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Job Insecurity and Children's Emancipation

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ABSTRACT

Job Insecurity and Children's Emancipation*

The age at which children leave the parental home differs considerably across countries. We present a theoretical model predicting that higher job security of parents and lower job security of children may delay emancipation. We then provide aggregate evidence which supports this hypothesis for 12 European countries. We also give microeconometric evidence for Italy, the single country for which we have access to household-specific information on job security and coresidence. It is a very interesting case to study since, in the late 1990s, approximately 75% of young Italians aged 18 to 35 were living at home and they had only a 4% probability of emancipation in the subsequent 3 years. We show that this probability would have increased by 4 to 10 percentage points if their fathers had gone from having a fully secure job to becoming unemployed for sure.

JEL Classification: J1, J2

Keywords: emancipation, job security, option value

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1 Introduction

The age at which children decide to leave the parental home differs considerably across countries. For instance, while in 2002 coresidence rates for males aged 25 to 29 years old were 20% in the Netherlands, 21% in the UK, and 22% in Ireland, in Italy this rate was as high as 73% (reaching a staggering 91% for males aged 20 to 24 years old). Close to this Italian record are other southern European countries like Greece and Spain, with coresidence rates of 70% and 67%, respectively, but also a Nordic country like Finland, where it was also equal to 73%. Such huge disparities across countries, which are observed in the analogous figures for females as well, call for an explanation.

The economic literature on this issue has focused on variables such as parents' and children's incomes, including unemployment, and on the sharing of public goods, such as housing, through coresidence. Empirical findings for the US and the UK indicate that parents provide their children with insurance against poor labor market opportunities (see, for instance, McElroy, 1985, Rosenzweig and Wolpin, 1993, or Ermish, 1996 and 1999). Several recent studies have analyzed the late emancipation age in southern European countries. For Italy, Manacorda and Moretti (2003) emphasize the income of parents, who are portrayed as bribing their children to stay at home longer, while Giannelli and Monfardini (2003) focus on the negative effect of housing costs on emancipation. The Spanish case is studied by Martinez-Granado and Ruiz-Castillo (2002), who also find a negative effect of housing costs, and it inspires the theoretical model of Díaz and Guilló (2003), which stresses the mother's housework as a public good inducing children to stay. In an international investigation, Ghidoni (2002) finds very few empirical regularities on the determinants of moving out across countries.

In this paper, we focus on one additional factor which has not received much attention so far, namely the degree of job security enjoyed by youths and their parents. Fogli (2000) has called attention to high parental job security as a determinant of late youth emancipation. She presents a model in which, due to credit market imperfections, granting high job protection to older workers is welfare improving. Young people remain with their parents, whose jobs are very secure, profiting from household consumption (a public good), thus avoiding the credit constraints they would face if they lived alone and went out to work.

While starting from her intuition, our interest in analyzing empirical evidence leads us to pursue a different model. Specifically, we study the residential decision of a child of altruistic parents in a dynamic setup; the child may stay or leave in a given period, but if she stays she gets the chance of revising her decision in the future (moving out is irreversible). Contrary to Fogli's model, the child may choose to work while coresiding with her parents. Further, we assume that she gets a higher fraction of family income while at home, but altruistic parents will still provide transfers to a poor but independent child. Lastly, in our setup there is uncertainty concerning the child's and the parents' future income. Our model, therefore, allows us to understand the roles of uncertainty and altruism on the child's residential decision, issues which are mostly absent in Fogli's setup given the general-equilibrium nature of her work.

We show that, for forward-looking children who are considering emancipation, given the irreversibility of moving out, staying home with the parents has an *option value* associated with waiting to see the realization of future incomes and deciding then whether or not to leave. Depending on her own and her parents' future income, a child who moved out may come to *regret* that she did. In the presence of partial altruism, there will be regret even if parents are wealthy enough to provide financial transfers to independent children. We also show that the joint consideration of parental and child's job security is a relevant explanatory factor for youth emancipation. In particular, we will argue that, under certain conditions, children's job insecurity lowers the probability that they leave home, while parental job insecurity raises it.

To test the effect of job insecurity on coresidence rates, we look first at macro data. In Section 3 we show that the aggregate evidence for 12 European Union (EU) countries in the 1980s and 1990s is consistent with our hypothesis. Micro-econometric evidence is given in Section 4 but only for Italy, the single country for which we have access to household-specific information on job insecurity and coresidence. We exploit the panel data structure of the Italian Survey of Household Income and Wealth (SHIW), collected by the Bank of Italy, which contains high quality data on individual-specific perceived job insecurity for 1995. Our particular sample and empirical strategy are described in detail in Sections 4.1-4.3. We estimate linear probability and probit models for whether children live independently after a given year, 1995, as a function of indicators of parental and child's job insecurity, and of a set of control variables measuring demographic, educational, and labor market characteristics of the children and their fathers (as well as other family background indicators). Very few papers in this literature have exploited in a similar way the panel structure of available microeconomic datasets. Indeed most papers present just cross-sectional evidence.

For parents, we use the answers to a question on the individual's perceived likelihood of having a job in the subsequent 12 months to construct an indicator of job insecurity. The high quality of this information is another valuable aspect of our data. Unfortunately this question was asked only to a randomly chosen 50% of the households, which leaves us with a small sample. In order to increase sample size, we impute perceived job insecurity to individuals who were not asked this question. To improve the reliability of our results, we implement Rubin's (1987) multiple imputation method. For children, due to problems of both sample size and potential endogeneity, and more in line with the literature, we cannot employ the same measure of perceived job insecurity which we use for parents. We therefore measure insecurity using unemployment rates at the province level, separate by gender and age classes. Our microeconometric results, discussed in Section 4.4, are consistent with the theoretical prediction that higher job security of the father deters emancipation.

In Section 5 we present our conclusions. Having established the quantitative importance of the effects of perceived job security on coresidence, we have uncovered an empirically significant link between labor market institutions –labor market rigidity in particular– and family demographics, which brings a new perspective on the long-standing debate on job security provisions.

2 A model of job insecurity and coresidence

In this section we illustrate how coresidence decisions are related to job insecurity of parents and children using a dynamic, two-period model of residential choice. Proofs of the results in this section can be found in Becker *et al.* (2004).

2.1 The family

The family in our model has n_0 parents and $n_1 + 1$ children. We assume that it has either one or two parents ("the parent" for short) and at least one child. Family size is denoted by n. Our focus is on the residential choice of one of the children, assuming that her siblings remain with their parents.

Direct utility is defined over consumption, only. We assume that, in the parental home, all individuals pool income and consume an equal fraction of total family income. We root this assumption on the difficulty of excluding children from the consumption of public goods in the household.[1] If all family members are coresiding, then consumption in the parental home is given by:

$$c_p^n = \frac{y_p + y_c - \gamma_p}{pn}, \tag{1}$$

where γ_p is the rent or the imputed cost of housing, p the price index of a representative basket of consumption goods, y_p parental income, and y_c the income of the child who is contemplating to move out (her siblings are assumed to earn no income). We denote the child's consumption by c_c. If she stays, she gets c_p^n. If she moves out, she will consume all of her income, net of housing costs under independence, γ_c, plus a non-negative transfer t from her parents:

$$c_i = \frac{y_c + t - \gamma_c}{p}.$$

Per capita consumption of the remaining family members is now given by:

$$c_p^i = \frac{y_p - t - \gamma_p}{p\,(n-1)}.$$

[1] Considering different sharing rules, provided that consumption of individual family members were monotonic in income, would not change our qualitative results.

We assume that parents are partially altruistic. They weigh their direct utility by a factor $\lambda \in (0.5, 1)$ and their children's utility by $(1-\lambda)$. Parental utility is then:

$$U_p = \lambda \left(n_0 + \frac{(1-\lambda)}{\lambda} n_1 \right) u(c_p) + (1-\lambda) u(c_c), \tag{2}$$

In what follows, we will in fact use the slightly modified functional form:

$$U_p = \lambda (n_0 + n_1) u(c_p) + (1-\lambda) u(c_c) = \lambda (n-1) u(c_p) + (1-\lambda) u(c_c), \tag{3}$$

which puts more weight on the utility of the n_1 children who always remain at home and simplifies the algebra significantly, while leaving our results qualitatively unchanged.

