WHY DO MOST ITALIAN YOUTHS LIVE WITH THEIR PARENTS? INTERGENERATIONAL TRANSFERS AND HOUSEHOLD STRUCTURE

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Abstract
More than 80% of Italian men aged 18–30 live with their parents. We argue that one contributing factor to this remarkably high rate of cohabitation is parents' tastes for coresidence. In order to investigate the role of parental preferences, we estimate the effect of exogenous changes in parental income on rates of cohabitation in Italy using Survey of Households' Income and Wealth (SHIW) micro data from 1989 to 2000. In order to identify a source of exogenous variation in parental income, we use changes in fathers' retirement age induced by the 1992 reform of the Italian Social Security system as an instrumental variable for parental income. By raising retirement age, this reform forced some fathers to remain in the labor market longer than they would have otherwise, therefore raising their disposable income. We use a two-sample instrumental variable (TSIV) strategy. Our TSIV estimates indicate that a rise in parents' income significantly raises the children's propensity to live at home: A 10% increase in annual parental income results in approximately a 10% rise in the proportion of men living with their parents. Although we cannot definitely rule out alternative interpretations, these results are consistent with our hypothesis that cohabitation is a normal good for Italian parents. (JEL: J120, J610, H550)

1. Introduction

Young Italians are considerably more likely to live with their parents than are their Northern European and U.S. counterparts. Data from the European Community Household Panel Survey (ECHPS) for example show that around 82% of Italian

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TABLE 1. Percentage of children living with parents. Men aged 18–30.

	Percentage
France	45.40
United Kingdom	53.01
Germany	44.99
Italy	82.19
Spain	65.26
U.S.	43.00
Portugal	78.14

Note: Sources: US: CPS monthly files 1996. Rest of countries in the sample: ECHPS, 1996.

men aged 18–30 live with their parents.[1] As Table 1 shows, this is not a feature unique to Italy: In other Southern European countries rates of coresidence are in the same order of magnitude. For Portugal, for example, the same survey gives an estimate of cohabitation rates on the order of 78%. For Spain the figure is around 65%. In contrast, CPS data show that only 43% of American men in the same age group live with their parents. French, German, and British men display rates of cohabitation between 45% and 53%, somewhat between Southern Europe and the U.S.

Obvious explanations for the high fraction of cohabiting young Italian men are the high youth unemployment rate and housing costs. Some authors have found, indeed, that youth labor market conditions are important determinants of young individuals' living arrangements. For example, Card and Lemieux (2000) find that poor labor market conditions in Canada explain why the fraction of youth living with their parents has increased in Canada relative to the U.S. in recent years. Other authors focus on different explanations for the remarkably high rates of coresidence among young Italians.[2] Becker et al. (2003), for example, concentrate on the role played by job insecurity on coresidence decisions of young Italians in the spirit of the theoretical model of Fogli (2000). They argue that moving-out decisions are irreversible and therefore higher job insecurity tends to decrease the probability of leaving the parental nest. In a recent paper, Giuliano (2002) argues that the more liberal attitudes brought by the sexual revolution have allowed Southern Europeans to cohabit with their parents without having to give up their sexual activity.

While we do not rule out the importance of these factors, in this paper we focus specifically on the role preferences and intra-household transfers play in shaping living arrangements. We show that if cohabitation is a "good" for parents and a

1. In this study we only concentrate on living arrangements of young men because our instrumental variable is not available for women for the whole period of observation.

2. Ruiz-Castillo and Martínez-Granado (2002) study the living arrangements decisions of Spanish youths. They jointly model the decisions of working, studying, and leaving the parental household, similar to the analysis in McElroy (1985).

"bad" for children, parents will be willing to trade off some of their consumption in order to "bribe" their children, that is, to compensate those children who remain at home by offering them higher consumption in exchange for their presence at home. One testable implication of our model is that, all else equal, an exogenous rise in parental income should be associated with a rise in the probability of coresidence.

We estimate the effect of parental income on Italian children's living arrangements using micro data from the Bank of Italy Survey of Households' Income and Wealth (SHIW) from 1989 to 2000. The key problem in estimating the effect of parental income on the children's propensity to cohabit is the potential endogeneity of parental income. Parental income is likely endogenous because of the endogeneity of parental labor supply (parents may decide to work more in order to support their cohabiting children) or the altruistic behavior of children (if parents suffer negative income or health shocks, their children may invite their parents to live with them) or other sources of unobserved heterogeneity.

To address this problem, we use arguably exogenous changes induced by a reform of social security as an instrumental variable for parental income. Specifically, our instrumental variable is based on the increase in normal retirement age mandated by the 1992 reform of social security. The reform gradually increased normal retirement age from 60 at the beginning of the period to 64 at the end of the period. This change in the retirement law forced some cohorts of parents to stay in the labor force longer than they would have stayed otherwise. Because replacement ratios are typically less than 1 and retirement is associated with an income loss, the reform should increase the income of some cohorts of parents but not others.

Consistently with this prediction, we show that the cohorts affected by the reform experienced an increase in disposable income relative to cohorts who were not affected by the reform. The effect is small, but precisely estimated. A feature of the reform is that mandated retirement age increased over time (from 60 at the beginning of the period to 64 at the end). This is useful because it allows us to include unrestricted father's age effects and time effects in the model. Identification comes from the interaction of father's age and year of the mandated changes.

Because our instrument is based on a change in retirement eligibility, not on the actual retirement decision, it is arguably exogenous with respect to other determinants of living arrangements. In support of this, we show that the instrument is orthogonal to many observable exogenous characteristics of parents and children. We also show results from several specification tests that indicate that the reform is unlikely to have had an effect on children living arrangements for reasons other than the effect on fathers' disposable income.

A second problem in estimating the effect of parental income on children's living arrangements arises because data on parental income are available only for cohabiting children. As with most existing household surveys, our data lack information on parental income for noncohabiting children. To address this issue, we use a two-sample instrumental variable strategy (Angrist and Krueger 1992). This strategy is feasible because the instrument is based on the interaction of father's age and year and our data have information on parental age for all men, both cohabiting and noncohabiting.

Our two-sample instrumental variable estimates suggest that parents' income is an important determinant of their children's propensity to live at home. A 1 million lira increase in annual parents' income (approximately \$500 or €500) raises the probability of cohabitation by between 3.5 and 3.9 percentage points. This is equivalent to an elasticity around 1.3–1.5. The estimates are robust to controls for local labor and housing market conditions, and for standard socio-economic characteristics, such as parents' education and children's age. Although different interpretations are possible, one possible interpretation that is consistent with our model is that cohabitation is a normal good for Italian parents and that this is an important factor in explaining the remarkably high rates of coresidence among young Italians.

Others before us have emphasized the noncooperative nature of relationship between parents and children. Cox (1987), for example, proposes an exchange model where parents derive some utility from services provided by children (care or housecleaning, for example) as well as "more subtle types of service that entails the behavioral constraints associated with attention to parents …; companionship and conforming to parental regulations." Parents might be willing to trade some consumption in exchange for these services. The evidence he provides on inter vivos transfers in the U.S. lends some support to the exchange model.[3]

Probably more controversial is our claim that Italian parents like to live with their children. Indeed, existing evidence suggests that the opposite might be true in the U.S. (Rosenzweig and Wolpin 1993, 1994). Estimates based on the Nation Longitudinal Survey show that cohabitation rates tend to fall as parental income rises suggesting that for U.S. fathers "privacy is a normal good" (Rosenzweig and Wolpin 1993, p. 104).

Using data from seven industrialized countries, in the concluding section of the paper we provide some additional evidence that indicates that parental tastes for cohabitation may differ across countries. Although Italian parents seem to be happier when they live with their children, the opposite seems to be true for parents in the U.S., U.K., and Germany.

3. Altonji, Hayashi, and Kotlikoff (1997) examine intrahousehold transfers among members of extended families in the U.S. In their model living arrangements are taken as given. Del Boca (1997) studies intrahousehold transfers in Italy but she also takes household structure as given. A somewhat related stream of literature analyzes the effect of pension and welfare transfers on living arrangements. Costa (1997) and McGarry and Shoeni (2000) study the effect of pensions on living arrangement decisions of the elderly. Recently, Bitler, Gelbach, and Hoynes (2001) study the effect of the U.S. welfare reforms of the 1990s on the living arrangements of children and women.

