the likelihood a child co-resides should increase with the quality and quantity of housing services supplied by the parent. We find evidence strongly consistent with this for sons: co-residence is substantially less likely with parents who rent, have smaller, and low-quality residences. We, however, find little consistent evidence of daughters' co-residence decisions based on the quality and quantity of housing supplied by their parents. Also, intergenerational living simultaneously may provide both non-employment insurance to the adult child and informal care insurance to the parent(s). This implies that, for a loss of full-time employment, the likelihood a child co-resides should increase as the health of the parent declines. Using a wide variety of measures of health and functional status, we find little robust evidence for this hypothesis for either sons or daughters.

The remainder of the paper is organized as follows. Section 2 briefly reviews the relevant literatures on the economics and demography of housing and living arrangements. Section 3 describes the HRS data and sample construction. Section 4 lays out the econometric framework and presents the basic estimation results on co-residence. Sections 5 and 6 discuss robustness checks and extensions, respectively. The paper concludes with a brief summary of findings and a discussion of caveats.

2. Related Studies

Our analysis is most closely related to five interconnected strands in the existing literature in economics and demography. The first is centered in urban and real estate economics, and has focused on the effect of household formation on homeownership rates (Dietz and Haurin, 2003). The homeownership rate is measured as the number of owner-occupier households divided by the total number of households. Economic conditions can affect both the number of owners versus renters via tenure choice (the numerator), as well as the number of households via household formation (the denominator). In an early, influential study, Börsch-Supan (1986) used data from the 1976-1977 waves of the American Housing Survey (AHS) and showed empirically that household formation was quite responsive to house prices and income. While this is a large literature, other notable work in this area includes Haurin, Hendershott, and Kim (1993), Yelowitz (2007), Haurin and Rosenthal (2008), and Lee and Painter (2013).

A second strand, located in labor economics, has treated the work and household formation decisions of young adults as jointly determined. McElroy (1985) is the best-known study in this area.[1] In particular, she specified a structural model of labor supply and household formation, estimated using data on 203 never-married white men with completed schooling between the ages of 19 and 24 from the 1971 National Longitudinal Survey of Young Men (NLS-YM) matched to data on their parents from the National Longitudinal Surveys of Mature Women (NLS-MW) and Men (NLS-MM), respectively. There are two important implications of her analysis. First, there is an ordered relationship between wage offers, employment, and residential independence. For wage offers below the adult child's reservation wage, the child does not work but co-resides with the parents; for offers above the reservation wage, the child works but still co-resides; for offers sufficiently above the reservation wage, the child works and lives independently. Second, the option of an adult child to co-reside with parents is a form of non-employment insurance.

McElroy's analysis abstracts from other forms of intergenerational support to young adults. A third strand of existing studies has treated co-residence as a type of intergenerational transfer that may be a substitute for (or complement to) financial transfers. This is a voluminous literature, whose focus is often on pinning down the motives for transfers. It spans both economics and demography, and is reviewed in detail in Bianchi et al. (2006). Ermisch (1999) is the best known paper in urban economics in this vein. He treats income as exogenous, but considers financial transfers in addition to co-residence. Rosenzweig and Wolpin (1993) treat human capital investments (and hence lifetime income), co-residence, and financial transfers as jointly determined. They estimate the parameters of a structural model based on a panel of 821 matched parent-son pairs for young men from the 1967 NLS-YM matched to parents in the NLS-MW and

[1] Haurin, Hendershott, and Kim (1993) also treat work decisions and income as endogenous.

NLS-MM, and followed up to eleven waves between 1967 and 1981. They found that an additional week of unemployment of the son increased the likelihood of parental co-residence by 3 percentage points, based on fixed effects logit estimation.

In economics, much of the existing work has examined the role of economic conditions on initial transitions out of the parental home and new household formation. In contrast, much effort has been spent in the demography literature in making distinctions between different types of transitions. This includes those who initially transition out of the parental home, as well as their complement, i.e. those who "fail to launch" (Goldsheider and DaVanzo, 1985; Billari and Liefbroer, 2007; Bianchi et al., 2006; among many others). More recently, researchers have begun to examine the link between the economic status and residential patterns of young and middle-aged adults after they have had a spell living independently, with those who return to the parental home termed "boomerang kids."

Kaplan (2012) embedded the option of the adult child to move in and out of the parental home into an intergenerational dynamic, stochastic lifecycle framework. Building on McElroy (1985), co-residence in his model is a form of non-employment insurance that allows adult children potentially to smooth non-housing consumption in the face of labor-market uncertainty. He estimated the structural model for 1,491 young men who did not attend college using monthly data from 1998 through 2002 on employment and co-residence drawn from the National Longitudinal Survey of Youth, 1997 Cohort (NLSY97) and found that labor market shocks affect the timing of moves into and out of the parental home. This type of insurance is particularly valuable for individuals with parents in the lower part of the income distribution, who have comparatively lower ability to supply financial transfers to the adult child in the face of an adverse labor-market shock. The option of moving back home also is associated with higher future earnings by allowing for more extensive search for jobs with higher earnings potential.

Although Kaplan's period of study was 1998-2002, his work was particularly well-timed in that it appeared just when the United States was in the midst of the largest macroeconomic contraction since the Great Depression. This period beginning with the financial crisis in 2007, often termed the Great Recession, saw a historically high level of co-residence among American individuals (Mykyta and Macartney, 2012; Mykyta, 2012). This has led to a number of very recent studies of the impact of the recession on living arrangements, all of which are closely related to our analysis below.