To obtain sharper results, we conduct all of our analysis with a specific functional form for the direct utility from consumption, namely the Constant Relative Risk Aversion (CRRA) form: $u(c) = (1-\alpha)^{-1} c^{1-\alpha}$, with $\alpha > 0$.[2]

2.2 Timing

There are two periods. In period 1, parent and child observe their income realizations, y_{p1} and y_{c1}. To insure non-negative consumption, we assume there is a lower bound on income realizations given by the housing costs, γ_p and γ_c. A positive income realization for the parent, interpreted as a draw of $y_{p1} > \gamma_p$, is equivalent to a job offer, and similarly for the child. Since there is no disutility from work, job offers are always accepted.[3] The child then decides whether or not to move out. Finally, consumption takes place as a function of the residential choice of the child.

The main difference across periods comes from assuming that moving out is irreversible, which can be justified on the grounds that the direct costs from moving, as well as the social stigma attached to going back to the parental house, tend to make independence a rather permanent state. Qualitatively similar results would emerge from considering finite costs instead. For a child who stayed with her parents in period 1, the

[2] The results generalize to other commonly used families of functions (such as the Constant Absolute Risk Aversion case).

[3] Family member j would require a positive income threshold above γ_j before accepting a job offer if there were disutility from work or if individuals were productive while unemployed (through household production, say). We are ignoring these cases.

period 2 timing of events and choices repeats itself. If the child has moved out in period 1, however, she faces no residential choice in period 2.

2.3 Period 2

We now characterize the resource allocation and residential decision in period 2. Assuming that the incomes of parent and child have taken the values y_{p2} and y_{c2}, the optimal transfer the parent would give the child if she decided to move out solves the following problem:

$$\max_{\tilde{t}_2 \geq 0} \left\{ \lambda \left(n-1\right) u \left(\frac{y_{p2} - \tilde{t}_2 - \gamma_p}{p \left(n-1\right)} \right) + \left(1-\lambda\right) u \left(\frac{y_{c2} + \tilde{t}_2 - \gamma_c}{p} \right) \right\}. \tag{4}$$

First-order conditions yield:

$$\lambda u' \left(c^i_{p2}\right) \geq \left(1-\lambda\right) u' \left(c_{i2}\right),$$

holding with equality when $\tilde{t}_2 > 0$. Since $\lambda > 0.5$, this implies that, when she receives a transfer, the child has lower consumption than the remaining family members. Under CRRA preferences, the positive transfer is given by:

$$t_2 = \frac{y_{p2} - \gamma_p - \left(y_{c2} - \gamma_c\right) \Gamma \left(n-1\right)}{1 + \Gamma \left(n-1\right)}, \tag{5}$$

where $\Gamma \equiv \left(\lambda / \left(1-\lambda\right)\right)^{\frac{1}{\alpha}} > 1$.

Using (5), we solve for consumption of the child when receiving a transfer:

$$c_{i2} \left(t_2 > 0\right) = \frac{y_{c2} + t_2 - \gamma_c}{p} = \frac{y_{p2} + y_{c2} - \gamma_p - \gamma_c}{\left(\Gamma \left(n-1\right) + 1\right) p}, \tag{6}$$

whereas at the parental home per capita consumption is:

$$c^i_{p2} \left(t_2 > 0\right) = \frac{\Gamma \left(y_{p2} + y_{c2} - \gamma_p - \gamma_c\right)}{\left(\Gamma \left(n-1\right) + 1\right) p}.$$

Comparing (1) and (6), it is easy to see that, when transfers are positive, the impact of higher income on the child's consumption is largest for a coresiding child. Indeed, a child who lives with her parents is entitled to the fraction $1/np$ of the goods consumed in the household; if she leaves, she will receive the smaller fraction $1/\left(\Gamma \left(n-1\right) + 1\right) p$ of the family's joint resources. This is the consequence of partial altruism and the parent's inability to limit the consumption of a coresiding child or, stated differently, it follows from the child's greater bargaining power at home relative to independence. While we find this a realistic assumption, the alternatives of full altruism or of identical bargaining power would not affect our qualitative results.

When a child leaves, the combined resources of the family drop by the cost γ_c. Further, since a transfer-recipient child will consume less than per capita consumption at the parent's house, a child with income low enough to trigger a transfer would be made worse off by moving out. In other words:

$$c_{i2}\left(t_2 > 0\right) < c_{p2}^n = \frac{y_{p2} + y_{c2} - \gamma_p}{np}.$$

We now address the moving out decision for the child who decided to stay at home in period 1. Define Δ_2 as the excess utility level when independent relative to coresiding, for period 2:

$$\Delta_2 \equiv u\left(c_{i2}\right) - u\left(c_{p2}^n\right).$$

The child moves out if $\Delta_2 > 0$. (If indifferent, $\Delta_2 = 0$, we assume she stays.)

How does Δ_2 change as a function of y_{c2}? Let us define $\tilde{y}_{c2}$ as the value such that parental transfers are zero, $t_2\left(\tilde{y}_{c2}\right) = 0$.[4] Under CRRA preferences,

$$\tilde{y}_{c2} = \frac{y_{p2} - \gamma_p}{\Gamma\left(n-1\right)} + \gamma_c. \tag{7}$$

Given the transfer function, we know that, for values of y_{c2} such that transfers were positive, the child's consumption would be lower if she were to move out. Consequently, the value that leaves the child indifferent between staying home and moving out, i.e. $\Delta_2\left(\bar{y}_{c2}\right) = 0$, must exceed $\tilde{y}_{c2}$. It is given by:

$$\frac{y_{p2} + \bar{y}_{c2} - \gamma_p}{np} = \frac{\bar{y}_{c2} - \gamma_c}{p}, \text{ or}$$

$$\bar{y}_{c2} = \frac{y_{p2} - \gamma_p}{n-1} + \frac{n}{n-1}\gamma_c > \tilde{y}_{c2}. \tag{8}$$

We now characterize formally how Δ_2 depends on the child's income.

[4]The notation $t_j\left(x\right)$ omits, for simplicity, other arguments of the function.

Lemma 1 *The function* $\Delta_2(y_{c2})$ *is strictly negative for* $y_{c2} \in [\gamma_c, \bar{y}_{c2})$ *and strictly positive for* $y_{c2} > \bar{y}_{c2}$*. Further,* $\Delta_2(y_{c2})$ *is strictly increasing in the range* $(\tilde{y}_{c2}, \bar{y}_{c2})$*. When the relative-risk aversion parameter* α *exceeds 1,* $\Delta_2(y_{c2})$ *is strictly increasing for* $y_{c2} \in (\gamma_c, \tilde{y}_{c2})$*. When* α *is below 1,* $\Delta_2(y_{c2})$ *is strictly increasing for* $y_{c2} > \bar{y}_{c2}$*.*

Figure 1A depicts a possible configuration of $\Delta_2(y_{c2})$. As Lemma 1 shows, utility from independence exceeds that under coresidence for $y_{c2} \geq \bar{y}_{c2}$. Thus, a child who did not move out in period 1, will now leave if her income exceeds $\bar{y}_{c2}$, otherwise she will stay. On the other hand, the set of income values $[\gamma_c, \bar{y}_{c2}]$ are a *regret region*: the child would prefer to go back home. The sources of regret are the rental cost γ_c and partial altruism.

It is worth discussing the ambiguity in the slope of $\Delta_2(y_{c2})$. Since Δ_2 corresponds to a difference in utility levels, changes in income affect this difference in two ways. First, income modifies consumption differently depending on the residential state. For example, for y_{c2} values such that no transfers would be provided to the child (i.e. above $\tilde{y}_{c2}$), higher y_{c2} implies that c_{i2} is changing by $1/p$ whereas consumption at home goes up by $1/np$. We label the impact of income changes on the child's consumption as the *sharing effect*. This is, however, not sufficient to ensure that Δ_2 varies positively with y_{c2}. The impact on Δ_2 depends also on the marginal utility that these changes in consumption entail. If, for example, $c_{i2} > c^n_{p2}$, the marginal utility of consumption at home is higher than under independence. In the range $y_{c2} > \bar{y}_{c2}$, this *marginal utility effect* counteracts the greater change in c_{i2} relative to c^n_{p2}. Consequently, although we know that c_{i2} will exceed c^n_{p2} for $y_{c2} > \bar{y}_{c2}$, we cannot be certain that Δ_2 is always positively sloped in this range. When $y_{c2} \in (\tilde{y}_{c2}, \bar{y}_{c2})$, by contrast, both effects go in the same direction, ensuring that Δ_2 is positively sloped. For $\alpha < 1$ (α governs the curvature of the utility function) one of the effects unambiguously dominates the other.[5]

The assumptions that the marginal utility effect dominates for low income and the sharing effect for high income are reasonable. For low income, it is likely that the

[5] If $\lambda = 0.5$ the parent gives an independent child the same fraction $1/n$ of total resources as at home and so, in the range of positive transfers, the sharing effect disappears.