Recent evidence also confirms that differences in preferences might play a role in explaining different rates of cohabitations across different cultures. For example, Giuliano (2002) shows that second-generation Southern Europeans in the U.S. are more likely to live with their parents than observationally equivalent individuals of different ancestry. This may indicate that labor and housing market conditions are not the only determinants of the differences in cohabitation rates across countries. Similarly, evidence from developing countries suggests that the determinants of cohabitation are different from the ones that have been documented for the U.S. In a paper that is somewhat similar in spirit to ours, Edmonds, Mammen, and Miller (2001) examine living arrangements decisions in South Africa. They show sizeable changes in household composition, as a grandparent becomes pension-eligible; however, contrary to the evidence from high-income countries, they do not find an effect on the propensity of grandparents to live alone, whereas they find sizeable changes in living arrangements of young and prime age-individuals.

The rest of the paper is organized as follows. In Section 2 we present a stylized model of living arrangements. In Section 3 we describe the data and present our main empirical results. In Section 4 we present some additional evidence. In Section 5 we discuss possible alternative interpretations of our evidence. Section 6 concludes.

2. A Simple Framework

In this section we present a stylized model of children's housing arrangements. We emphasize that the point of this simple model is not to specify a structural equation to be estimated, but rather to illustrate the implications of different assumptions and to make sense of the empirical results in the next section.

We think of living arrangements as the outcome of a noncooperative game between parents and children. We assume that children value their independence and, everything else equal, prefer to live on their own. Parents prefer to live with their children and are willing to offer them an income transfer (the "bribe") if children decide to live with them. The model shows that a rise in parental income tends to increase the children's propensity to live at home.[4] This result depends on the assumption of decreasing marginal utility of consumption. The loss in consumption associated with the bribe results in a lower utility loss for rich parents than for poor parents. As income rises, parents are more willing to bribe their children to keep them at home. This is generally true if parents are selfish, and it remains true under some conditions if parents are altruistic (i.e., they internalize their children's distaste for cohabitation).

4. The assumption that children prefer to live on their own is not crucial to our results. If both parents and children prefer cohabitation, there is no bargaining between parents and children and this result obviously still holds.

We start by assuming that parents are selfish. We model coresidence decisions as a noncooperative game in which parents are first movers. For simplicity, we consider a household composed of one child and one parent. The parent's utility is a function of consumption and living arrangements. The parent offers an income transfer to his child but only if he decides to cohabit. For simplicity, we assume Stone Geary preferences. The parent's problem can be written as

$$\begin{aligned} \text{Max} \quad & U_P(C_P, H) = [\log(C_P) + H \log(a_P)] \\ \text{Subj. to:} \quad & C_P = Y_P - b_1 H \end{aligned} \tag{1}$$

where C is consumption, H is a dummy variable equal to 1 if the child is living at home, and the parameter $\log(a_P)$ denotes the parents' marginal utility of cohabitation (with $a_P \geqslant 1$). The parameter b_1 is the transfer to cohabiting children. The budget constraint states that parents' consumption is a function of their income minus any transfer to their cohabiting children.

Children maximize the following utility function:

$$\begin{aligned} \text{Max} \quad & U_K(C_K, H) = \log(C_K) + H \log(a_K) \\ \text{Subj. to:} \quad & C_K = Y_K - R(1 - H) + b_1 H, \end{aligned} \tag{2}$$

where Y_K is children's income and R their housing costs. Children's consumption is financed either by parental transfers or their own income. $\log(a_K)$ is the children's disutility of living at home ($0 < a_K < 1$). The model assumes that children pay for their own housing cost if they leave the parental home.[5]

Parents set the optimal transfer to cohabiting children, $b_1{}^*$, so that the children are indifferent between living at home or on their own: $U_K(Y_K + b_1{}^*, 1) = U_K(Y_K - R, 0)$. It follows that $b_1{}^* = (Y_K - R)/a_K - Y_K$. Conditional on $b_1{}^*$, parents are willing to bribe their children into staying at home if the utility they derive from cohabitation is higher than the utility from separation, $U_P(Y_P - b_1^*, 1) \geqslant U_P(Y_P, 0)$.

In equilibrium,

$$Pr(H = 1) = Pr(Y_P \geqslant A_1 Y_K - A_2 R), \tag{3}$$

where $A_1 = [(1 - a_K)a_P]/[a_K(a_P - 1)] > 0$ and $A_2 = a_P/[a_K(a_P - 1)] > 0$. The model predicts that if cohabitation is a good for parents and a bad for children,

5. The model should also include parents' housing costs. We assume that these are borne by parents irrespective of their children's living arrangements and that they are unchanged when children move out of the parental household. So parents' housing costs do not affect the children's marginal decision as to whether to live with their parents or not. To simplify the notation, they are not explicitly included in the model.

conditional on children income and housing costs (plus preferences), a rise in parents' income is associated with a rise in cohabitation rates.[6]

What happens if parents are altruistic? One way to think about altruistic parents is to assume that children's utility enters the parents' utility function, so that parents internalize their children's distaste for cohabitation. When parents are altruistic and internalize their children's distaste for cohabitation, the effect of parental income on children living arrangements depends on the parents' utility function. If parents have a sufficiently high desire for coresidence, the model generates the same prediction obtained above for the selfish case: A rise in parental income is associated with an increase in cohabitation rates. We formally show this point in the Appendix.

3. Empirical Evidence

In this section we present some empirical evidence on the relationship between parental income and children's living arrangements. In Section 3.1 we describe the data. In Section 3.2 we outline our two-sample instrumental variable strategy and discuss our identification assumption. The main empirical results are in Section 3.3. In Section 3.4 we present additional results intended to probe the robustness of our estimates. In Section 4 we present suggestive evidence that lends further credibility to our claim that Italian parents draw utility from cohabitating with their children. We also show that the same is probably not true in other countries, especially the U.S.

3.1. Data

We use data from the individual records of the SHIW for 1989, 1991, 1993, 1995, 1998, and 2000. In addition to a wide array of socioeconomic variables, for each household-head the SHIW provides information on the year of birth of the father and mother, and whether they are still alive. This information is important because it allows us to recover parental age both for cohabiting children and for noncohabiting children. Although we have no information on the income or identity of parents of children living on their own, we have information on their parents' age. This piece of information will prove valuable for our two-sample instrumental variable strategy.

We construct two samples. The *children sample* includes men aged 18–30 whose parents are both alive, whose father is aged between 40 and 74 and whose

6. Note that this result only holds if $Y_K > R$. If $Y_K < R$ children will live at home irrespective of their parents' income and they will receive a zero bribe. Also, note that parents could in principle offer a zero compensation to those children who have low disutility of cohabitation; however, the condition in (3) remains unchanged. To keep the model simple we ignore this case.

TABLE 2. Descriptive statistics.

Men	
Living with parents	0.86
Age, years	23.65
Age of father, years	55.00
Age of mother, years	51.15
Working	0.50
Student	0.28
Income*	9.47
Observations	11,782
Parents	
Father's age, years	55.40
Mother's age, years	51.69
Father retired	0.39
Father working	0.60
Father's income*	24.80
Mother retired	0.22
Mother working	0.29
Mother's income*	8.06
Observations	21,381

Notes: The children sample includes all men 18–30 with a living father between the ages of 40 and 74 and a living mother between the ages of 37 and 71. The parents sample includes married couples with a husband between the ages of 40 and 74 and a wife between the ages of 37 and 71. Data source: SHIW individual records, 1989–2000. *All monetary variables are in 1995 million lira.

mother is aged 37 to 71.[7] We include in this sample both children living with their parents and those living on their own. We focus on men because the information on parental age for noncohabiting children was only collected for household heads, who are mostly men.[8]

The top panel in Table 2 reports descriptive statistics for children. Overall there are 11,782 observations in the children sample. About 86% of men aged 18 to 30 in the SHIW live with their parents.[9] Their average age is 24 and their fathers and mothers are, respectively, aged 55 and 51. About half of these young adults

7. These correspond approximately to the bottom and top percentiles of the distribution of parents' age for the children in the sample.

8. Parental age was only collected in the SHIW starting in 1989. We use all available data from 1989 onwards. Data on the year of birth of the parents of the head of household's spouse (mainly women) are collected only starting in 1993, which severely restricts our sample size. However, we provide estimates for women based on this restricted sample below. Because there is no way to identify grandchildren, parents, or grandparents of the heads of household in our data, we ignore children living in three-generation households headed by one of the children's grandparents. Similarly, we treat those children living with their parents or grandparents (or both) who are classified as heads of household as living on their own. Finally, we ignore children living only with one parent.

9. This proportion is higher than the one recorded in the ECHPS. One reason for this is that in the SHIW sample we condition on children whose parents are both alive.

work, suggesting that rationing in the labor market alone is unlikely to explain living arrangements.[10] Average annual income, expressed in lira at 1995 prices and defined net of taxes and social security contributions, amount to just less than 9 million lira per year. About 30% of these individuals are enrolled in school.