Lee and Painter (2013) examined the impact of recessions on household formation. Specifically, they used data on young men and women from the Panel Study of Income Dynamics (PSID) from 1975-2009, which spans four recessions as dated by the National Bureau of Economic Research (NBER), including the Great Recession. They found that being unemployed has a substantial negative impact on the transition out of the parental home for young adults. In particular, treating employment status as exogenous, their multinomial logit estimates indicate that unemployment lowers the likelihood of a transition by 50%. In addition, they found that, independent of the individual's unemployment status, more general macroeconomic and housing-market conditions matter for household formation, including the state unemployment rate, average wage, GDP growth, median gross rent, median house value, respectively, as well as whether the national economy is in a recession year. They also found much stronger effects for the Great Recession than earlier economic downturns.[2]

[2] Wiemers and Bianchi (2014) found substantial differences in co-residence among women by birth cohort, primarily driven by changes in life expectancy across cohorts.

Rogers and Winkler (2014) used American Community Survey (ACS) data from 2005 and 2011 and examined how the differential timing of the downturn in both labor and housing markets across metropolitan areas in the Great Recession affected household formation by young adults. Although their focus was on the impact of broader market conditions, treating employment status as exogenous, their logit estimates indicated that not being employed reduced the likelihood of living independently by 11%.

In a closely related study, Wiemers (2014) examined the impact of unemployment on household composition and doubling up. She used data from the 1996, 2001, 2004, and 2008 Survey of Income and Program Participation (SIPP) panels, which span the period from December, 1998, through December, 2010. With repeated observations on individuals, she estimated individual-specific fixed effects linear probability models and found that becoming unemployed raises the likelihood of living in a shared arrangement by 0.3%. While statistically significant, this is an economically small effect.

3. Data and Descriptive Statistics

The empirical analysis in this paper focuses on boomerang kids. Specifically, for children who were observed living independently from their parents and working full-time, we examine the impact of reductions in employment on subsequent co-residence with parents.

We construct our analysis dataset from the HRS, which is a stratified random sample of over 25,000 individuals 50 and older, and their spouses (regardless of age). The study began in 1992, and respondents are interviewed in waves that occur every even-numbered calendar year. Every six years (e.g., 1998, 2004, 2010, 2016, etc.), a new birth cohort of individuals in their mid-50s enters the study, refreshing the panel. In our application, the HRS respondents are the parents.

Every wave, we obtain a rich set of information on the demographic, economic, and health characteristics of the parents from the core survey. In addition, every wave the parents report on the economic, locational, demographic, and co-residence status of their children, the great majority of whom are adults.

We use the 1998-2012 HRS waves on parents paired with their reports on their children to form a matched panel dataset on parents and their adult children. The 1998 wave is the earliest one for which our analysis variables are measured consistently. To focus on boomerang kids, we restrict ourselves to children who satisfied the following criteria: they were between the ages of 25 and 44, were employed full-time, and lived independently from their parents at "baseline," which is defined as the year when their parents entered the HRS.[3] We exclude children who were ever reported being students during the sample period. Once the child returns to co-reside with a parent, that child is censored in subsequent sample years.[4]

The sequence of labor-market questions in the HRS on each child is as follows. Initially, the parent was asked whether the child worked 30 hours or more per week, less than 30 hours, or not all. We define the child as "full-time employed" (*F*) if the parent reported the child to be working 30 hours or more per week; the child as "part-time employed" (*P*) if the parent reported the child to be employed less than 30 hours a week; and the child as "not employed" (*U*) if the parent reported the child is not employed. Importantly, "not employed" could be either unemployed (and actively seeking work) or out of the labor force; the HRS did not distinguish between the two. Then, the parent was asked to indicate the child's total family income in ranges:

[3] For members of the original HRS (b. 1931-41), CODA (b. 1924-1930), and War Babies (b. 1942-46) cohorts, the baseline year is 1998; for the Early Baby Boomer cohort (b. 1947-1953), the baseline year is 2004.

[4] That is, we do not allow for repeated departures and returns, whereby a child starts off living independently, then "boomerangs" back to the parental residence, then departs to a second spell of independent living, then boomerangs a second time. In practice, there are very few of these cases. When we expand the sample to include these censored observations, the point estimates and standard errors shown in the tables below do not change appreciably.

less than $10,000, $10,000-$35,000, $35,000-$70,000, $70,000-$100,000, and over $100,000. While most theories of intergenerational transfers, such as those motivated by altruism or exchange, rely on the relative income of children and parents as one of the key determinants of behavior, we believe that the child employment measures in the HRS are substantially more accurate than the income measures. Specifically, we think it is much more likely that the parent knows the employment status, as opposed to the income, of the child. This is reflected in the data: about 20% of parents who report employment status for a child do not know the income of the child. In addition, as described above, most related studies have concentrated on the role of employment. For these reasons, we focus on employment transitions in the econometric analysis. For robustness, we also show selected results using income.

Columns 1-4 of Panel A of Table 1 present basic summary statistics on sons in the sample, which have been the focus of the bulk of the previous literature in economics (McElroy, 1985; Rosenzweig and Wolpin, 1993; Kaplan, 2009, 2012, among others). There are 43,097 person-year observations on 9,279 sons. Although, by construction, all sons start the sample full-time employed, almost 10% experience an employment change in the sample period that leaves them part-time or not employed. Average family income, defined for 7,460 sons, was about $57,000, but declines substantially for those part-time and not employed. Panel B shows summary statistics for the outcome variable: the fraction of men who co-reside with parent. Although, by construction, all sons start the sample living independently, 8% of those who transition to part-time employment co-reside. Just over 6% of those not employed co-reside. So, for sons there appears to be a positive correlation in the raw data between declines in employment and parental co-residence.

Columns 5-8 show summary statistics on daughters. There are 34,715 person-year observations on 7,345 daughters. Although, by construction, all daughters start the sample full-time employed, 20% experience an employment change in the sample period that leaves them part-time or not employed. Average family income, defined for 5,870 daughters, was about $56,000. However, unlike for sons, family income does not fall as labor force attachment falls for daughters, perhaps suggesting they are more likely to be the secondary earners in the household. While all daughters start the sample living independently, 2% of those who transition to part-time employment co-reside. Almost 3% of those not employed co-reside. So, there appears to be a much weaker correlation between employment and parental co-residence for daughters than for sons.