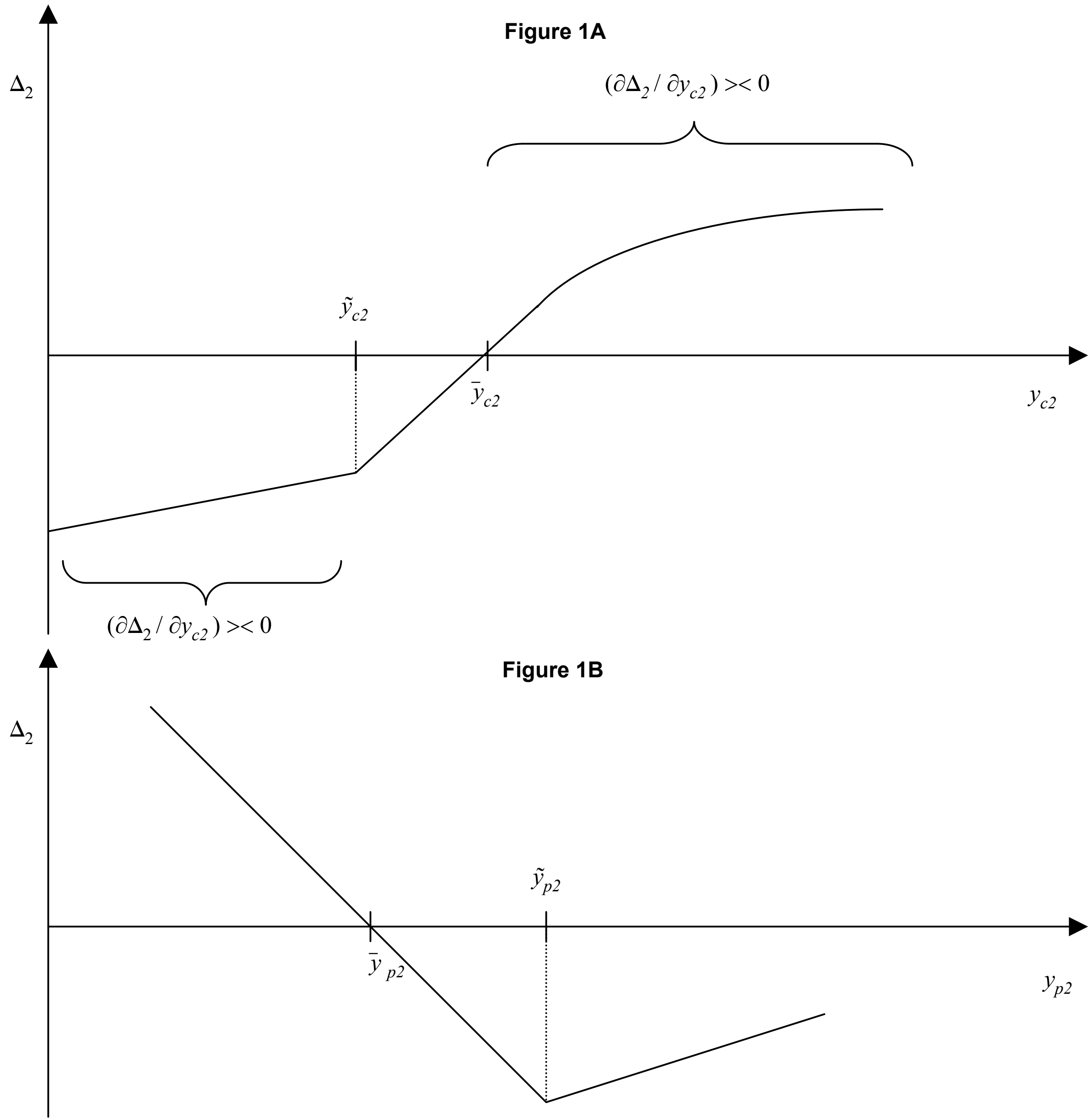

Figure 1. The moving out decision in period 2

marginal utility effect will assume an important role in the comparison of utilities across residential states since CRRA preferences satisfy Inada conditions.

The dependence of the residential choice on y_{p2} is also interesting. Define $\tilde{y}_{p2}$ as the level of parental income such that $t_2(\tilde{y}_{p2}) = 0$, and $\bar{y}_{p2}$ as the parental income level that leaves the child indifferent between moving out and coresiding, $\Delta_2(\bar{y}_{p2}) = 0$. It is straightforward to show that $\bar{y}_{p2} < \tilde{y}_{p2}$. Then,

Lemma 2 *The function $\Delta_2(y_{p2})$ is strictly decreasing for $y_{p2} \in [\gamma_p, \tilde{y}_{p2})$ and strictly negative for $y_{p2} > \bar{y}_{p2}$. For the latter range of y_{p2} values, when the relative-risk aversion parameter α exceeds unity, $\Delta_2(y_{p2})$ is strictly increasing.*

In Figure 1B we depict a possible configuration for $\Delta_2(y_{p2})$. Whether or not $\Delta_2(\gamma_p)$ is positive depends on parameter values (specifically, a large number of family members n and a small rental cost γ_c make $\Delta_2(\gamma_p)$ positive). As Lemma 2 shows, however, for $y_{p2} > \bar{y}_{p2}$, $\Delta_2(y_{p2}) < 0$ holds unambiguously, and children of wealthy parents who stayed home will not move out. Just as with $\Delta_2(y_{c2})$, higher parental income does not necessarily raise the child's willingness to stay home. In Figure 1B, $\Delta_2(y_{p2})$ is consistent with assuming that the marginal utility effect dominates the sharing effect for the range of positive transfers.

In Figure 2 we plot the curves $\tilde{y}_{c2}$ and $\bar{y}_{c2}$ in (y_c, y_p) space. To the right of the $\bar{y}_{c2}(y_{p2})$ schedule, the child moves out; to the left she stays. From the point of view of the moving-out decision taken in period 1, we can also divide the coresidence area into two parts. To the left of $\tilde{y}_{c2}(y_{p2})$, children who became independent in period 1 will receive a transfer, i.e. $t_2(y_{c2}, y_{p2}) > 0$, while to the right they will not. Recall that the coresidence area is a regret area.

2.4 Period 1

A simplified presentation of the model's structure is given in Figure 3. In period 1 the residential choice is more involved than in period 2, due to irreversibility and the possibility of regret. Naturally, the latter depends on the likelihood that period 2 incomes