The second sample is the *parents sample*. It includes married individuals in couples whose head of household is a man aged 40–74 married to a woman aged 37–71. This sample includes both couples with children and without children, as no information on number of children is recorded consistently throughout our period of observation.[11]

The lower panel of Table 2 reports descriptive statistics for the parents. We have information on 21,381 parental households. Individuals in the parents' sample are on average slightly older than parents of those in the children sample, a reflection of the fact that older parents in the sample are less likely to have children in the age group 18–30.[12] About 60% of fathers work and the remaining 40% are retired. Average father's income is almost three times the average children's income and in the order of 25 million per year. About 30% of mothers work and 20% are retired. Mothers' average income is just 8 million per year.

3.2. Identification Strategy

The goal of this paper is to estimate the effect of changes in parental income on cohabitation. To do so, we estimate the following equation:

$$H_{it} = \beta_0 + \beta_1 Y_{Pit} + X'_{it}\beta_2 + u_{it} \tag{4}$$

H is a dummy equal to one if child i lives with his parents at time t; Y_P is parental income; X is a set of controls; u is the residual. The coefficient of interest is β_1. If parents have a sufficiently high taste for cohabitation, increases in parental income should raise the probability that children live with their parents: $\beta_1 > 0$.

A two-sample instrumental variable strategy. In estimating equation (4) we face two problems. First, data on both parental income and children's living arrangements are needed. Typically, however, household data do not contain information on parental income for those children who live on their own and our data are no exception. Although parental income for cohabiting children is observed in the SHIW, parental income of noncohabiting children is not available.

10. Work is defined as at least one employment spell in the year.

11. From both samples we also drop the observations with missing income, with income below the first percentile or above the ninety-ninth percentile and with missing age.

12. For an analysis of fertility (and women's labor supply) in Italy, see Del Boca (2002).

But even if parental income for noncohabiting children was available, parental income would arguably be endogenous to housing arrangements, and Ordinary Least Squares (OLS) estimates of β_1 would be inconsistent. There are many sources of unobserved heterogeneity. Consider, for example, unobserved shocks to local labor market conditions. In areas that are economically depressed, parental income is lower. At the same time, children's economic opportunities are also lower, inducing some of the children to live with their parents. In this case, one might find a negative correlation between parental income and coresidence rates, even in the absence of a causal effect of parental income and coresidence. Alternatively, it is also possible that the opposite bias arises, if children from depressed areas are more likely to leave the area and look for jobs elsewhere. In this case, one might find a positive spurious correlation between parental income and coresidence rates.

Another example of unobserved heterogeneity is represented by unobserved shocks to parents' health. A negative health shock may reduce parental income, and at the same time increase the probability that a child decides to live with his parents to take care of them. Heterogeneity in tastes may also induce spurious correlation between parental income and living arrangements. For example, it is possible that children who grew up in rich families have a tighter relationship with their parents than children who grew up in poor families, or vice versa.

In all these cases, one would find a relationship between parental income and children living arrangements for reasons that are unrelated to the causal effect we are interested in identifying. The sign of the bias can not be determined a priori. OLS estimates of β_1 are upward or downward biased depending on whether u is positively or negatively correlated with Y_P.

We use an instrumental variable strategy to address the problem of endogeneity of parental income. The ideal instrument is correlated with parental income but uncorrelated with all other factors that determine living arrangements (including children's labor market opportunities, tastes, unobserved health shocks, etc.). We describe our proposed instrument in the next section. But before doing so, we note that parental income of noncohabiting children is not available, and a standard instrumental variable strategy is not feasible.

Instead, we use a two-sample instrumental variable strategy (TSIV). To address the problem of missing parental income for noncohabiting children, the instrument must be available for all children (both cohabiting and noncohabiting) as well as for parents. With such an instrument, we estimate β_1 in two steps. First, we use the sample of parents to estimate the effect of the instrument on parental income:

$$Y_{Pit} = \gamma_0 + \gamma_1 Z_{it} + X'_{it}\gamma_2 + e_{it}, \tag{5}$$

where Z denotes the instrument. Equation (5) is the first-stage equation. Second, we use the sample of children to estimate the effect of the instrument on the probability that a child lives with his parents:

$$H_{it} = \theta_0 + \theta_1 Z_{it} + X'_{it}\theta_2 + v_{it}. \tag{6}$$

Equation (6) is the reduced form equation. Angrist and Krueger (1992) show that a consistent estimate of β_1 is given by the TSIV estimator, which is the ratio of the reduced form coefficient over the first stage coefficient:

$$est(\beta_1^{IV}) = est(\theta_1^{OLS})/est(\gamma_1^{OLS}). \tag{7}$$

We now describe the instrumental variable Z that we use in this paper.

Using pension reform to identify the effect of parents' income on coresidence. We propose using changes in Social Security eligibility and retirement age introduced in Italy in 1992 as an instrument for parental income. Retirement typically reduces disposable income, because replacement ratios are generally below one. We show that changes in the normal retirement age introduced by a 1992 reform of Social Security had a significant effect on parents' disposable income, and we use these changes as an instrument for parental income. We argue that (conditional on a number of covariates) the reform is uncorrelated with other determinants of living arrangements. An advantage of this instrument is that it is based on parental age, which is available for both cohabiting and noncohabiting children.

Italian workers can retire if they have accumulated enough years of Social Security contributions or when they reach a certain age, called *normal retirement age*. Normal retirement age and the minimum number of years of social security contributions are set by law. Before 1992, normal retirement age was 60 for most men. In 1992, a major reform of the Social Security system gradually increased normal retirement age for men from 60 in 1992 to 65 in 2002. We show that the rise in normal retirement age effectively forced some individuals to remain in the labor force longer than they would have otherwise. Because retirement typically reduces disposable income, individuals in the affected cohorts experienced an arguably exogenous increase in disposable income compared with individuals in the previous cohorts who were not affected by the reform.[13]

We use the change in father's retirement eligibility mandated by the reform as a source of variation in parents' income that is arguably exogenous to children's living arrangements.[14] Specifically, the instrument is a dummy equal to one if the father is affected by the reform, that is, if he is younger than normal retirement

13. To the extent that working lowers utility, and that without the reform some individuals in the affected cohorts would have retired earlier, the reform presumably reduced the utility of the cohorts affected.

14. The reason for focusing on the effect of the father' pension reform is that in our sample most mothers do not work (less than 30% do), and that mothers' income account for less than 25% of total household income.

age. The normal retirement age increases three times between 1989 and 2000. In particular, the instrument equals 1 for fathers younger than 60 in 1989, 1991 and in 1993, 61 in 1995, 63 in 1998 and 64 in 2000.[15] We expect the reform to reduce the probability of being retired for affected fathers, and therefore increase their incomes. In other words, we expect the first-stage coefficient on the instrument in equation 5 to be positive: $\gamma_I > 0$.

The key identification assumption is that changes in retirement age mandated by the reform are uncorrelated with other unobserved determinants of coresidence. Because our source of identification depends on changes in social security eligibility, not actual retirement decisions, it is arguably exogenous to children's living arrangement decisions and it should be orthogonal to the sources of potential endogeneity outlined in the previous section. Subsequently we show some specification checks that lend some credibility to this assumption.

Identification stems from the interaction of father's age and time, namely, the changes over time in the retirement age mandated by the law. It is important to note that the fact that the law mandated increases in the retirement age over time allows us to control for father's age dummies and year dummies. This is obviously important, because father's age is arguably an important determinant of children living arrangement decisions.[16]

In interpreting the results, it is important to note that our TSIV estimator can only include as controls (X) those variables that are available both in the children sample and in the parent sample. These include only father's age, mother's age, and year. If one makes the further assumption that individuals live in the same region as their parents (probably not a bad assumption for Italy), one can additionally control for region of residence, which acts as a proxy for the state of the local labor market and housing market. The lack of additional controls is potentially problematic. However, if the instrument is orthogonal to children's and parents' characteristics that affects living arrangements, our estimator in equation (7) is still consistent.

15. The reform established a progressive rise in normal retirement age for individuals with a minimum number of years of accumulated social security payments. The details of minimum retirement age and minimum years of social security contributions are presented in the first two columns of Table B1 in the appendix. The reform was amended in 1994. The details of these amendments are presented in columns 3 and 4 of Table B1. Column 5 shows the cutoff age for our instrument: The instrument is equal to one for individuals up to the reported age and zero otherwise. For a detailed account of the changes introduced by the 1992 social security reform see Brugiavini (1999) and Attanasio and Brugiavini (2003). We are grateful to Agar Brugiavini for having clarified some details of the Social Security system in Italy.