4. Econometric Framework and Estimation Results

To move beyond the simple comparisons in Table 1, we use a regression framework. Let i index the child, j index the family, and t index the calendar year. We follow previous studies and use a linear-in-parameters econometric specification:

$$Y_{ijt} = \xi + \beta U_{it} + \phi P_{it} + \boldsymbol{\delta}\mathbf{X}_{ijt} + \gamma_t + \eta_i + \upsilon_{ijt}. \tag{1.1}$$

The dependent variable, Y, is an indicator that takes on a value of one if the child co-resides at period t with a parent and zero otherwise.[5] Our primary interest is in the impact of the loss of full-time employment on co-residence. As there are two primary alternatives to full-time employment, there are two focal explanatory variables. The first is U, an indicator variable that takes on a value of one if the individual is not employed and zero otherwise. The second is P, an

[5] We interpret this specification as a reduced-form. We note, however, that most of the structural analyses in the literature specify indirect utility from residential choice as linear in parameters, yielding in effect a similar specification as the reduced-form we adopt above.

indicator variable that takes on a value of one if the individual is part-time employed and zero otherwise. Those who are full-time employed (F) are the excluded group in (1.1). Therefore, ξ represents the (conditional) mean fraction of men who are full-time employed and co-residing with a parent. By construction, all children begin the sample living independently, so $Y = 0$, and are full-time employed, so $U = 0$ and $P = 0$. Then a change in employment occurs and co-residence decisions are made. Therefore, β represents the impact on co-residence of a move from full-time to non-employment; ϕ represents the impact on co-residence of a move from full- to part-time employment. The central objective is to obtain consistent estimates of β and ϕ.

This specification includes a full set of calendar-year indicators, γ, and a vector of control variables, $\mathbf{X}$, that consist of the child's age, plus the parental-level characteristics shown in Panel C of Table 1, all of which are time-varying.[6] The controls include parental age, income, wealth, as well as the number of bedrooms in the parents' residence; indicator variables for employment, homeownership, and being married; and three measures of parental health status. The first measure is the number of limits to the parent's Activities of Daily Living (ADLs). The HRS collects information on five activities—bathing, eating, dressing, walking across a room, and getting in and out of bed—each designed to measure various dimensions of an individual's ability to function in his or her residential space. For each of the five tasks, the HRS records a 1 if the respondent had difficulty with that task and a zero otherwise. The scores are summed for the five tasks, so that the ADL variable ranges from 0 (no difficulties with any of the tasks) to 5 (difficulties with

[6] The education of the child is not included as a regressor, because, by construction in the sample, it is time-invariant and swept out by the fixed-effects estimator we use. The child's marital and homeownership statuses are not included as regressors, because they are endogenous outcomes. We examine the impact of employment changes on those outcomes in a separate analysis not discussed here. We control for family attributes of the biological parents separately to allow for parent's no longer residing together. Unfortunately, the HRS did not gather information on the biological parents of the spouses of married adult children of HRS respondents. So, we only have parental information for the parents of young men.

all of the tasks). The second measure is a count of the number of medical conditions a doctor had ever told the parent that he or she had. The eight conditions were high blood pressure, diabetes, cancer, lung disease, heart disease, stroke, psychiatric problems, and arthritis. The index ranges from 0 (the absence of all eight conditions) to 8 (the presence of all eight conditions) where, as before, a larger index value indicates poorer health. The third measure is a count of the number of limits to Instrumental Activities of Daily Living (IADLs), which are the ability to use a telephone, meal preparation, shopping, handle the household finances, and take medication.

One empirical challenge is that there may be unobserved heterogeneity in a child's taste for co-residency with parents, represented by η in (1.1). If labor-market transitions are correlated with this unobserved heterogeneity, then OLS estimates of β and ϕ will be biased and inconsistent. This could occur, for example, if children with a high taste for leisure also have a high taste for co-residency. Then it might appear that those who become unemployed are more likely subsequently to co-reside with a parent, but this would not necessarily be a causal effect.

To account for this, we present fixed-effects linear probability model estimates of β and ϕ. Similar to Wiemers (2014), this allows us, in principle, to obtain consistent estimates of the parameters. With this estimator, the estimates are identified by within-child differences across time in employment trajectories. Our key identifying assumption is that these are (conditionally) uncorrelated with the error term, υ, in (1.1).

Specifically, to ensure consistent estimates, the differences across time in employment trajectories should be due to differential shocks to labor demand between individuals across time, and not to differential shocks to labor supply. Since most estimates in the literature indicate that prime-age male labor supply is quite inelastic, we believe this assumption holds for sons, as changes in employment for men in the 1998-2012 sample period were most likely due to shocks to labor demand across the 2000 recession and the Great Recession. Traditionally, women have had more elastic labor supply, especially on the extensive margin. However, Blau and Kahn (2007) and Kumar (2015) have argued that female labor supply elasticities have converged toward male labor supply elasticities over time with the secular increase in female labor force participation. To reflect any potential differences in responsiveness to employment changes by sex, we present separate estimates for sons and daughters below.

Column 1 of Table 2 shows fixed-effect estimates of the parameters in (1.1) for the full sample of 9,279 sons. The estimate of β is $\hat{\beta} = 0.017$: on average moving from full-time employment to being not employed raises the likelihood of co-residing with a parent by 1.7 percentage points for sons. With a heteroscedasticity-robust standard error (clustered at the individual level) of 0.004, this estimated effect is statistically different from zero at conventional significance levels. Given that the incidence of co-residence among sons working full-time is very low (Table 1), this is an economically sizable impact. Similarly, $\hat{\phi} = 0.021$. That is, on average moving from full-time employment to being part-time employed raises the likelihood of co-residing with a parent by 2.1 percentage points. With a standard error of 0.007, this estimated effect is also statistically different from zero at conventional significance levels.