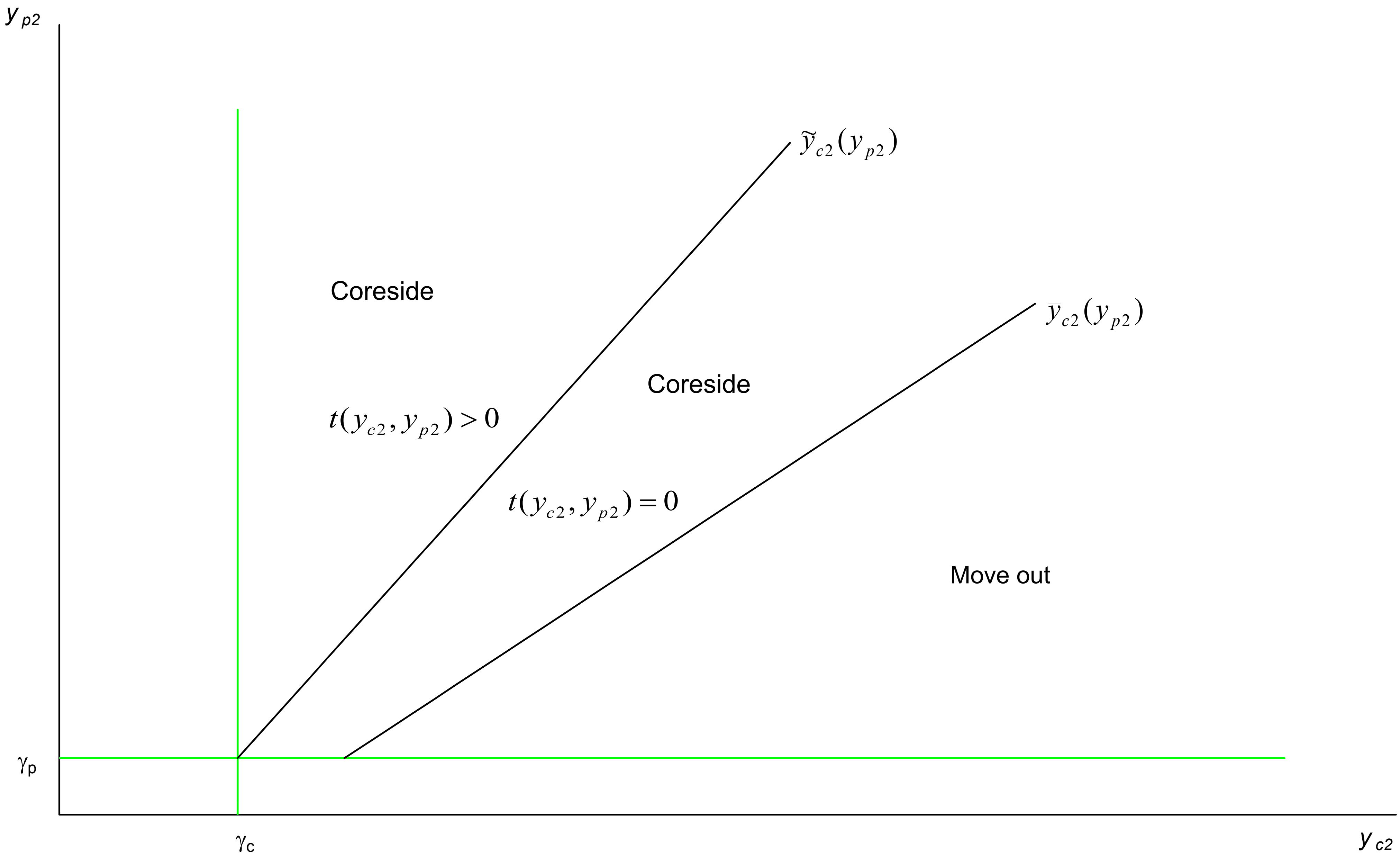

Figure 2. Period 2 residential regimes

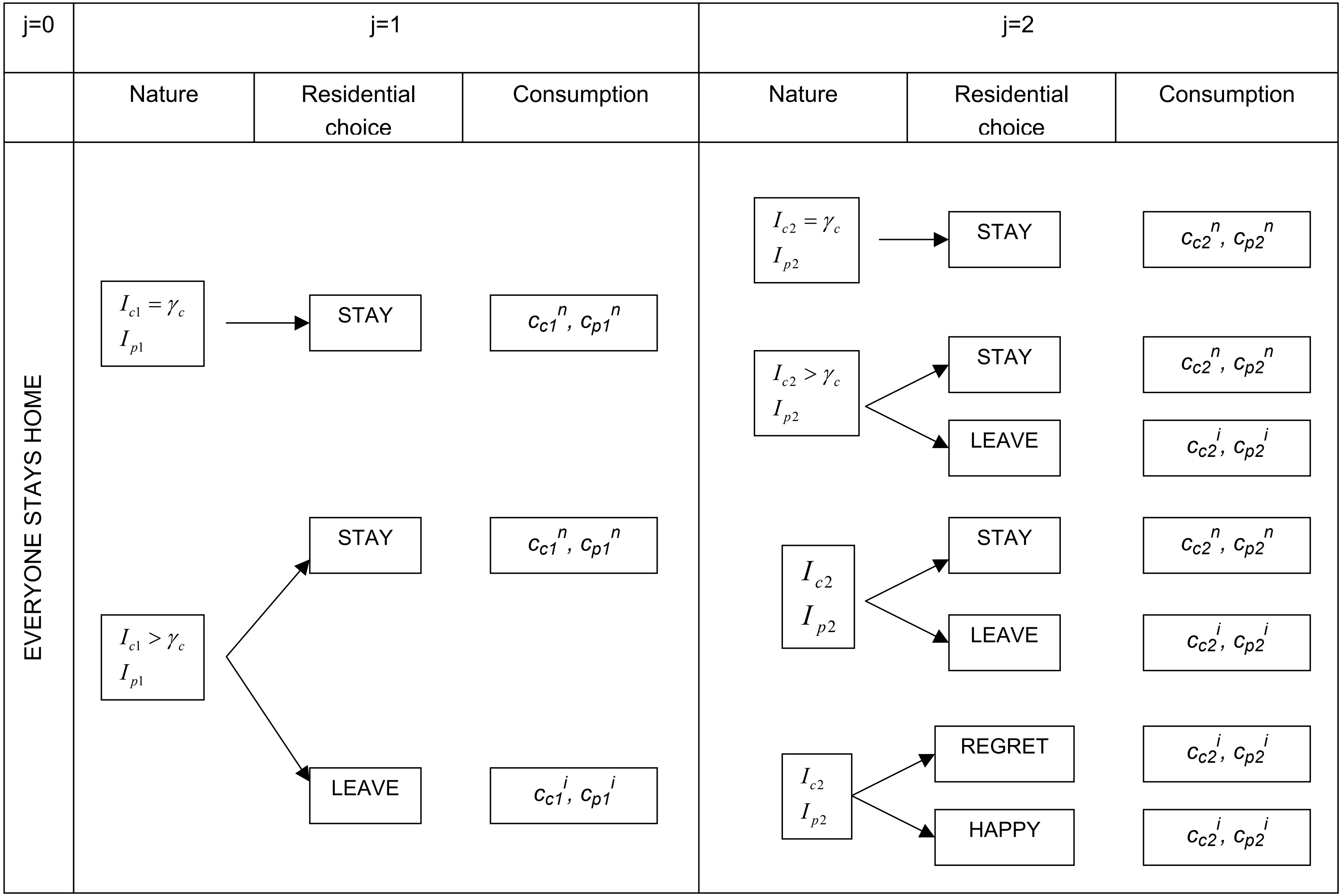

Figure 3. Structure of the model

will fall to the left of the schedule $\bar{y}_{c2}(y_{p2})$. So far, we have been silent about the distribution of income, since period 2 decisions are taken after y_{c2} and y_{p2} are observed. We now assume that $(y_{c2}, y_{p2}) \sim F(y_{c2}, y_{p2})$, where $F(\cdot)$ is the joint cumulative distribution function (cdf) of period 2 income (y_{c2}, y_{p2}), with marginal cdf's $F_c(y_{c2})$ and $F_p(y_{p2})$. We denote the corresponding probability density functions (pdf's) by $f(y_{c2}, y_{p2})$, $f_c(y_{c2})$ and $f_p(y_{p2})$. $F(\cdot)$ has support over $[\gamma_c, \infty) \times [\gamma_p, \infty)$.

If $F(\cdot)$ assigns positive probability to the regret region, staying home in period 1 has an *option value*, the value associated with waiting to see the realization of the period 2 income and deciding then whether or not to move out. Just like with any real option, this value has to be weighted against the potential gains from moving out early on.

Define as Δ_1 the expected excess utility from moving out relative to staying home conditional on making the optimal residential choice in period 2:

$$\Delta_1 \equiv u(c_{i1}) + \int_{\gamma_p} \int_{\gamma_c} u(c_{i2})\, dF(y_{c2}, y_{p2})$$

$$- \left\{ u\left(c_{p1}^n\right) + \int_{\gamma_p} \left[\begin{array}{c} \int_{\gamma_c}^{\bar{y}_{c2}(y_{p2})} u\left(c_{p2}^n\right) dF_c(y_{c2}|y_{p2}) \\ + \int_{\bar{y}_{c2}(y_{p2})}^{\infty} u(c_{i2})\, dF_c(y_{c2}|y_{p2}) \end{array} \right] dF_p(y_{p2}) \right\}.$$

Δ_1 is defined over the set $[\gamma_c, \infty) \times [\gamma_p, \infty)$.

The first two terms in Δ_1 represent the expected utility from moving out in period 1. Given that the child becomes independent in period 1, period 2 utility is also computed for $c_{c2} = c_{i2}$. The terms preceded by a minus sign represent the expected utility from staying home in period 1. In this case, the child retains the possibility of choosing the best residential arrangement in period 2. Thus, given y_{p2}, for $y_{c2} \leq \bar{y}_{c2}(y_{p2})$, the child remains with her parents and $c_{c2} = c_{p2}^n$; otherwise she moves out and $c_{c2} = c_{i2}$. The child will move out if $\Delta_1 > 0$. When $y_{c2} > \bar{y}_{c2}(y_{p2})$, having moved out in period 1 does not carry any utility loss; therefore, in this range, the terms concerning period 2 utility while independent cancel out and the moving out condition simplifies to:

$$u(c_{i1}) - u\left(c_{p1}^n\right) > \int_{\gamma_p} \left[\int_{\gamma_c}^{\bar{y}_{c2}(y_{p2})} \left(u\left(c_{p2}^n\right) - u(c_{i2}) \right) dF_c(y_{c2}|y_{p2}) \right] dF_p(y_{p2}). \qquad (9)$$

It is worth examining equation (9) in detail. First of all, the right-hand side is nonnegative. It represents the difference between expected utility under coresidence and under independence, i.e. the gain in expected utility associated with waiting for period 2 before choosing whether or not to move out (the option value). It will be strictly positive if the cdf $F(\cdot)$ places positive mass on the regret region. The left-hand side represents the difference in period 1 utility from being independent relative to moving out. The child will move out when this gain exceeds the expected benefit from waiting.