16. In a series of papers that use an identification strategy that is similar to ours, Bertrand, Miller, and Mullainathan (2001), Duflo (2000), and Edmonds, Mammen, and Miller (2001) exploit a special feature of the South African pension system as a source of exogenous variation in household income. Their instrument is based on differences in pension entitlement across age (and gender) groups. Their identification comes from a comparison of individuals of different ages. In contrast, our approach, which depends on changes over time in the age for pension eligibility, allows us to control for age differences. A second difference with these papers is that whereas in the South African case old-age pension recipients experienced large income increases, in Italy retirees experience income losses.

TABLE 3. The effect of pension reform on parents' income, parents sample.

	(1)	(2)	(3)	(4)
Father is affected by reform	2.227 (0.497)	2.077 (0.518)	1.861 (0.440)	2.107 (0.572)
% change	0.068	0.063	0.057	0.064
Father's age	yes	yes	yes	yes
Year dummies	yes	yes	yes	yes
Mother's age		yes	yes	yes
Region dummies × year dummies			yes	yes
Parents' education				yes
Father occupation and industry				yes

Notes: Standard errors in parentheses are clustered by father's age and year. The equation estimated is equation (5). The dependent variable is the sum of father's and mother's income. Entry in the first row is the coefficients on the instrumental variable (γ_1). Entry in the second row is the percent increase in the dependent variable for a unit increase in the instrument. The instrument is a dummy equal to 1 if the father is affected by the reform. The sample used is the parents sample. Age is a set of 5-year age dummies. Region is a set of 10 region dummies. Education is a set of five dummies: No education, primary, junior high, high school, or college and more. Industry is a set of seven dummies: manufacturing, agriculture, trade, transport, banking, services to firms and public sector, or services to families. Occupation is a set of five dummies: blue collar, white collar, manager, professional, self-employed, or entrepreneur. All regressions are weighted by sampling weights. Number of observations: 21,381.

There is no obvious reason to expect that changes in retirement age mandated by the reform are correlated with children's or parents' characteristics that affect living arrangements. It is important to note that, by using the two samples separately, we can test this assumption. In particular, we estimate the first stage, equation (5), conditioning on parents' characteristics; and the reduced form, equation (6), conditioning on children's characteristics. Below, we show that the estimated effect of the instrument do not change when parents and children characteristics are included, suggesting that the instrument is orthogonal to these characteristics. This result is important, because it lends some credibility to our assumption that the instrument is exogenous. It also indicates that the instrumental variable estimates are not affected by the failure to control for more children's and parents' characteristics.

3.3. Main Empirical Results

The effect of the reform on parental income. Table 3 reports the OLS estimates of equation (5), our first stage, estimated using the parent sample. As we explain in detail in the previous section, the instrument is a dummy equal one if the respondent is affected by the reform, that is, if the respondent is younger than normal retirement age in any given year. Because normal retirement age mandated by the reform increases over time, identification comes from the interaction of father's age and year. Throughout the paper, standard errors are clustered by father age and year.

The dependent variable is parents' total income (the sum of father's and mother's earnings and pension income). In column 1 we regress parents' income on the instrumental variable, year dummies, and 5-year father age dummies (40–44, 45–49, etc.).[17] The reform appears to have the expected effect. The estimate suggests that households where the husband was in the cohorts affected by the reform experienced an increase in income of about 2.2 million lira (or about $1,100), relative to individuals who were in cohorts not affected by the reform. This corresponds to about a 6% rise in parents' income. This is obtained by dividing the point estimates in the first row by the average value of parents' income (32,865 million lira). In column 2 we additionally control for five-year mother age dummies. Column 3 adds 10 region fixed effects interacted with year dummies. These dummies control for local labor market and the housing market conditions. Finally, in column 4 we include additional parents' controls: Dummies for father and mother's education as well dummies for father's industry and occupation of current (previous) job if employed (retired). Controlling for industry and occupation is important because the reform mainly affected employees in the private sector.

Estimates in Table 3 are stable across specifications. The inclusion of additional controls in columns 2–4 has little effect on the estimated effect of the reform on parents' income. The point estimate for the effect of the pension reform fall from about 2.2 million lira in column 1 to about 2 million lira in column 4. This corresponds to a rise in parents' income of about 6%. This finding lends some credibility to our assumption that changes in the reform are orthogonal to individual characteristics.

In Table 4 we further investigate the effect of the pension reform on each parents' employment, income and savings. Although we will not use these estimates in our TSIV estimation strategy, these results are helpful in assessing the effect of the reform and the validity of the instrument. All models in Table 4 control for the whole set of household controls included in the model in columns 4 of Table 3.

The first row reports the effect of the reform on father's retirement. The proportion of retired fathers falls by around 6 percentage points as an effect of the reform, which corresponds to a fall of about 16% in the probability of retirement. Interestingly, the point estimate of the effect of the reform on retirement is similar to the rise in father employment reported in row 2, indicating that the reform was effective at forcing some individuals to delay retirement and remain in the labor force. Row 3 indicates that as a consequence of the reform, father's income rises by about 1.7 million lira. In rows 4 to 7, we show the effect of the reform on mothers' income and employment. (Remember that we assign the instrument to households based on father's age, not mother's age. Therefore results in rows 4

17. We have also run regressions with unrestricted 1-year age dummies for both the father and the mother. Results are qualitatively similar to the ones reported below. Although the point estimates for the effect of changes in parental income on coresidence rates are generally higher than the ones reported below, standard errors are also higher.

TABLE 4. The effect of pension reform on different parents' outcomes, parents sample.

Dependent Variable	Coefficient	% change
1. Father retired	−0.063	−0.162
	(0.026)	
2. Father working	0.059	0.098
	(0.026)	
3. Father's income	1.747	0.070
	(0.566)	
4. Mother retired	−0.052	−0.244
	(0.024)	
5. Mother working	0.006	0.021
	(0.09)	
6. Mother's income	0.321	0.040
	(0.247)	
7. Parents' savings rate	−0.006	−0.028
	(0.009)	
8. Parents' income – Controls for father's early retirement	2.122	0.064
	(0.586)	

Notes: Standard errors in parentheses clustered by father age and year. Entries in the first column are the coefficients on the instrumental variable (γ_1), which are converted in percent changes in column 2. The instrument is a dummy equal to 1 if the father is affected by the reform. The sample used is the parents sample. All specifications control for year dummies, father's age, mother's age, region dummies interacted with year dummies, parents' education, and father's occupation and industry. Age is a set of 5-year age dummies. Region is a set of 10 region dummies. Education is a set of five dummies: No education, primary, junior high, high school or college and more. Industry is a set of seven dummies: manufacturing, agriculture, trade, transport, banking, services to firms and public sector, or services to families. Occupation is a set of five dummies: blue collar, white collar, manager, professional, self-employed, or entrepreneur. Savings rates are computed as the ratio of total household disposable income minus consumption over total disposable income. All regressions are weighted by sampling weights.

to 6 are to be interpreted as the effect of changes in fathers' mandated retirement age on mothers' labor supply.) We find some slight evidence of complementarity between father's and mother's consumption of leisure. The reform appears to have a small (but insignificant) negative effect on mothers' retirement status. It also appears to have a small positive effect on mothers' employment, mothers' earnings and mothers' income.

Findings in rows 1 to 6 lend support to the idea that changes in parents' disposable income documented in Table 3 are generated by increased fathers' employment induced by the reform.

We then test whether the reform affected savings rates. Effectively, the reform affected not only normal retirement age but also the rules for the computation of pension income, essentially reducing replacement ratios. For this reason, it is possible that the cohorts affected by the reform discounted the loss in future pension income which resulted from the reform. In this case the effect of the reform on cohabitation rates could mask changes in parents' permanent income and the interpretation of our results would be affected.[18] We do not expect this to be

18. For example, it is in theory possible that increases in cohabitation rates associated with the reform are the result of a decrease in parents' permanent income and the inability to buy a home for their children.

a major problem for our estimates. The effect of this change for the cohorts at the margin of retirement was almost insignificant, since the changes in replacement ratios occurred very gradually over time (Attanasio and Brugiavini 2003). However, to see whether this is a problem in our sample, we regress parents' savings rates on the instrument and all the controls. Savings rates are obtained as in Attanasio and Brugiavini as the difference between total household disposable income minus total consumption over total disposable income. The estimate in row 7 is statistically not distinguishable from zero. We conclude that there is little evidence that savings rates of the cohorts of fathers at the margin of retirement were affected by the reform.

One feature of the reform is that not only did it affect the eligibility rules for normal retirement but also those determining early retirement. As a last check for our results, in row 8 of Table 4 we report the effect of the reform on parents' income once we include in the regressions a dummy variable that controls for changes in early retirement.[19] Results are essentially unchanged with respect to Table 3 (column 4), implying that the omission of this additional institutional feature is unlikely to bias our results.