Column 2 shows the fixed-effect estimates from the alternative specification,

$$Y_{ijt} = \xi + \theta \ln(Income)_{it} + \boldsymbol{\delta}\mathbf{X}_{ijt} + \gamma_t + \eta_i + \upsilon_{ijt}, \tag{1.2}$$

in which the focal explanatory variable is the natural logarithm of the child's household income, and θ represents the semi-elasticity of parental co-residency with respect to income. If independent living is a normal good (holding parental income constant in **X**), then $\theta < 0$. For the subsample of 7,460 sons for whom parents were able to report an income value, $\hat{\theta} = -0.021$, which suggests that if the son's income were to double, the rate of co-residency would fall by 2.1

percentage points. Evaluated at the mean co-residency rate for those working full-time implies an income elasticity of co-residency of -1.3.

Columns 3 and 4 repeat the estimation of (1.1) and (1.2), respectively, for the sample of daughters. As suggested by the comparison of simple means in Table 1, daughters appear less responsive than sons. In column 3, $\hat{\beta} = 0.006$, which implies that on average moving from full-time employment to being not employed raises the likelihood of co-residing with a parent by 0.6, or just over one-half, of a percentage point for daughters. With a standard error of 0.002, this effect is statistically different from zero at conventional significance levels, however it is much smaller than the estimated impact for sons. Indeed, the null hypothesis that the impacts for sons and daughters are equal can be rejected at the 1% level of significance. Similarly, for daughters, $\hat{\phi} = 0.005$. That is, on average moving from full-time employment to being part-time employed raises the likelihood of co-residing with a parent by one-half of a percentage point. With a standard error of 0.003, this effect is not statistically different from zero at conventional significance levels. Furthermore, column 4 shows the estimate of θ in (1.2) for daughters. It is about half the size of the income estimate for sons in column 2. The implied income elasticity of co-residency for daughters is -0.3. Overall, it appears sons are more responsive to changes in labor-market activity in their parental co-residence decisions than daughters.

5. Robustness to Alternative Estimators

The empirical results thus far show evidence that transitions away from full-time employment generate significant increases in co-residence for sons, and less so for daughters. Column 1 of Table 3 shows alternative estimates using the estimator of Arellano and Bond (1991) on the following dynamic specification based on (1.1),

$$Y_{ijt} = \xi + \alpha Y_{ijt-1} + \beta U_{it} + \phi P_{it} + \boldsymbol{\delta}\mathbf{X}_{ijt} + \gamma_t + \eta_i + \upsilon_{ijt}\,, \tag{1.3}$$

in which the lagged co-residence status is allowed to enter as an explanatory variable. Similarly, column 2 shows Arellano-Bond estimates for an isomorphic variant of (1.2),

$$Y_{ijt} = \xi + \alpha Y_{ijt-1} + \theta \ln(Income)_{it} + \boldsymbol{\delta}\mathbf{X}_{ijt} + \gamma_t + \eta_i + \upsilon_{ijt}\,. \tag{1.4}$$

The Arellano-Bond estimates of the impacts of employment and income, respectively, are very similar to those in Table 2.

Another potential challenge is that there might be unobserved factors that vary across families and are correlated with child labor-market dynamics. To represent this, υ in (1.1) can be decomposed as $\upsilon = \mu + \varepsilon$,

$$Y_{ijt} = \alpha + \beta U_{it} + \phi P_{it} + \boldsymbol{\delta}\mathbf{X}_{ijt} + \eta_i + \mu_j + \gamma_t + \varepsilon_{ijt}\,, \tag{1.5}$$

where μ represents time-invariant unobserved heterogeneity at the family level. For example, across extended families, parents may differ in permanent income (Lee and Painter, 2013), which is typically difficult to measure, but might be correlated with child labor-market attachment.[7] Alternatively, across extended families, parents might differ in their provision of informal insurance via co-residence in a way that is correlated with child labor-market attachment. In the presence of such an unobserved family fixed effect, estimates of β and ϕ, would be biased and inconsistent.

To address this, we construct a sibling panel dataset on sons for HRS respondents with more than one son, as the survey gathers child information for all children. With a matched sibling-pair-to-parent dataset, we track residential and employment status of brothers and their parents over time. Hence, we can control both for time-variant unobserved heterogeneity at the family

[7] Or in a lambda-constant dynamic model of private income transfers, the unobserved heterogeneity could represent the parents' marginal utility of (full) income.

level and at the individual level with a two-way fixed-effect estimator: within a son over time, and between sons. With this estimator, controlling for calendar-year effects, γ, the estimates of β and ϕ are identified by differences across time in the employment trajectories between sons in the same family. Our central identifying assumption is that these are (conditionally) uncorrelated with the error term, ε, in (1.5). In principle, to ensure consistent estimates, the sibling differences in employment trajectories should due to differential shocks to labor demand between sons across time, and not to differential shocks to labor supply. Of the full sample of 9,279 sons, 5,642 also have a brother in the created matched panel.

Since the two-way fixed effect estimates can be calculated only for men in the sample with brothers, as a point of reference column 3 of Panel A in Table 3 shows the (one-way) individual fixed-effect estimates for the subsample of men from column 2 with brothers. A comparison of column 1 in Table 2 (full sample) to column 3 in Table 3 (sub-sample with brothers) shows that the one-way fixed effect estimates are very similar.[8] Column 4 of Panel A in Table 3 shows estimates with both individual and family fixed effects. To the third decimal place, the results are identical to those with just individual fixed effects.[9]

One potential concern about the identification in column 4 is that it depends on the exogeneity of differential shocks to labor demand between sons across time. These shocks could be spatially correlated for two reasons. First, brothers might be exposed to the same local economic conditions (in the extreme, for example, laid off by the same employer). Second, a parent might be more aware of the employment status of sons who live close by, rather than far

[8] We show a parallel set of estimates using income as the focal explanatory variable in column 6.