Now we discuss the determination of $\bar{y}_{c1}$, the child's period 1 moving-out threshold. As before, it is defined as the root of Δ_1, namely, $\Delta_1(\bar{y}_{c1}) = 0$. We note the following facts:

Lemma 3 *The period 1 schedule $\tilde{y}_{c1}(y_{p1})$ such that, for $y_{c1} > \tilde{y}_{c1}(y_{p1})$, parents provide no transfers, is identical to the period 2 function $\tilde{y}_{c2}(y_{p2})$.*

This lemma says that, on the space (y_c, y_p) of incomes, the functions $\tilde{y}_{c2}(y_p)$ and $\tilde{y}_{c1}(y_p)$ exactly overlap. Given that the transfer-giving threshold is identical for periods 1 and 2, the left-hand side of equation (9) coincides with the function Δ_2 –now with arguments (y_{c1}, y_{p1}). The properties of function Δ_2 in terms of (y_{c2}, y_{p2}) were characterized in Lemmas 1 and 2.

The regret region, namely period 2 income pairs such that the child prefers coresidence, is formally defined as:

$$R \equiv \{(y_{c2}, y_{p2}) \in [\gamma_c, \infty) \times [\gamma_p, \infty) : y_{c2} \leq \bar{y}_{c2}(y_{p2})\}.$$

If the probability of regret is zero, then today's residential choice becomes static and depends only on the comparison of the current utility levels attained under the two residential arrangements. Thus, the residential decision of period 1 is identical to that of period 2. For the case when there is a positive probability of regret ($F(R) > 0$), we can show the following:

Proposition 4 *The period 1 moving-out threshold correspondence* $\bar{y}_{c1}(y_{p1})$*, on* (y_{c1}, y_{p1}) *space, lies strictly to the right of the corresponding period 2 schedule* $\bar{y}_{c2}(y_{p2})$ *if and only if* $F(R) > 0$*. When* $\alpha < 1$*,* $\bar{y}_{c1}(y_{p1})$ *is single-valued.*

Intuitively, the proposition says that the possibility of future regret makes people more conservative about moving out.

We are interested in the impact of y_{p1} on the period 1 moving-out threshold. First, define expected regret as:

$$\bar{R} \equiv \int_R \left(u\left(c_{p2}^n\right) - u\left(c_{i2}\right)\right) dF\left(y_{c2}, y_{p2}\right).$$

We now rewrite the condition $\Delta_1 = 0$ including the arguments of c_{i1} and c_{p1}^n:

$$u\left(\frac{\bar{y}_{c1} - \gamma_c}{p}\right) - u\left(\frac{y_{p1} + \bar{y}_{c1} - \gamma_p}{np}\right) = \bar{R},$$

or, in more compact notation,

$$\Delta_2\left(\bar{y}_{c1}, y_{p1}\right) = \bar{R}. \tag{10}$$

We are ready to state the following result:

Lemma 5 *Higher period 1 parental income* y_{p1} *raises the child's moving-out threshold* $\bar{y}_{c1}(y_{p1})$ *for* $\alpha < 1$*.*

As before, $\alpha < 1$ is a sufficient condition. The result in the lemma would also follow under other parameter values provided the sharing effect dominates the marginal utility effect for $y_{c1} > \bar{y}_{c2}$.

Let us consider now the impact of expectations. How do different distribution functions $F(\cdot)$ affect $\bar{y}_{c1}$? For simplicity, we assume that the distributions of period 2 incomes of parent and child are independent, and so: $F(y_{c2}, y_{p2}) = F_c(y_{c2}) F_p(y_{p2})$. Our preferred interpretation of insecurity is a measure of the probability that income will assume a low future realization. In this sense, a natural way of comparing income distributions is to rank them in terms of first-order stochastic dominance. Let $F_c^1(y_{c2})$ first-order stochastically dominate $F_c^2(y_{c2})$. Then, $F_c^1(y_{c2}) \leq F_c^2(y_{c2})$, for all y_{c2} values, and the expectation over nondecreasing functions of y_{c2} computed under the distribution F_c^1 exceeds the same expectation under F_c^2. (The same applies to comparisons between $F_p^1(y_{p2})$ and $F_p^2(y_{p2})$).[6]

Intuition suggests that if the child received the good news that the distribution of her second period income had changed from F_c^2 to F_c^1, the wealthier child (in expectation) would be more willing to move out in period 1. Likewise, if the parent's income was known, in period one, to be distributed according to $F_p^1(y_{p2})$ rather than $F_p^2(y_{p2})$, we would expect his child to be more reluctant to leave the parental home. As it turns out, this is not always the case. The reason is the ambiguous impact of higher income on the excess utility from independence, Δ_2, described in Lemmas 1 and 2.

To illustrate the problem, go to Figure 2 and fix a particular value y_{c2} to the right of $\bar{y}_{c2}(y_{p2})$. Starting at $y_{p2} = \gamma_p$ and moving up while holding y_{c2} constant, we move from the moving-out region into the regret region. Below $\tilde{y}_{c2}(y_{p2})$ parents cannot provide financial help and higher parental income reduces Δ_2. But if we increase y_{p2} to the point of moving into the positive transfer region, higher parental income may increase or decrease Δ_2. If we restrict the distributions F_p^1, F_p^2, and F_c to put mass only on combinations of (y_{c2}, y_{p2}) below the schedule $\tilde{y}_{c2}(y_{p2})$, then the result that F_p^1 induces a higher period 1 moving-out threshold relative to F_p^2 goes through. Similarly, as we fix a value for y_{p2} and increase the child's income, Δ_2 may be increasing or decreasing, but it unambiguously increases within the no-transfers subset of the regret region. Again, if we restrict the distributions F_c^1 and F_c^2 not to put mass on the positive-transfers subset of the regret region, the intuitive result outlined above will prevail. We formalize it below.

Let $\mathcal{F}$ be the set of all pairs of independent distributions functions (F_c, F_p) with support over $([y_c, \infty), [y_p, \infty))$, such that no mass is placed on the positive-transfers subset of the regret region. Then:

Proposition 6 *Let* $\left(F_p^1, F_c\right)$ *and* $\left(F_p^2, F_c\right)$ *be two elements of* $\mathcal{F}$*, and assume that* F_p^1

[6] The results associated with first- and second-order stochastic dominance require the additional assumption that income is bounded. That is, $y_{p2} \leq y_p^* < \infty$ and $y_{c2} \leq y_c^* < \infty$, so that $F_p(\gamma_p) = F_c(\gamma_c) = 0$ and $F_p\left(y_p^*\right) = F_c\left(y_c^*\right) = 1$. We omit making this assumption explicit for simplicity.

first-order stochastically dominates F_p^2. Let the period 1 moving-out threshold corresponding to F_p^j be denoted $\bar{y}_{c1}\left(F_p^j\right)$. Then, when $\alpha < 1$, $\bar{y}_{c1}\left(F_p^1\right) \geq \bar{y}_{c1}\left(F_p^2\right)$.

Proposition 7 *Let (F_p, F_c^1) and (F_p, F_c^2) be two elements of $\mathcal{F}$, and assume that F_c^1 first-order stochastically dominates F_c^2. Let the period 1 moving-out threshold corresponding to F_c^j be denoted $\bar{y}_{c1}(F_c^j)$. Then, when $\alpha < 1$, $\bar{y}_{c1}(F_c^1) \leq \bar{y}_{c1}(F_c^2)$.*

Appendix A provides a discussion of how Propositions (6) and (7) can be generalized.[7]

Lastly, in Appendix A we also address the impact of higher variance of the future income streams of parent and child on the moving-out decision. We compare the variance of income under different distributions while holding the expected value constant (i.e. second-order stochastic dominance). We are able to show that a lower variance in the child's future income stream makes her less reluctant to leave the parental home. We get the opposite reaction concerning the variance of parental income. For the parent's income, however, since changes in the distribution of income affect the period 1 moving out threshold only to the extent that they affect period 2 income values within the regret area, we need to ensure that the distribution of his income shifts so as to raise the variance of income values in that region specifically, as opposed to the requirement that it becomes more volatile over its global range.