Based on Tables 3 and 4, we conclude that the reform was effective in forcing some fathers to remain in the labor force and in increasing their incomes. Mothers' labor supply did not adjust, and as a consequence, household income increased.

The effect of the reform on probability of coresidence. We now turn to the children sample, and estimate the effect of the reform on the probability of coresidence (equation (6)). In column 1 of Table 5 we report the results of an OLS regression where the dependent variable is a dummy equal to one if the child lives with his parents. The right-hand side variable is the instrument. We additionally control for father age dummies and year dummies. The estimates suggest that the reform raised cohabitation rates by about 7.5 percentage points. Estimates are significant at conventional significance levels. In columns 2 and 3, we add controls for mother's age and the interaction of region dummies and year dummies, respectively. Inclusion of these additional controls has little effect on our estimates, suggesting that these are uncorrelated with the reform. In column 4 we also control for child's age. The point estimate remains virtually unchanged. The second

19. Changes in early retirement affected differentially self-employed workers and public and private employees. As in the case of normal retirement, requirements were also imposed on the years of contributed Social Security payments. Again, we ignore this latter requirement because information on Social Security contributions in the SHIW data is only available from 1995 onward and—most important—only in the parents' sample. Based on the institutional features of the reform, we define a dummy for normal retirement that takes value 1 for individuals aged less than 52 from 1989 to 1995, less than 54 in 1998, and less than 55 in 2000 if private employees. The dummy takes a value 1 for individuals aged less than 52 from 1989 to 1995, less than 53 in 1998, and less than 54 in 2000 if public employees; and a value 1 for individuals aged less than 56 from 1989 to 1995, and less than 57 in 1998 and 2000 if self-employed.

TABLE 5. The effect of pension reform on sons' coresidence, OLS estimates, children sample.

	(1)	(2)	(3)	(4)
Father is affected by reform	0.076	0.073	0.073	0.074
	(0.019)	(0.017)	(0.020)	(0.018)
% change	0.088	0.085	0.084	0.086
Father's age	yes	yes	yes	yes
Year dummies	yes	yes	yes	yes
Mother's age		yes	yes	yes
Region dummies × year dummies			yes	yes
Child's age				yes

Notes: Standard errors in parentheses clustered by father age and year. The equation estimated is equation (6). The dependent variable is a dummy equal to 1 if the respondent lives with his parents. Entry in the first row is the coefficient on the instrumental variable (θ_1). Entry in the second row is the percent increase in cohabitation for a unit increase in the instrument. The instrument is a dummy equal to 1 if the father is affected by the reform. The sample used is the children sample. Age is a set of 5-year age dummies. Region is a set of 10 region dummies. All regressions are weighted by sampling weights. Number of observations: 11,782.

row of Table 5 shows that the reform was associated to rise in cohabitation rates of more than 8% (this is obtained by dividing the point estimates in the first row by the average level of coresidence, 0.86).

Two-sample instrumental variable estimates. We are now in a position to compute our TSIV estimates defined in equation (7). The TSIV estimates of the effect of parents' income on the probability that a child lives with his parents, β_1, are reported in Table 6. The model in the first column controls for father's age and year dummies. Recall from equation (7) that the TSIV coefficient is the ratio of the reduced form estimates in Table 5 over the first stage estimates in Table 3. For example, the coefficient in column 1 of Table 6 is the ratio of the coefficient in column 1 of Table 5 over the coefficient in column 1 of Table 3, and is equal to $0.076/2,227 = 0.034$. Standard errors are adjusted to account for the fact that data are from two samples.[20] Columns 2 and 3 add controls for mother's age and region dummies interacted with year dummies, respectively. Estimates are not particularly sensitive to these additional controls.

Increases in parental income seem to have a significant effect on the probability of coresidence, and the magnitude of the effect appears to be nontrivial. The estimate in column 3, for example, when all the available controls are

20. In order to compute the standard errors we have used the procedure suggested by Jappelli, Pischke, and Souleles (1998). First we estimate the first stage regression of Y_P on Z using the parents sample (all regressions are conditional on X). Second, we regress H on the estimated value of Y_P from the first stage regression using the children sample. Let us denote the variance of the estimator by var_1. Third we predict the value of H in the parents sample using this last equation. Finally, we regress this predicted value of H on the predicted value of Y_P from the first stage equation using the parents sample. Let us denote the variance of this estimator by var_2. We compute the standard error of the TSIV estimator as the square root of $var_1 + var_2$. Standard errors in all the regressions are clustered by father's age and time.

TABLE 6. The effect of parents' income on sons' coresidence, two-sample instrumental variable estimates.

	(1)	(2)	(3)
Parents' income	0.034	0.035	0.039
	(0.011)	(0.012)	(0.014)
Partial elasticity	1.300	1.352	1.487
Father's age	yes	yes	yes
Year dummies	yes	yes	yes
Mother's age		yes	yes
Region dummies × year dummies			yes

Notes: The equation estimated is equation (7). Entries in the first row are the TSIV estimates of the effect parents' income on children's coresidence. Entries are the ratio of the coefficients in Table 5 over the coefficients in Table 3. Partial elasticity is the percent increase in probability of cohabitation for a percentage increase in parents' income. See Angrist and Krueger (1992) for the definition of TSIV. TSIV standard errors are obtained following the procedure described in Jappelli Pischke, and Souleles (1998).

included, indicates that an extra 1 million lira of parents' income (about $500) raises the probability of cohabitation by about 3.9 percentage points. The second row in the table reports the implied partial elasticity, defined as the percent increase in probability of cohabitation for a percentage increase in parents' income. Average parents' income is 32.86 million liras, and average cohabitation rate is 0.86. According to the estimate in column 1, a 10% increase in parents' income (3.28 million liras) would result in an increase in the probability of cohabitation of around 12.80 percentage points (this is 3.28×0.039), that is, a rise in cohabitation rates of around 15%, with an implied elasticity of around 1.5.[21]

Based on the evidence in Tables 3 to 6, we conclude that the 1992 reform of Italian Social Security had a significant impact on parents' disposable income. It also increased cohabitation rates between parents and children. A 1 million lira increase in parents' income raises coresidence rates by 3.5–4 percentage points. The reform appears to be largely uncorrelated with other observable determinants of coresidence rates.

To put the estimated effect in perspective, it is useful to calculate how much our estimate would predict cohabitation to rise by in the last 10 years given the increase in parental income, and how this predicted rise compares with the actual rise. In the period under consideration (from 1989 to 2000), total parental income increased in real terms (at 1995 prices) by about 78,000 liras per year. Based on our estimates in Table 6, this increase in income would imply an annual increase of at least 0.26 percentage points in cohabitation rates. The actual increase in cohabitation rates over that period was 0.42 percentage points a year. The predicted increase is then 62% of the actual increase.

21. As an additional check for our results, We have also reweighted data from the parents sample by the probability that a father has a child in the age group 18–30 and used this reweighted sample to compute our first stage estimates. This procedure leads essentially to identical results.

3.4. Robustness Checks

In this section we present results from several additional specifications intended to probe the robustness of our estimates and to verify the validity of our identifying assumption.

One concern is that father's retirement may directly affect children propensity to cohabit if children's disutility of living with their parents increases when parents are retired. Retired parents are likely to spend most of their time at home, and this could reduce the children's propensity to live there. If this was true, our instrument would be invalid. To assess whether this is a serious problem, we re-estimated equation (7) including only working children. Children who have a job are likely to spend most working hours outside the home. For working children, it should make little difference whether their parents are at home during working hours or at work. In row 1 of Table 7, we find that estimates of equation (7) based on the sample of working children are similar to the ones for the whole sample. This suggests that the possibility that children increased aversion to live in their parents' home when their parents retire does not explain our results.[22,23]

A second concern regards the effect of changes in parents' income on children's schooling and labor supply. It is in theory possible that the rise in parental income induced by the reform makes it possible for children to attend college, by relaxing parents' liquidity constraints. This is potentially problematic because many Italian children live at home while attending college. The rise in cohabitation rates that we uncovered could in theory be the by-product of a higher probability of school enrollment. Although this would not change our results, this would weaken our interpretation of Italian children staying at home in exchange for some financial transfers on the part of their parents. In row 2 of Table 7 we report estimate of the effect of changes in parents' income on children's school enrollment. We find little evidence that the reform affected enrollment.

Rows 3 and rows 4 report the effect of changes in parents' income on children's labor supply. We find a negative but not highly significant effect of changes in parents' income on children's employment. Results are more precise when we consider earnings: We find that a 1 million lira rise in parents' income is associated to a fall in children's income of about 1.3 million lira, suggesting that one way

22. Obviously there are potential problems with this strategy because child's labor force status (our conditioning variable) is likely to be endogenous to parents' income (and the transfers children receive from their parents).