[9] For the two-way individual and family fixed-effect estimates in column 4, heteroscedasticity-robust standard errors clustered at the family level are shown in parentheses.

away from the parent.[10] To account for the correlation of real or reported employment shocks across space and siblings, column 5 provides a further robustness check by controlling for a sibling-pair locational fixed effect. Specifically, for each child, the HRS asks the parent whether the child lives within a 10-mile radius of the parent. The sibling-pair locational effect takes on a value of one if both sons in the sibling pair live within 10 miles of the parent and zero otherwise. Since even within metropolitan areas, much of the Great Recession's effect was localized (e.g., unemployment, foreclosures, etc.), this fixed effect accounts for whether the sons were subject to the same local economic conditions. As the results in column 5 indicate, the estimated impacts of employment loss are effectively unchanged.

In panel B, we repeat this exercise for daughters. Of the full sample of 7,345 daughters, 3,864 also have a sister. Hence, we show in columns 3 and 4 the one-way individual fixed effects and two-way individual and family fixed effects estimates on this subsample of 3,864 women. Those estimates are very similar to the full sample estimates shown in Table 2. Overall, unobserved family fixed effects do not appear to be a threat to the validity of our estimates.

6. Extensions

Next, we examine the extent to which the impact of employment transitions differs by subgroups based on a variety of child and parent characteristics measured in the baseline year (i.e., the year the parent entered the HRS). Since our individual fixed-effects estimates are quite robust, we show only estimates based on individual fixed effects in these specifications.

a. Differences by Baseline Child Characteristics

[10]Moreover, the parent likely knows far more about the employment of children who are co-residing than those living independently.

Panel A in Table 4 shows fixed-effect estimates for subsamples of sons split by the son's baseline marital status, education, and homeownership status, respectively. The response of co-residence to a transition away from full-time employment differs by the son's educational attainment (columns 3-4), but not by the son's marital status (columns 1-2) or homeownership status (columns 5-6). A parallel set of estimates are presented for daughters in Panel B. The response of co-residence to a transition away from full-time employment differs by the daughter's homeownership status (columns 3-4), but not by the daughter's marital status (columns 1-2) or education (columns 3-4). In particular, daughters who were renters at baseline were more likely to move back in with a parent after a transition away from full-time employment.

b. Differences by Baseline Parental Characteristics

Tables 5-7 examine how the response of co-residence to employment differs by the characteristics of parents. In Table 5, we report estimates from stratifying the sample by three key measures of parental socio-economic attributes: educational attainment, wealth, and race. Although some of the point estimates across sub-samples are statistically different based on the *p*-values shown in the table, there are few consistent difference among sons and daughters.

In Table 6, we report estimates from stratifying the sample by housing attributes of the parents in the baseline year, as an important factor in the co-residence decision is the quality and quantity of housing services supplied by the parent. In columns 1-2, we split the sample by the parents' housing tenure status; in columns 3-4, we split by the number of bedrooms in the residence (less than and greater than or equal to the median); and, in columns 5-6, we split by quality (low vs. not low). Importantly, the quality measure is based on observations made by the interviewer about the condition of the parental residence during the interview. It is not based on the parents' or children's self-assessments of quality. We define "low" quality to be a residence that was dirty or of poor quality. All other residences were deemed "not low quality." Based on the estimates in the table, we find that for sons the impact of employment changes on co-residence depends on the quality and quantity of housing services supplied by the parent: in particular, co-residence is substantially less likely with parents who are renters, have smaller, and low-quality residences. We do not find these patterns for daughters.

Intergenerational living simultaneously may provide both non-employment insurance to the adult child and informal care insurance to the parent(s), so that adult children who become unemployed and have parents in poor health or having experienced adverse health shocks may be differentially more likely to co-reside. We examine this in Table 7, where we split the sample above and below the medians of three baseline parental health measures: ADLs, IADLs, and the number of medical conditions. We find little evidence that the responsiveness of parental co-residency to changes in child employment depends on the health of the parent.

7. Summary

We present new estimates of the impact of transitions away from full-time employment on co-residence between parents and adult children. We have four primary findings. First, on average, a young man moving from full-time employment to being not employed raises the likelihood of co-residing with a parent by 1.7 percentage points. Similarly, on average moving from full-time employment to being part-time employed raises the likelihood of co-residing with a parent by 2.1 percentage points. The implied income elasticity of co-residence is -1.3. These effects are economically large, statistically significant, and suggest that the ability to move in with parents is an important way that sons adjust during economic downturns. We find more muted responses for daughters. Second, methodologically we assess the robustness of our main individual fixed-effect estimates with a set of two-way fixed effect estimates, based on both individual- and family-level effects. Third, the impact of employment changes on co-residence depends on the quality and quantity of housing services supplied by the parent: for sons, co-residence is substantially less likely with parents who rent, have smaller, and low-quality residences. Finally, we find little robust evidence that the health characteristics of the parents have differential impacts on how changes in employment translate into shared living arrangements. Overall, these results confirm and extend the findings of Kaplan (2009, 2012), Lee and Painter (2013), and others that parents use housing and shared living arrangements as a form of risk-sharing in the face of labor-market uncertainty for their adult children.

We temper these conclusions with two important caveats. First, the transitions from full-time to non-employment that we measure using the HRS data could be either from unemployment or departures from the labor force. These two possibilities have distinct implications for labor-market dynamics that we cannot separately identify. Second, our sample runs 1998-2012, and much of the employment loss that identifies our estimates can be traced to the 2000 recession and the Great Recession. The full impacts of the latter on housing and living arrangements have yet to play out completely, so that our findings should be best thought of as the short- to medium-term effects of employment loss on housing and living arrangements. As more post-recovery data become available, we will be able to estimate longer-run impacts. These are important avenues of future research.