Let us now summarize the key results provided by the model as a guide for empirical work:

(i) *The option value.* For forward-looking children who are considering emancipation, given the stochastic nature of future income, staying home with the parents has an *option value*. This is the value associated with waiting to see the realization of future incomes and deciding then whether or not to move. The reason is that moving out and then back in is costly; in the stylized case of our model, moving out is irreversible. Thus, depending on her own and her parent's future income, a child who moved out may come to *regret* that she did. In the presence of partial altruism (parents value the child's

[7] As to family size, only for the case of $\alpha < 1$ and very altruistic parents ($\lambda = 0.5$) is it possible to show that the first period moving-out threshold is lower than the corresponding threshold for a smaller family (see Becker *et al.*, 2004).

consumption less than theirs), there will be regret even if parents are wealthy enough to provide financial transfers to independent children.

(ii) *Job insecurity*. We have considered two dimensions of insecurity or income risk: changes in expected income and changes in the variance of the future income processes. Changes in income influence future regret sometimes in unintuitive ways. The reason is that higher income not only induces changes in consumption which depend on the residential state (the sharing effect) but these consumption changes, in turn, affect the marginal utility from consumption (the marginal utility effect) which also differs across residential choices. When these effects work in opposite directions, only by imposing some restriction on parameter values or income distributions is it possible to determine the impact of income on regret. If the child faces higher expected income, provided the marginal utility effect dominates at low levels of consumption but is replaced by the sharing effect at high consumption levels, she will be more eager to leave the parental home. The effects of higher expected parental income are not prone to a similar generalization; only under some restrictions on the distribution of future parental income can we make the statement that higher expected future income will raise the child's moving-out threshold. The ambiguous effect of income on regret again limits the generality of statements concerning the effects of uncertainty on the residential choice. Only under particular income distributions can we conclude that higher variance in the child's income and lower variance in the parent's (holding the expected value constant in both cases) make the child more reluctant to leave home.

Thus, the model tells us that job insecurity affects the child's residential choice. The direction of the effects of insecurity depends on the aforementioned conditions. Establishing whether these conditions hold, as well as their quantitative relevance, is largely an empirical issue which makes it worth considering both macro and microeconomic evidence. We next turn to this task.

3 Job insecurity of parents and children in the EU

In this section, we test some of the model's predictions at the macro level. Specifically, we aim at testing the theoretical result that higher own expected income induces children to leave the parental home earlier, while higher expected parental income has the opposite effect. This is done by estimating the aggregate relationship between coresidence rates and survey measures of perceived job insecurity for 12 European countries.

For coresidence, we analyze the aggregate fractions of males and females, aged 20-24 and 25-29 years old, who live at the parental home. For this purpose, data from the European Labor Force Survey is at hand for most EU countries over the period 1983-2002, but the availability of a measure of perceived job insecurity is much more restricted. We construct it using the European Commission's Eurobarometer. The survey asked individuals whether they had worried over the last year about losing their job or not finding one (1984) or to assess the risk of losing their job (1992 and 1997).[8] We transform the available set of individual responses into a 0-1 dummy variable (not insecure-insecure) and then compute the age- and gender-specific fractions of individuals who report being insecure. Data are available for 4 countries in 1984 (3197 individual observations) and 12 countries in 1992 and 1997 (4982 and 5067 observations, respectively). We end up with 100 cells (16, 36, and 48, for the three years, respectively). Appendix B gives more details. We also compute job insecurity for people aged 50-59 years old, who are representative of parents.

Table 1 summarizes the variables for the endpoints, 1984 and 1997. The coresidence rate naturally falls with age, it is lower for females than for males, and it is higher in Mediterranean, predominantly Catholic countries –Italy, Spain, Portugal, and Greece– and in Finland, than elsewhere in Europe.[9] Long-term trends cannot be gauged from Table 1, which only contains data on four countries for 1984. Patterns are not uniform

[8]The wording of questions is not the same in all surveys, so that comparability is a relevant concern. See Appendix B.

[9]It is striking to find that a Nordic country like Finland is in this respect very similar to the Mediterranean countries. But then note that job insecurity perceived by youth is the second highest in the Table.

Table 1: Descriptive statistics on coresidence and perceived job insecurity, 1984 and 1997 (%)

	Coresidence rate		Job insecurity perceived by youth		Job insecurity perceived by 50-59 y.o.	
	1984	1997	1984	1997	1984	1997
A. All	37.6	51.4	26.0	35.5	42.7	44.9
B. Age						
20-24 y.o.	55.0	67.0	28.3	36.9		
25-29 y.o.	20.2	35.7	23.7	34.1		
C. Gender						
Male	46.6	59.0	27.6	37.3		
Female	28.5	43.7	24.4	33.7		
D. Country						
Belgium		40.6		33.6		71.0
West Germany	34.5	37.2	17.6	25.2	21.2	47.2
Greece		63.1		34.7		48.6
Spain		75.0		51.7		47.6
France	28.1	35.2	29.2	53.6	59.1	67.2
Ireland		44.8		30.3		33.3
Italy	56.1	75.0	20.1	40.4	36.7	25.9
Netherlands		30.9		34.5		35.6
Austria		47.8		24.8		29.6
Portugal		68.7		29.5		23.9
United Kingdom	31.6	30.8	37.2	24.9	53.9	52.9
Finland		67.1		42.7		55.7

Note: Coresidence rate: percentage of youth population living at parental home. Unemployment risk: percentage of respondents who think that their job is at risk (See Appendix B for definitions). Sources: Coresidence rate: European Labor Force Survey. Unemployment risk: Eurobarometers 20 (1984) and 47.1 (1997).

Table 2: Coresidence and perceived job insecurity (%)

	Coefficient	Std. Error
Constant	58.31	3.16
Germany	-0.93	2.44
Greece	21.12	2.64
Spain	31.77	2.91
France	-5.87	2.52
Ireland	6.03	2.72
Italy	26.99	2.79
Netherlands	-8.63	2.62
Austria	6.53	3.27
Portugal	22.00	2.70
United Kingdom	-4.84	2.54
Finland	25.75	3.25
Age 2529	-32.66	1.03
Female	-15.79	1.02
1992	3.85	1.81
1997	6.59	1.76
Insecurity	13.00	5.16
Insecurity 5059	-8.76	4.24
No. of observations	100	
Adjusted R^2	0.956	

Note: The table reports, in percentage form, OLS estimates of equation (11). The dependent variable is the coresidence rate (in %). Descriptive statistics for the data are given in Table 1 and a description of the databases is given in Appendix B. The reference cell is that of males aged 20-24 y.o. living in Belgium in 1984.

across the 12 countries in the full period (1983-2002). The above 5 countries plus France experience upward trends; Belgium and the Netherlands feature mild upward trends only in the 1990s; and Ireland, Germany, and the UK show mild reductions among those aged 20-24 y.o. and stability in the 25-29 y.o. group. Thus, there is a majority of countries experiencing an increase in coresidence among the latter group in the 1990s.

Perceived job insecurity is lower among those aged 25-29 y.o. than among the youngest group, and highest among older people (50-59 y.o.), except in Italy, Portugal, and Spain. It is also somewhat higher for males than for females. Across countries, young workers feel most insecure in France, Spain, Finland, and Italy, while older workers do so in Belgium, France, Finland, and the UK.[10]

Figure 4 represents coresidence and perceived job insecurity for each of the two youth age and gender groups, pooling together all three years.[11] The graph suggests, *prima facie*, a positive relationship between the two variables. We test this relationship in a multivariate framework, by estimating the following regression of $Coresidence$ rates on job insecurity perceived by youth (*Insecurity*) and by the older age group (*Insecurity 5059*):

$$\begin{aligned} Coresidence_{ijt} &= \alpha_i + \beta \; Age\ 2529 + \delta \; Female + \lambda \; 1992 + \phi \; 1997 \\ &\quad + \gamma \; Insecurity_{ijt} + \rho \; Insecurity\ 5059_{it} + e_{ijt} \end{aligned} \tag{11}$$

where the subindex i denotes countries, j age-gender cells, and $t = 1984, 1992$, 1997. The α_i are country effects, Age 2529 and $Female$ are dummy variables for cells in those age and gender groups, respectively, 1992 and 1997 are year dummies, and e_{it} is random noise. The reference cell is that of males aged 20-24 y.o. living in Belgium in 1984.