23. This result also helps us to rule out an additional alternative interpretation for our results. One might be concerned that older children may provide child-care for their younger siblings. This could pose a problem if father's retirement makes it possible to relieve the older children from their child-care responsibilities, therefore making it easier for the older children to leave the parental household. Because working children are unlikely to be able to look after their younger siblings, this result suggests that sibling's care in an unlikely explanation for our results in Table 6.

TABLE 7. Robustness checks. The effect of parents' income on children's behavior, TSIV estimates.

Dependent variable	Coefficient	Elasticity
1. Son lives with parents (sample includes only working children)	0.049 (0.019)	2.177
2. Son enrolled in school	0.005 (0.012)	0.570
3. Son works	−0.036 (0.026)	−2.385
4. Son's earnings	−1.309 (0.584)	−4.541
5. Son lives with parents (controls for parents' birth cohort)	0.033 (0.021)	1.241
6. Son lives with parents (controls for changes in pension indexation)	0.053 (0.015)	2.007
7. Son lives with parents (north)	0.078 (0.053)	3.250
8. Son lives with parents (south)	0.015 (0.006)	0.468
9. Son lives with parents (1993–2000)	0.035 (0.023)	1.314
10. Daughter lives with parents (1993–2000)	−0.058 (0.052)	−2.572
11. Son lives with parents (controls for parents' education; 1993–2000)	0.040 (0.040)	1.499

Notes: Entries in the first row are the TSIV estimates of the effect parents' income on children's coresidence. All specifications control for year dummies, father's age, mother's age, and region dummies interacted with year dummies. Standard errors are in parentheses.

children spend the transfer they receive from their parents is through increased leisure consumption.[24]

Row 5 reports the effect of changes in parents' income on cohabitation rates with additional controls for father's birth cohort. In particular we include in the regression a third-order polynomial in the father's year of birth. The goal is to control for unobserved heterogeneity across cohorts. Giuliano (2002), for example, argues that more liberal attitudes of the parents who experienced the 1968 sexual revolution might explain the increase in coresidence rates among young Italians. It is also possible that different cohorts of fathers differ in their permanent level of income or wealth, thus inducing spurious correlation. The point estimate falls

24. Ideally, if we observed these parental transfers, we could directly test the notion that parents "bribe" their children in exchange for cohabitation. Unfortunately our data do not provide information on consumption of excludable goods by children. Purely based on casual observation, though, we speculate that parental transfers might take the form of higher children's consumption of clothing, leisurely activities (i.e., travel and "going out"), and durable and semidurable goods (cars, motorcycles, mobile phones). External observers are often surprised to note how young Italian men seem to have high rates of conspicuous consumption despite the remarkably high rate of youth unemployment. Indeed, our analysis suggests that low labor supply and high consumption levels could be two sides of the same coin, that is, the effect of parental transfers in exchange for children coresidence.

slightly (from about 0.039 in Table 6 to 0.033 in Table 7), although the precision of the estimate declines due to an increase in the standard erros of the estimates of the first-stage equation.

Overall, we conclude that differences across cohorts are unlikely to be major sources of bias.

Next, we check whether changes in pension indexation that took place over the same period affected the income profile of workers relative to retirees. The concern is that such changes could in theory lead to biased Two Sample Least Squares (TSLS) estimates. In row 6, we allow the income profiles for retirees and workers to differ across years. Results are very similar to the ones in Table 6: The point estimate is 0.053 (0.015).

In Rows 7 and 8, we test whether there is any substantial difference in the effect of parental income of cohabitation rates between the two macro areas of the country: The North, characterized by high income and low youth unemployment, and the South, characterized by lower income and higher youth unemployment.[25] We find that the effect of parental income on cohabitation is positive in both areas, although the estimated coefficient is higher in the North than in the South (the point estimates are, respectively, 0.078 and 0.015).

The remaining estimates are based on the subsample of children (and their parents) observed between 1993 and 2000. Starting in 1993, the SHIW provides information on parents' age for both the head of household and the their spouse. This allows us to compute the effect of changes in parents' income on women living arrangements. On the other hand, by dropping data for 1989 and 1991, we lose about 35% of the sample. The number of observations for sons falls from 11,782 to 7,438, and the number of observations for parents falls from 21,381 to 13,350.

Row 9 reports the results for men in the period 1993–2000 for a model similar to the one in column 3 of Table 6. The point estimate remains virtually unchanged, although there is a substantial loss in precision. Row 10 reports results for women. The estimate is negative but not statistically significant. The imprecision of the estimate precludes firm conclusions.

We note that it is in theory possible that the determinants of daughters' living arrangements differ from the ones of their male siblings. First, women's living arrangements are highly correlated with marital status. In most cases, Italian women live with their parents until they marry and then leave their parental home. This makes it unlikely that the effect of parental income would be as large for women as it is for men. Second, it is difficult to interpret this estimate, because a woman's marriage market potential may be affected by parental income.

25. North includes: Piedmont, Val d'Aosta, Liguria, Lombardy, Trentino Alto Adige, Veneto, Friuli Venezia Giulia, Emilia Romagna, Tuscany, Umbria, Marches, and Latium. South includes: Campania, Abruzzi, Molise, Apulia, Basilicata, Calabria, Sicily, and Sardinia.

Finally, we turn to the question of whether failure to control for parental education is an important source of bias. We have already shown above that parental education is uncorrelated with the instrument in the parents' income equation. Although this finding is reassuring, it does not rule out some correlation in the reduced form equation. The 1993–2000 subsample provides information on the level of education of the parents' of the head's and the head's spouse if they do not live with their parents. We use this information to estimate equation (7) with additional controls for father's and mother's education.[26] The point estimate in row 11 is similar to the one in Table 6, although much less precise, due to the smaller sample size.[27]

Taken together, results in Table 7 indicate that the reform is unlikely to have had an effect on cohabitation rates for reasons other than the effect on parents' income.

4. Happiness and Coresidence

Our model in Section 2 is based on the assumption that parents draw utility from their children's presence at home. The evidence in Section 3.3 is generally consistent with this assumption. Here we present some additional (suggestive) evidence that indicates that this assumption may be reasonable for Italy, although not necessarily for other countries.

In the model of Section 2, parents have all the bargaining power. Any surplus will then accrue to parents. One example of such surplus is children's housing cost. In equilibrium, parents end up appropriating what is left of the surplus after the bribe is paid. This implies that parents should be better off if children cohabit. On the other hand, in equilibrium children should be indifferent between living at home and living on their own.

Evidence suggests that both these two results might be true for Italy, but not necessarily for other countries, particularly the U.S. Table 8 shows the results of an OLS regression of a measure of parental happiness on a dummy variable indicating whether at least one child lives with the parents and a set of covariates. The coefficient on the cohabitation dummy is reported for the same set of countries as in Table 1. Data are from the World Values Survey 1981–1984, which includes the question: "Taking all things together, would you say you are: Not

26. Caution must be used in interpreting these results, because the information on parental education is not ideal. The question refers to the level of education achieved by the parents' when they were of the same age as the child at the time of the interview.

27. The data also provide information on the sector and occupation of the parents. However, this information refers again to the occupation and industry of employment of parents at the time they were of the same age as their children currently are. We prefer not to include this additional controls because these are likely to be severely error ridden measures of parents' current occupation and industry of employment.

TABLE 8. Coresidence and happiness.

	Basic Controls	Additional Controls
Parents		
France	−1.089	−2.074
	(1.301)	(1.286)
Great Britain	−1.490	−1.935
	(1.281)	(1.257)
West Germany	−0.345	−1.744
	(1.222)	(1.203)
Italy	5.420	4.084
	(1.590)	(1.566)
Spain	3.123	1.892
	(1.549)	(1.520)
U.S.	−3.167	−2.706
	(1.527)	(1.498)
Portugal	1.805	1.287
	(3.106)	(3.044)
Children		
France	1.454	2.635
	(2.195)	(2.319)
Great Britain	−5.031	−1.851
	(2.053)	(2.188)
West Germany	−4.557	−2.953
	(2.086)	(2.178)
Italy	−6.312	−3.278
	(2.139)	(2.262)
Spain	−3.886	−0.688
	(2.322)	(2.433)
U.S.	−1.181	1.800
	(2.616)	(2.699)
Portugal	−6.116	−2.830
	(4.348)	(4.407)

Notes: Standard errors in parentheses. The dependent variable is a categorical variable which takes four values between 0 (not at all happy) and 1 (very happy). The coefficient for parents is the coefficient($\times 100$) on a dummy variable equal to 1 if at least one child lives with the parents. The coefficient for children is the coefficient on a dummy equal 1 if living with parents. Basic controls include family income, gender, age and age squared, and country-specific dummies. Additional controls include a full set of dummies for marital status (married, cohabiting, single, divorced or separated, and widowed), dummies for their employment status (employed, unemployed, housewife, student, pensioner, or other inactive) and health status. Regressions are weighted by sampling weights. Sample includes men aged 40–74 and women aged 37–71 with children. Number of parent observations: 6,021. Data: World Values Survey, 1981–1984. Number of children observations: 2,247.

at all happy, Not very happy, Quite happy, Very happy." We classify the possible answers into equally spaced values between zero and one, with "Not at all Happy" being zero and "Very Happy" being one. In the first column of the table we report the results of a regression with basic controls including family income, parents' gender, parents' age, and age squared plus country-specific dummies. In the second column we include additional controls, namely a full set of dummies for parents' marital status, dummies for their employment status, and a health status variable.