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Table 1. Sample Means for Variables Used in Analysis

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	All Sons				All Daughters			
		Subsample:				Subsample:		
Characteristics	Full Sample	Working Full-Time	Working Part-Time	Not Working	Full Sample	Working Full-Time	Working Part-Time	Not Working
A. Individual Characteristics								
Age	38.0	37.9	37.8	38.7	37.9	37.9	37.6	37.8
Working full-time	0.905	1	0	0	0.800	1	0	0
Working part-time	0.023	0	1	0	0.069	0	1	0
Not working	0.072	0	0	1	0.131	0	0	1
Income	56,956	59,111	27,173	26,793	56,363	56,919	56,325	52,349
B. Outcomes								
Co-resides with a parent	0.021	0.016	0.080	0.064	0.019	0.017	0.020	0.027
C. Parental Characteristics								
Father is working	0.460	0.465	0.454	0.380	0.456	0.456	0.475	0.447
Mother is working	0.390	0.396	0.338	0.323	0.386	0.394	0.384	0.337
Father is homeowner	0.866	0.874	0.811	0.773	0.861	0.859	0.915	0.840
Mother is homeowner	0.800	0.811	0.716	0.673	0.797	0.799	0.838	0.757
Father’s age	65.9	65.9	66.6	66.0	65.8	65.8	65.9	65.6
Mother’s age	63.6	63.5	63.4	64.1	63.6	63.6	64.0	63.4
Father’s ADL limits	0.254	0.239	0.373	0.444	0.252	0.247	0.222	0.294
Mother’s ADL limits	0.280	0.262	0.334	0.511	0.293	0.287	0.250	0.365
Father’s IADL limits	0.108	0.101	0.161	0.185	0.121	0.120	0.104	0.138
Mother’s IADL limits	0.083	0.077	0.122	0.156	0.092	0.086	0.098	0.128
Number of father’s medical conditions	1.8	1.8	2.0	2.1	1.8	1.8	1.8	1.8
Number of mother’s medical conditions	1.8	1.8	2.0	2.3	1.9	1.8	1.8	2.0
Father's household income	78,293	79,996	64,593	57,249	73,405	72,461	81,758	74,543
Mother's household income	61,650	63,461	49,817	40,167	60,172	59,493	71,170	58,269

Father's family wealth	526,254	545,502	344,745	296,762	487,067	477,989	556,255	504,180
Mother's family wealth	428,580	449,365	249,343	197,988	395,692	386,133	494,954	401,039
Number of bedrooms in father’s residence	3.2	3.2	3.6	3.4	3.3	3.3	3.5	3.5
Number of bedrooms in mother’s residence	3.3	3.3	3.6	3.5	3.4	3.4	3.5	3.5

Note: Authors’ calculations based on HRS sample described in text.

Table 2. Individual Fixed-Effect Estimates of the Impact of Employment Status on Parental Co-Residence for All Sons and Daughters, Respectively, Robust Standard Errors in Parentheses

	(1)	(2)	(3)	(4)
	Sample:			
Explanatory Variable	All Sons	All Sons	All Daughters	All Daughters
Focal Explanatory Variables				
Not employed	0.017 (0.004)	---	0.006 (0.002)	---
Part-time employed	0.021 (0.007)	---	0.005 (0.003)	---
Log of income	---	-0.021 (0.004)	---	-0.011 (0.003)
Control Variables				
Age	-0.004 (0.005)	0.003 (0.003)	-0.0008 (0.002)	-0.00005 (0.002)
Father works	0.0001 (0.002)	-0.006 (0.002)	0.0005 (0.002)	0.007 (0.003)
Father is married	-0.017 (0.005)	-0.019 (0.010)	0.003 (0.006)	-0.008 (0.010)
Father’s limits to activities of daily living	0.005 (0.002)	-0.005 (0.003)	-0.001 (0.001)	0.0004 (0.004)
Father’s limits to instrumental activities of daily living	0.005 (0.003)	0.006 (0.007)	0.002 (0.003)	0.001 (0.006)
Father’s number of medical conditions	0.0002 (0.001)	-0.00008 (0.002)	-0.0006 (0.001)	0.004 (0.002)
Father’s age	0.0002 (0.0001)	0.0002 (0.0003)	0.0000009 (0.0001)	0.0002 (0.0003)
Number of father’s children	0.0005 (0.0008)	0.002 (0.002)	0.001 (0.001)	0.0009 (0.002)
Father’s income	0.0000003 (0.0000003)	0.00002 (0.00002)	0.000005 (0.000006)	0.00003 (0.00002)
Father’s wealth	0.0000007 (0.0000006)	0.0000008 (0.000001)	0.00000003 (0.0000008)	0.000001 (0.000001)
Father owns home	0.001 (0.004)	0.010 (0.008)	0.0006 (0.005)	-0.014 (0.011)
Number of bedrooms in father’s residence	0.001 (0.001)	0.002 (0.002)	0.002 (0.006)	0.002 (0.002)
Mother works	0.001 (0.002)	0.003 (0.003)	-0.002 (0.002)	-0.001 (0.004)

Mother is married	-0.016	-0.028	-0.023	-0.029
	(0.004)	(0.008)	(0.006)	(0.011)
Mother’s limits to activities of daily living	-0.002	-0.0007	0.004	0.003
	(0.001)	(0.003)	(0.002)	(0.003)
Mother’s limits to instrumental activities of daily living	-0.002	0.003	0.0007	-0.0004
	(0.051)	(0.006)	(0.003)	(0.005)
Mother’s number of medical conditions	0.003	0.002	0.003	0.006
	(0.001)	(0.002)	(0.001)	(0.003)
Mother’s age	0.00009	0.0004	0.0002	0.0003
	(0.0002)	(0.0003)	(0.0002)	(0.0003)
Number of mother’s children	0.001	0.001	-0.0009	-0.003
	(0.001)	(0.002)	(0.002)	(0.004)
Mother’s income	-0.000001	-0.00001	0.00001	-0.00001
	(0.000009)	(0.00001)	(0.000008)	(0.00002)
Mother’s wealth	-0.0000004	-0.0000004	-0.00000005	-0.000002
	(0.0000007)	(0.000001)	(0.0000009)	(0.000002)
Mother owns home	-0.007	-0.016	-0.016	-0.007
	(0.004)	(0.009)	(0.006)	(0.010)
Number of bedrooms in mother’s residence	0.004	0.006	0.005	0.006
	(0.002)	(0.002)	(0.002)	(0.002)
Number of individuals	9,279	7,460	7,345	5,870

Note: The dependent variable for all regressions takes on a value of one if the son co-resided with the parent and zero otherwise. Columns 1 and 3 show the parameter estimates of beta and phi in equation (1.1) from separate regressions, with the associated sample shown in the column heading; columns 2 and 4 show parameter estimates of theta in (1.2). There are three employment statuses: full-time, part-time, and not employed. Therefore, the excluded group is comprised of those who are full-time employed. All regressions included a set of controls for the calendar year. All standard errors are heteroscedasticity-robust and clustered at the individual-level.