Estimation results are shown in Table 2. The age- and gender-specific dummies confirm the features apparent in Table 1, and the year dummies suggest an upward

[10]Böckerman (2002), using individual data from a survey for 15 EU countries plus Norway in 1998, finds that perceived job insecurity increases with age but does not vary with gender, and also large cross-country differences after controling for a long list of variables (it is largest in Spain, Greece, and Italy).

[11]Legenda: BE Belgium, GE Germany, GR Greece, SP Spain, FR France, IR Ireland, IT Italy, NL Netherlands, AT Austria, PT Portugal, FI Finland, UK United Kingdom.

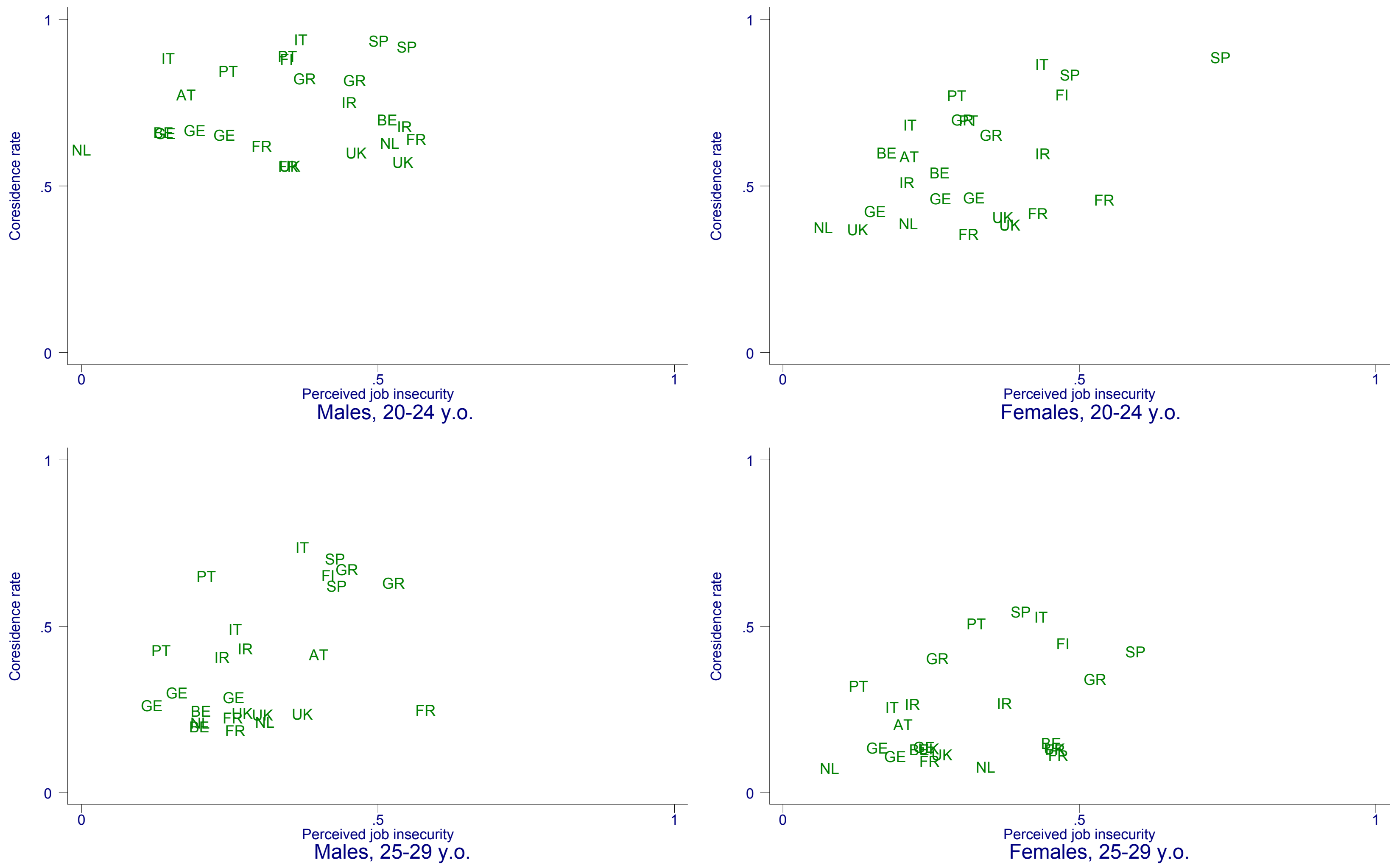

Figure 4. Coresidence and youth job insecurity (1984,1992,1997)

trend. The country dummies should mainly capture cross-country cultural differences, i.e. differences in preferences about coresidence –importance of family ties, attitudes regarding partnership formation, taste for independence, etc.–, and also in the costs of moving out for youths –availability and cost of rental and owner-occupied housing–, to the extent that these aspects have remained constant over time. They are significant and, net of age and gender composition effects, they identify the same five countries mentioned above as having high coresidence rates.

Most importantly, Table 2 reveals that once these characteristics are controlled for, job insecurity perceived by youth significantly raises the coresidence rate, while job insecurity of older age groups –proxying for the parents– lowers that rate. In particular, a 10 percentage point increase in the fraction of youths who perceive their job to be insecure is associated, *ceteris paribus*, with an increase in the coresidence rate of 1.3 percentage points, while a corresponding increase in the fraction of workers in their 50s is associated with a reduction of 0.9 percentage points. In other words, if the percentage of youths feeling insecure went from 0 to 100, the coresidence rate would increase by 13 percentage points, and the same change in the percentage of older wokers would reduce coresidence by 9 percentage points. This is not a negligible change, given that the average coresidence rate in the data is 47%. Moreover, the inclusion of the two job insecurity variables reduces the standard deviation of the country dummy coefficients by 3.1%, i.e. it explains part of the variation in coresidence rates across countries.

There is a large literature analyzing the determinants of perceived job insecurity in the US (e.g. Manski and Straub, 2000), the UK (e.g. Green *et al.*, 2000) and other European countries (OECD, 1997; Böckerman, 2002). Differences in perceived job insecurity depend on national idiosyncrasies and on differences in youth unemployment rates, as well as in underlying labor market institutions, such as employment protection legislation and unemployment benefits. Taking the degree of perceived job insecurity as given, our goal is to show that it is related to coresidence. The aggregate evidence for European countries we have uncovered indicates that, once cross-country differences in culture and institutions have been controlled for, job insecurity of parents and children influence in a significant way the coresidence choices of European youth.

Next, we test with micro-data the possibility that job insecurity of parents and children influences the moving out decision of the latter. We can do this only for Italy because this is the single country for which we found suitable data, which are described in the following section.

4 Microeconometric evidence for Italy

In this section we integrate the macro evidence just presented with micro evidence, by estimating the extent to which job insecurity of parents affects the decision of children to leave home in Italy.

4.1 Data and sample design

We use a representative sample of Italian individuals (the "children") of working age (more precisely, between 18 and 35 y.o.) and living with their parents. This sample has been extracted from the Italian Survey of Household Income and Wealth (SHIW). We use the 1995 wave of this survey, which contains information on 8135 households and 23924 individuals, to select our baseline sample. We then use the 1998 wave to obtain information on whether a child has left home between 1995 and 1998. Our goal is to test whether a measure of job insecurity of the father affects this decision, controlling for observable confounding factors measured in 1995 or before.[12]

A first difficulty in the construction of our data comes from the fact that the emancipation decision of children (the *outcome*) can only be observed for households which have been interviewed in both 1995 and 1998. Only 2699 of the 8135 households satisfy this condition and, thus, our analysis can only focus on this subset of the original SHIW sample.[13] Note, however, that since these panel households were randomly selected,

[12]We focus on job insecurity of fathers, not of both parents, because the labor participation rate of married women is low in Italy. Nevertheless, we also control for whether the mother works in our empirical analysis, in order to capture the availability of public goods such as household services.

[13]See Banca d'Italia (1997, 2000).

they are still representative of the reference population. So this data limitation should only reduce the efficiency of our estimates, not their reliability.

In our empirical investigation we will use an unusual measure of *perceived job insecurity*, which we will describe in detail below, constructed from the answers to a survey question in which individuals are asked about the probability of having a job in the following year. This question has been designed carefully to measure various dimensions of how uncertainty is perceived and is the main reason why the 1995 wave of the SHIW is particularly useful for our purposes.