According to the estimates in column 1, parents in Italy and Spain seem to be significantly happier if their children live with them, whereas the opposite is true in the United States. When we introduce further controls in column 2, the coefficient becomes insignificant for Spain, but remains significant and positive for Italy.

Evidence from the same survey also supports our claim that Italian children, all things equal, are indifferent between living at home and living independently. At the bottom of Table 8 we present results from a separate regression where the dependent variable is children's happiness. The sample includes males aged 18 to 30, consistent with the sample that we use in the rest of the empirical analysis. The variable of interest here is a dummy variable for whether the child lives with his parents. In all countries, including Italy, the effect of cohabitation is negative, although it tends to become statistically insignificant when all controls are included.

Table 8 indicates that children are indifferent between living at home and living on their own in all countries included in this analysis while parents draw some utility from cohabitation in Italy but not in the rest of the countries in our sample. However, these findings need to be interpreted as merely suggestive. First, their interpretation is not terribly clean, because cohabitation is endogenous. Second, Table 8 says little about the relationship between parental income and cohabitation.

5. Interpretation

Our findings in Section 3 indicate that an arguably exogenous increase in parental income result in an increase in the probability of cohabitation. Cohabitation is also associated with a decrease in children's labor supply. Our findings in Section 4 suggest that Italian parents are better off if children cohabit and Italian children are indifferent between living at home and living on their own. Taken together, these results seem consistent with the notion that Italian parents have a preference for co-habitation, and are willing to transfer resources to their children to keep them at home. Although our empirical results are consistent with the model presented in Section 2, other interpretations are also possible. In this section, we discuss some alternative interpretations of the empirical evidence.

First, recent work by Becker et al. (2003) indicates that job insecurity is an important determinant of living arrangements for Italian men. Although the evidence presented by Becker et al. is convincing, the empirical evidence in this paper captures something different. Job insecurity could in theory explain our results if the increases in average parents' income were associated with decreases in job insecurity. In this case, a fall in job insecurity among the cohort of parents affected by the reform would make it more likely that their children stay at home. If this were the case, failing to account for changes in job insecurity would lead us to obtain biased estimates. But our estimates are identified by changes in mean income due to changes in retirement eligibility, and the changes in retirement eligibility are unlikely to be negatively correlated with job insecurity. The reason is that while changes in retirement eligibility affect mean income of some cohorts of fathers, there is no reason to expect that changes in retirement eligibility reduce job insecurity. If anything, the cohorts of parents affected by the reform are facing more income uncertainty, relative to the cohorts of fathers not affected by the reform.[28] In other words, although we think that this job insecurity is an important factor affecting coresidence rates in Italy, job insecurity is unlikely to be an important explanation for our results.

Similarly, it is possible that parents are happier about cohabitation (as in Table 8) not because they like cohabitation per se, but because they are altruistic and job insecurity is a bad for children. We cannot completely rule out this interpretation. We point out, however, that if this was the case, children should be happier when they live with their parents. This does not appear to be the case from Table 8.

A second possible alternative explanation of our results is that children can exert some influence over their parents and extract resources from them. If cohabitation makes children more productive in extracting such resources from their parents, this model can in theory explain the empirical evidence in Section 3. As parental income increases, the potential gain from cohabitation increases, and children will be more likely to move in with their parents. On the other hand, this model would predict that parents are less happy if they live with their children, which is not consistent with the evidence in Section 4.[29]

One issue on which we remain agnostic regards cross-country differences in cohabitation rates. Income has risen around the world. Why don't American children live at home longer? Why is Italy such an outlier? We don't have a definitive answer. We speculate that this might be due to differences in preferences and culture. We note that the evidence in Table 8 is consistent with this view. However, as discussed above, the evidence in Table 8 is very indirect and precludes firm conclusions.

28. Whereas social security income is fixed, private sector earnings do have some variability (although this variability is quite small for older workers in Italy).

29. The interpretation of our results crucially depends on the assumption that parents derive some utility from cohabitation. A third alternative story is that parents derive disutility from cohabitation but they are altruistic towards their offspring. One unappealing feature of this story is that it is not clear why children need to cohabit in order to benefit from parental altruism. If parents are altruistic, children will benefit irrespective of their living arrangements.

6. Conclusions

Rates of cohabitation between parents and children are much higher in Southern Europe than in the U.S. and in Northern Europe. In this paper we try to address one of the forces behind such a high rate of cohabitation by focusing on the effect of parental resources on children's coresidence choices in Italy. It is important to notice that, although we believe that our analysis has some potential to help in rationalizing the dramatic differences in cohabitation rates across different countries, in this paper we cannot affirm whether and to what extent our analysis is able to shed light on such differences.

We use Bank of Italy micro data to empirically estimate the effect of parental income on children living arrangements decisions. The key econometric issue is the endogeneity of parental income. To obtain consistent estimates, we use a TSIV estimator which exploits the 1992 reform of Social Security to generate an arguably exogenous source of variation in parents' income.

We show that this reform is associated with a significant rise in father's disposable income and with a rise in the probability that children live with their parents. We interpret this finding as evidence that parental income affects positively children's living arrangements. We find that a one million lira rise in parents' annual income (approximately $500) increases the probability that children live with their parents by about 3.5–3.9 percentage points, which corresponds to an elasticity of coresidence with respect to parental income of about 1.3 to 1.5. Based on our estimates, we calculate that the increase in parents' income between 1989 and 2000 is responsible for a substantial fraction (between 60% and 70%) of the increase in cohabitation rates experienced by Italian young men.

Although we cannot completely rule out the possibility that unobserved heterogeneity may play a role in explaining the relationship between parental income and living arrangements, we show that our instrument is orthogonal to several children's and parents' observable characteristics. This finding lends some credibility to the assumption that our instrument is exogenous. We also try to address the concern that retirement might affect coresidence for reasons that are unrelated to changes in parental disposable income; we find no evidence of our reform affecting cohabitation rates because of permanent changes in parents' income, the probability that children are enrolled in school, the disutility of parents' leisure on the part of children or the fact that children's and parents' leisure are substitutes in the household utility function.

In the light of our theoretical model, we interpret our results as being consistent with the idea that cohabitation is a normal good for Italian parents and that parental preferences might contribute to explain the remarkably high rate of cohabitation between Italian children and their parents. Evidence that children cut their labor supply in response to exogenous rises in parents' income is consistent with our hypothesis that children indirectly benefit from a rise in their parents income and trade some of their independence in exchange for higher consumption. Additional evidence from the World Value Survey lends some additional support to this conclusion. Our empirical results help rationalize some puzzling facts about the Italian economy and society: low youth employment, high coresidence rates, deferred marriage, and low fertility. Implicitly we suggest that the structure of welfare entitles parents to some bargaining power towards their offspring and this has potentially some undesirable policy implications.

However, at the end of the paper, we also caution the reader that alternative explanations of our findings cannot be entirely ruled out. For example, we cannot completely rule out the possibility that cohabitation is undesirable for Italian parents, but children prefer to live with richer parents because of the potential gains from such cohabitation.

Appendix A

In this Appendix, we consider the case where parents are altruistic. As in Section 2, we assume that parents derive utility from cohabitation. But unlike Section 2, we also assume that parents' utility depends on their children's utility. We show that when parents are altruists and internalize their children's distaste for cohabitation, a sufficiently high desire for coresidence will generate the same result obtained in Section 2: A rise in parental income will be associated with an increase in cohabitation rates.