Table 3. Estimates of the Impact of Employment Status on Parental Co-Residence for All Sons and Daughters, Respectively, Using Alternative Estimators, Robust Standard Errors in Parentheses

	(1)	(2)	(3)	(4)	(5)	(6)
	Sample and Estimator:					
	Full Sample		Subsample with Same-Sex Siblings			
Explanatory Variable	Arellano-Bond	Arellano-Bond	Individual Fixed Effects	Individual and Family Fixed Effects	Individual, Family, and Sibling-Location Fixed Effects	Individual Fixed Effects
A. All Sons						
Not employed	0.010	---	0.015	0.015	0.015	---
	(0.004)		(0.005)	(0.006)	(0.006)	
Part-time employed	0.016	---	0.023	0.023	0.023	---
	(0.007)		(0.009)	(0.010)	(0.010)	
Log income	---	-0.018	---	---	---	-0.013
		(0.004)				(0.005)
Number of men	9,167	6,795	5,642	5,642	5,642	4,389
B. All Daughters						
Not employed	0.004	---	0.004	0.004	0.004	---
	(0.002)		(0.004)	(0.004)	(0.004)	
Part-time employed	0.005	---	0.002	0.002	0.002	---
	(0.003)		(0.004)	(0.004)	(0.004)	
Log income	---	-0.010	---	---	---	-0.010
		(0.004)				(0.005)
Number of women	7,274	3,157	3,864	3,864	3,864	2,917

Note: The dependent variable for all regressions is an indicator that takes on a value of one if the son co-resided with the parent and zero otherwise. Column 1 shows the parameter estimates of beta and phi in equation (1.3); column 2 shows the parameter estimate of theta in equation (1.4). Columns 3-5 show the parameter estimates of beta and phi in equation (1.5). Column 6 shows the two-way fixed-effect parameter estimate of theta in equation (1.2). There are three employment statuses: full-time, part-time, and not employed. Therefore, the excluded group is comprised of those who are full-time employed. All regressions included a set of controls for the calendar year and the variables for each biological parent: age, number of ADLs, number of IADLs, number of medical conditions, number of bedrooms in residence, income, wealth, and indicators for homeownership, employment, and marital status. Standard errors in columns 1-3 and 6 are heteroscedasticity-robust and clustered at the individual-level; standard errors for columns 4-5 are clustered at the family level.

Table 4. Individual Fixed-Effect Estimates of the Impact of Employment Status on Parental Co-Residence for All Sons and Daughters, Respectively, by Selected Subgroups based on Child Characteristics at Baseline in the Sample, Robust Standard Errors in Parentheses

	(1)	(2)	(3)	(4)	(5)	(6)
	Sample Split by Child's Baseline Characteristic:					
	Marital Status		Education		Homeownership	
Explanatory Variable	Married	Not Married	High School or Less	More than High School	Rent	Own
A. All Sons						
Not employed	0.012	0.023	0.010	0.031	0.018	0.014
	(0.005)	(0.007)	(0.005)	(0.007)	(0.005)	(0.006)
Part-time employed	0.017	0.024	0.023	0.013	0.022	0.018
	(0.008)	(0.011)	(0.009)	(0.012)	(0.009)	(0.010)
p-value for test of equality of parameters across sub-samples	0.335		0.034		0.864	
Number of men	6,195	3,084	4,861	4,418	4,220	5059
B. All Daughters						
Not employed	-0.007	0.005	0.007	0.005	0.011	-0.001
	(0.005)	(0.003)	(0.004)	(0.003)	(0.004)	(0.002)
Part-time employed	0.008	0.004	0.007	0.005	0.012	0.001
	(0.006)	(0.003)	(0.004)	(0.003)	(0.005)	(0.003)
p-value for test of equality of parameters across sub-samples	0.803		0.824		0.011	
Number of women	4,830	2,515	3,317	4,028	3,212	4,133

Note: The dependent variable for all regressions takes on a value of one if the son co-resided with the parent and zero otherwise. All columns show the parameter estimates of beta and phi in equation (1.1) from separate regressions, with the associated sample shown in the panel and column heading. There are three employment statuses: full-time, part-time, and not employed; the excluded group is comprised of those who are full-time employed. All regressions included a set of controls for the calendar year and the variables for each biological parent: age, number of ADLs, number of IADLs, number of medical conditions, number of bedrooms in the residence, income, wealth, and indicators for homeownership, employment, and marital status. All standard errors are heteroscedasticity-robust and clustered at the individual-level.

Table 5. Individual Fixed-Effect Estimates of the Impact of Employment Status on Parental Co-Residence for All Sons and Daughters, Respectively, by Selected Subgroups Based on Parental Socio-Economic Status at Baseline in the Sample, Robust Standard Errors in Parentheses

	(1)	(2)	(3)	(4)	(5)	(6)
	Sample Split by Parental Baseline Characteristic:					
	Highest Education Achieved by Parents		Parental Wealth		Race of Parents	
Explanatory Variable	High School or Less	More than High School	Below Median	Above Median	At Least One Non-White	Both White
A. All Sons						
Not employed	0.012	0.019	0.008	0.031	0.005	0.022
	(0.009)	(0.004)	(0.005)	(0.007)	(0.006)	(0.005)
Part-time employed	0.010	0.025	0.026	0.012	0.035	0.018
	(0.016)	(0.008)	(0.009)	(0.011)	(0.017)	(0.008)
p-value for test of equality of parameters across sub-samples	0.596		0.016		0.070	
Number of men	2,311	6,968	4,407	4,872	1,199	7,364
B. All Daughters						
Not employed	-0.001	0.008	0.004	0.007	0.013	0.005
	(0.006)	(0.003)	(0.003)	(0.004)	(0.006)	(0.002)
Part-time employed	0.006	0.006	0.006	0.006	-0.008	0.007
	(0.005)	(0.003)	(0.004)	(0.003)	(0.011)	(0.003)
p-value for test of equality of parameters across sub-samples	0.306		0.851		0.176	
Number of women	1,872	5,473	3,639	3,706	1,124	5,621

Note: The dependent variable for all regressions takes on a value of one if the son co-resided with the parent and zero otherwise. All columns show the parameter estimates of beta and phi in equation (1.1) from separate regressions, with the associated sample shown in the panel and column heading. There are three employment statuses: full-time, part-time, and not employed; the excluded group is comprised of those who are full-time employed. All regressions included a set of controls for the calendar year and the variables for each biological parent: age, number of ADLs, number of IADLs, number of medical conditions, number of bedrooms in the residence, income, wealth, and indicators for homeownership, employment, and marital status. All standard errors are heteroscedasticity-robust and clustered at the individual-level.

Table 6. Individual Fixed-Effect Estimates of the Impact of Employment Status on Parental Co-Residence for All Sons and Daughters, Respectively, by Selected Subgroups Based on Parental Housing Characteristics at Baseline, Robust Standard Errors in Parentheses

	(1)	(2)	(3)	(4)	(5)	(6)
	Sample Split by Parental Housing Attributes					
	Parental Homeownership		Number of Bedrooms in Parental Residence		Quality of Housing	
Explanatory Variable	Rent	Own	Less than Three	Three or More	Low Quality	Not Low Quality
A. All Sons						
Not employed	0.002	0.022	0.021	0..017	-0.016	0.019
	(0.006)	(0.005)	(0.033)	(0.004)	(0.009)	(0.004)
Part-time employed	0.013	0.024	-0.021	0.022	0.102	0.017
	(0.012)	(0.009)	(0.029)	(0.007)	(0.047)	(0.007)
p-value for test of equality of parameters across sub-samples	0.025		0.335		0.003	
Number of men	1,756	7,523	8,666	613	227	9,052
B. All Daughters						
Not employed	0.009	0.004	-0.008	0..006	0.044	0.005
	(0.006)	(0.003)	(0.016)	(0.003)	(0.018)	(0.003)
Part-time employed	0.003	0.006	0.033	0.005	0.045	0.004
	(0.006)	(0.003)	(0.025)	(0.003)	(0.025)	(0.003)
p-value for test of equality of parameters across sub-samples	0.727		0.533		0.053	
Number of women	1,485	5,860	6,876	469	197	7,148

Note: The dependent variable for all regressions takes on a value of one if the son co-resided with the parent and zero otherwise. All columns show the parameter estimates of beta and phi in equation (1.1) from separate regressions, with the associated sample shown in the panel and column heading. There are three employment statuses: full-time, part-time, and not employed; the excluded group is comprised of those who are full-time employed. All regressions included a set of controls for the calendar year and the variables for each biological parent: age, number of ADLs, number of IADLs, number of medical conditions, income, wealth, and indicators for employment and marital status. All standard errors are heteroscedasticity-robust and clustered at the individual-level.

Table 7. Individual Fixed-Effect Estimates of the Impact of Employment Status and Income on Parental Co-Residence for All Sons and Daughters, Respectively, by Selected Subgroups Based on Parental Functional and Health Characteristics at Baseline, Robust Standard Errors in Parentheses

	(1)	(2)	(3)	(4)	(5)	(6)
	Sample Split by Parental Functional and Health Attributes					
	Number of Limits to Activities of Daily Living a Parent Has		Number of Limits to Instrumental Activities of Daily Living a Parent Has		Number of Medical Conditions a Parent Has	
Explanatory Variable	None	One or More	None	One or More	Less than Three	Three or More
A. All Sons						
Not employed	0.017	0.018	0.017	0.015	0.021	0.012
	(0.005)	(0.006)	(0.004)	(0.011)	(0.005)	(0.006)
Part-time employed	0.018	0.021	0.021	0.025	0.029	0.010
	(0.008)	(0.014)	(0.007)	(0.021)	(0.010)	(0.009)
p-value for test of equality of parameters across sub-samples	0.689		0.963		0.210	
Number of men	7,702	1,577	8,601	678	6,860	2,419
B. All Daughters						
Not employed	0.006	0.005	0.005	0.012	0.005	0.008
	(0.003)	(0.003)	(0.003)	(0.017)	(0.003)	(0.005)
Part-time employed	0.005	0.006	0.006	0.004	0.004	0.005
	(0.006)	(0.003)	(0.003)	(0.007)	(0.003)	(0.005)
p-value for test of equality of parameters across sub-samples	0.727		0.592		0.476	
Number of women	6,054	1,291	6,738	607	5,326	2,019

Note: The dependent variable for all regressions takes on a value of one if the son co-resided with the parent and zero otherwise. All columns show the parameter estimates of beta and phi in equation (1.1) from separate regressions, with the associated sample shown in the panel and column heading. There are three employment statuses: full-time, part-time, and not employed; the excluded group is comprised of those who are full-time employed. All regressions included a set of controls for the calendar year and the variables for each biological parent: age, number of bedrooms in the residence, income, wealth, and indicators for homeownership, employment, and marital status. All standard errors are heteroscedasticity-robust and clustered at the individual-level.