Parents maximize a linear combination of their own private utility and their child's utility with weights $(1-\rho)$ and $\rho (0 \leqslant \rho \leqslant 1)$, where ρ reflects the degree of parental altruism. The parents' problem can now be written as

$$\begin{aligned}
\text{Max} \quad V_P(C_P, C_K, H) &= (1-\rho)U_P(C_P, H) + \rho U_K(C_K, H) \\
&= (1-\rho)[\log(C_P) + H\log(a_P)] \\
&\quad + \rho[\log(C_K) + H\log(a_K)] \\
\text{Subj. to:} \quad C_P &= Y_P - b_0(1-H) + b_1 H \\
C_K &= Y_K + (b_0 - R)(1-H) + b_1 H, \qquad \text{(A.1)}
\end{aligned}$$

where the notation is the same as above and b_0 is the altruistic transfer to non-cohabiting children. The children's utility function remains unchanged and it is described in (2). As in Section 2, we assume that parents set b_1 so that children will be indifferent between living at home or on their own. Formally, they will set a value of b_1 such that $U_K(b_1, 1) = U_K(b_0, 0)$. But unlike the model in Section 2, children now know that they will receive some compensation b_0 even if they do not cooperate. This is because parents are altruistic, and will transfer income to their children even if they live on their own. The level of this transfer is unilaterally set by parents who maximize their own utility in (4) conditional on $H = 0$. One can check that the altruistic transfer to non-cohabiting children is $b_0^* = \rho Y_P - (1-\rho)(Y_K - R)$. From the indifference condition for children, the optimal bribe to cohabiting children will then be $b_1^* = \rho(Y_P + Y_K - R)/a_K - Y_K$. Conditional on b_1^* and b_0^*, parents will be willing to bribe their children into staying at home if the utility they derive from cohabitation is higher than the utility from separation, $V_P(Y_P - b_1^*, b_1^*, 1) \geqslant V_P(Y_P - b_0^*, b_0^*, 0)$:

$$Pr(H = 1) = Pr(A_0 Y_P \geqslant A_1 R - A_0 Y_K), \tag{A.2}$$

where $A_0 = [a_K(a_P - 1) - \rho(a_P - a_K)]$ and $A_1 = (a_K - \rho a_K - \rho a_P)$.

It is easy to see that a rise in parents' income Y_P is associated with a rise in the probability of cohabitation ($A_0 \geqslant 0$) if $a_P > a_P^* > 1 > a_K$, where $a_P^* = a_K(1-\rho)/(a_K - \rho)$.[30] A necessary and sufficient condition for a rise in parents' income to generate a rise in the rate of cohabitation is that parents have a sufficiently high direct utility of cohabitation.[31] Notice that because parents are altruistic, they internalize their children's disutility of cohabitation. In equilibrium, if a rise in parental income is associated with a rise in the rate of coresidence, parents' private marginal utility of cohabitation is strictly positive ($a_P^* > 1$).

Relative to the non-altruistic case described in Section 2, the model with altruism generates less unambiguous conclusions. With altruism, higher parental income may or may not result in higher probability of coresidence, depending on the utility function. If utility of cohabitation is sufficiently high, a rise in parental income is associated with a rise in coresidence.

30. Notice that this equation only holds for $a_K > \rho$. For values of a_K below this threshold children of richer parents live away from home irrespective of parental income. In this case there is no bribe that is sufficient to create an incentive for children to stay at home as parents' income rises.

31. One can check that for $a_P < a_K^*$, that is, if altruistic parents have stronger preferences for privacy than their children (once the latter has been adjusted for parental altruism) then cohabitation will fall as parental income rises.

Appendix B

TABLE B.1. Requirements for men's normal retirement (*pensione di vecchiaia*) in Italy.

	1992 reform		1994 up-date of 1992 reform		Instruments = 1 if age less than
	(1)	(2)	(3)	(4)	(5)
Time period	Age	Years of social security payments	Age	Years of social security payments	
Before 1993	60	15	60	15	60
1/1/93–31/12/93	60	16	60	16	60
1/1/94–31/12/94	61	16	61	16	61
1/1/95–30/6/95	61	17	61	17	61
1/7/95–31/12/95	61	17	62	17	61
1/1/96–31/12/96	62	17	62	17	62
1/1/97–31/12/97	62	18	63	18	62
1/1/98–30/6/98	63	18	63	18	63
1/7/98–31/12/98	63	18	64	18	63
1/1/99–31/12/99	63	19	64	19	64
1/1/00–31/12/00	64	19	65	19	64

Notes: The table reports the details of the Italian 1992 pension reform and its 1994 update. Columns 1 and 2 report the minimum age and the minimum years of contributed social security payments to access normal retirement over different time periods as established by the 1992 reform. Columns 3 and 4 report the same information as implied by the 1994 update of the 1992 reform. Column 5 reports the cutoff age for our instrumental variable. This variables takes the value 1 if individuals are aged less than the cutoff age and zero otherwise.

References

Altonji, Joseph, Fumio Hayashi, and Laurence J. Kotlikoff (1997). "Parental Altruism and Inter Vivos Transfers: Theory and Evidence." *Journal of Political Economy*, 105, 1121–1166.

Angrist, Josh, and Alan Krueger (1992). "The Effect of Age at School Entry on Educational Attainment: An Application of Instrumental Variables with Moments from Two Sample." *Journal of American Statistical Association*, 87, 328–336.

Attanasio, Orazio, and Agar Brugiavini (2003). "Social Security and Households' Savings." *Quarterly Journal of Economics*, 118, 1075–1119.

Becker, Sascha, Samuel Bentolila, Ana Fernandes, and Andrea Ichino (2003). "Parental Job Insecurity and Children's Emancipation: The Italian Puzzle." Working Paper, IUE.

Bertrand, Marianne, Douglas Miller, and Sendhil Mullainathan (2001). "Public Policy and Extended Families: Evidence from South Africa." Working Paper, Princeton University.

Bitler, Marianne, Jonah Gelbach, and Hilary Hoynes (2001). "The Impact of Welfare Reform on Living Arrangements." Working Paper, UC Davis.

Brugiavini, Agar (1999). "Social Security and Retirement in Italy." In *Social Security and Retirement Around the World*. NBER Conference Report series, edited by J. Gruber and D. Wise. University of Chicago Press.

Card, David, and Thomas Lemieux (2000). "Adapting to Circumstances: The Evolution of Work, School, and Living Arrangements Among North American Youth." In *Youth Employment and Joblessness in Advanced Countries*, edited by D. Blanchflower and R. Freeman. University of Chicago Press.

Costa, Dora (1997). “Displacing the Family: Union Army Pensions and Elderly Living Arrangements.” *Journal of Political Economy*, 105, 1269–1291.
Cox, David (1987). “Motives for Private Income Transfers,” *Journal of Political Economy*, 95, 508–546.
Del Boca, Daniela (1997). “Intrahousehold Distribution among Italian Families.” In *The Economics of the Family*, edited by I. Parsson and C. Jonung. Routledge.
Del Boca, Daniela (2002). “The Effect of Child Care and Part Time Opportunities on Participation and Fertility Decisions in Italy.” *Journal of Population Economics*, 15, 549–573.
Duflo, Esther (2000). “Grandmothers and Granddaughters: Old Age Pension and Intrahousehold Allocation in South Africa.” NBER Working Paper, No. 8061.
Edmonds, Eric, Kristin Mammen, and Douglas Miller (2001). “Rearranging the Family? Household Composition Responses to Large Pension Receipts.” Working Paper, Dartmouth College.
Fogli, Alessandra (2000). “Endogenous Labor Market Rigidities and Family Ties.” Working Paper, NYU.
Giuliano, Paola (2002). “The Peter Pan Paradox: Why Mediterranean Youth Stay at Home, Do Not Marry, Do Not Have Children (and May Not Work).” Working Paper, UC Berkeley.
Jappelli, Tullio, Jorn-Steffen Pischke, and Nicholas S. Souleles (1998). “Testing For Liquidity Constraints In Euler Equations With Complementary Data Sources.” *The Review of Economics and Statistics*, 80, 251–262.
McElroy, Marjorie B. (1985). “The Joint Determination of Household Membership and Market Work: The Case of Young Men.” *Journal of Labor Economics*, 3, 293–316.
McGarry, Kathleen, and Robert F. Shoeni (2000). “Social Security, Economic Growth and the Rise in Elderly Widows’ Independence in the Twentieth Century.” *Demography*, 37, 221–236.
Rosenzweig, Mark, and Kenneth I. Wolpin (1993). “Intergenerational Support and the Life-Cycle Incomes of Young Men and Their Parents: Human Capital Investments, Coresidence, and Intergenerational Financial Transfers.” *Journal of Labor Economics*, 11, 84–112.
Rosenzweig, Mark, and Kenneth Wolpin (1994). “Parental and Public Transfers to Young Women and Their Children.” *The American Economic Review*, 84, 1195–1212.
Ruiz-Castillo, Javier, and Maite Martínez-Granado (2002). “The Decisions by Spanish Youth.” *Journal of Population Economics*, 15, 305–330.

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5. Piero Cipollone, Alfonso Rosolia. 2007. Social Interactions in High School: Lessons from an Earthquake. *American Economic Review* **97**:3, 948-965. [CrossRef]