It has two problems, however. First, only individuals who are either working or unemployed are asked about their job prospects. This excludes retired "house-husbands" and students. In principle, we could have considered retired fathers as having a sort of perfectly secure job, since they are in large part individuals who enjoy perfectly safe incomes. We do not do so because retired fathers are more likely to be at home all day, and this might affect the emancipation decision of children for reasons different from the pure effect of job security. Another reason to drop households with retired fathers is that to be completely sure about having no unemployment in the subsequent year is not equivalent to being sure for life because of retirement. Since we are interested in emancipation, we also restrict the sample to children aged up to 35 years old in 1995. These criteria, while required by the focus of our analysis, reduce the sample considerably, to 1142 children, but note that this sample is still representative of the population of children living in households where fathers are not retired in 1995.[14] The characteristics of this sample are described in the first two columns of Table 3.

The second data problem related to the question on perceived uncertainty is that to limit the length of the questionnaire this question has not been asked to labor force members in *all* households, but only to labor force members in households in which the head-husband was born in an odd year.[15] Thus, our measure of perceived uncertainty

[14]To be included in the sample, children must also be still alive, not in jail and not long-term hospitalized in 1998; these restrictions only affect a marginal number of observations.

[15]Other questions were asked in households where the father was born in even years.

is available only for household members belonging to the intersection between the panel subset of the SHIW and the subset in which information on job insecurity was collected.

As a result, only for 479 of the households described in the first two columns of Table 3 we observe a measure of perceived job insecurity of fathers and only for 212 of these households we have also an analogous measure for children. It is important to note that while the sampling design ensures that the 479 households for which paternal insecurity is available are on average observationally equivalent to the 1142 households for whom we have two years of data (see the last two columns of Table 3), this is not true for the 212 households for which the information is available for both fathers and children. The reason is that the question has not been asked to all individuals born in odd years, but to all individuals belonging to households in which the father was born in odd years.[16]

These features of the data bring good and a bad news. The good news is that we are in the perfect condition to apply the Multiple Imputation method proposed by Rubin (1987)[17] to impute perceived job insecurity to the fathers for whom this information is missing. As explained in Section 4.3 the reason is that, by construction, the missing information is missing "completely at random", which is a crucial necessary condition for the application of Rubin's method. The bad news is that since for children the information is *not* missing at random we cannot use the same imputation procedure. Moreover, the number of observations for which the information is available for children and fathers together is so small that any other imputation strategy for children would give unreliable results. Thus, for children we cannot use the measure of job insecurity constructed on the basis of the SHIW question.

Before moving to a more detailed description of our indicators of job insecurity and children emancipation, let us note again that while data limitations force us to use a relatively small sample, this sample is still representative of the population of interest

[16] A careful reader may wonder why the information on job insecurity is available only for 479 fathers and not for 571 = 1142/2 fathers, if it was asked to fathers born in odd years. The reason is that some fathers did not respond to the question. The comparison of observables in Table 3 clearly shows, however, that the cases of non-response are randomly distributed in the data.

[17] See also, more recently, Rubin (1996).

Table 3: Descriptive statistics for the sample used in the econometric analysis. 1995

	Full sample		Restricted sample	
Variable	Mean	Std. Dev.	Mean	Std. Dev.
Age	22.57	3.61	22.59	3.54
Female	0.45	0.50	0.44	0.50
Homeownership	0.71	0.45	0.72	0.45
Father's age	52.23	5.37	51.85	5.11
Father's years of schooling	9.24	4.12	9.40	4.11
Number of kids	2.37	1.04	2.36	0.87
Wealth	0.34	0.50	0.34	0.55
Mother employed	0.35	0.48	0.36	0.48
Rental index	6.44	2.11	6.30	2.19
Northwest	0.14	0.35	0.13	0.33
Northeast	0.19	0.39	0.16	0.37
Center	0.20	0.40	0.17	0.38
South	0.36	0.48	0.43	0.50
Islands	0.12	0.32	0.11	0.31
U rate in kid's age-gender-province cell	0.39	0.23	0.41	0.23
Father's perceived job insecurity	–	–	0.16	0.28
Out in 1998	0.04	0.18	0.04	0.20

Descriptive statistics of variables measured in 1995 for the *full sample* of 1142 children who:

- live with both of their parents in 1995,
- belong to households interviewed in both 1995 and 1998,
- are aged between 18 and 35 years old in 1995,
- have a father who is either employed or unemployed (i.e. not retired),
- are still alive, not in jail and not long-term hospitalized in 1998,

and for the *restricted* sample of 479 children whose father answered the question concerning perceived uncertainty. Therefore father's perceived job insecurity is recorded only for the restricted sample. Monetary variables are in millions of Italian Liras.

(see Table 3). Moreover, its timing structure is suitable for exploring the relationship between job insecurity (in particular paternal insecurity) and the subsequent (not contemporaneous) decisions of children to leave home controlling for a large set of individual and family background characteristics.

4.2 The indicators of job insecurity and the outcome variable

The first key variable in our analysis is the reply to the following question, posed to employed and unemployed individuals:[18]

> *What are the chances that in the next 12 months you will keep your job or find one (or start a new activity)? In other words, if you were to assign a score between 0 and 100 to the chance of keeping your job or of finding one (or of starting a new activity), what score would you assign? ("0" if you are certain not to work, "100" if you are certain to work).* [A graphic scale going from 0 to 100 is shown to the respondent.]

In this paper, we use the complementary probability, namely, the probability of unemployment.

As described in Guiso *et al.* (2002), the full sample of individuals who were asked this question in 1995 contains 4799 individuals, which become 4205 after non-respondents (8%) are excluded. Note also that those who expected to voluntarily retire or drop from the labor force are excluded. Their answers attest to the high degree of job security enjoyed by workers in Italy: the 4th decile is zero, the median is 30%, a 50% chance of unemployment is reached only in the 8th decile, and only 3% of individuals are certain to be unemployed in the year following the interview.[19] The authors also compare this source, restricting the sample to those employed, with the Survey of Economic

[18]Note that those who answer "yes" to the question *"Do you expect to voluntarily retire or stop working in the next 12 months?"* are not asked this question.

[19]The authors point out that it is not clear if employed respondents report only involuntary job losses or any change in employment status (including job mobility).

Table 4: The indicator of perceived job insecurity: observed and imputed with the multiple imputation method

Value of the indicator	Observed		Observed and imputed	
	Percent	Cumul.	Percent	Cumul.
0.0	60.13	60.13	45.88	45.88
0.1	11.27	71.40	12.22	58.11
0.2	7.93	79.33	11.70	69.81
0.3	1.88	81.21	7.83	77.64
0.4	2.71	83.92	6.71	84.34
0.5	4.80	88.73	5.95	90.30
0.6	0.63	89.35	2.80	93.10
0.7	2.09	91.44	2.12	95.22
0.8	2.71	94.15	1.96	97.18
0.9	1.88	96.03	1.02	98.20
1.0	3.97	100.00	1.80	100.00
Total	100.00		100.00	

Distribution of the indicator of job insecurity of fathers in the sample 479 observations used in the econometric analysis (see Table 3). The indicator measures the probability assigned by the individual to the event that he does not work in the following year. Job insecurity imputations reported in the right-hand side of the table were estimated using the multiple imputation method. The column 'Observed' refers to the 479 observations for which information on father's perceived job insecurity is non-missing. The column "Observed and imputed" refers to all 1142 observations in the sample.

Expectations (SEE), which contains a similar question for employed workers in the US.[20] While in Italy 59% of individuals report a zero chance of unemployment, in the US only 31% do so. The cumulated fraction of respondents for each probability of unemployment is systematically lower in the US than in Italy up to a 10% probability (at the 7th decile), after which it becomes similar.

The first two columns of Table 4 report the distribution of the perceived uncertainty indicator for the fathers of our sample of 479 households in Table 3. It is easy to see that, as expected given the sample design, our sample is not very different from the full sample

[20] *"I would like you to think about your employment prospects over the next 12 months. What do you think is the percent chance that you will lose your job during the next 12 months?"*. See Manski and Straub (2000).

used by Guiso *et al.* (2002). In our case, the average perceived unemployment probability of fathers is slightly smaller (20% vis-à-vis 22%) but this makes sense, since in our sample individuals are older (they must have a child of working age) and the perceived probability of unemployment drops with age. For the same reason, the probability mass at zero expected unemployment is also lower (53%).

As previously explained, we cannot use the same question of the SHIW to obtain an indicator of perceived uncertainty for children. Thus we simply include a measure of labor market tightness for children among the controls. Specifically, we use unemployment rates at the province level (separate by gender and age class) obtained from Quarterly Labor Force statistics. In the case of fathers, who are typically characterized by employment rates which are very high and constant across provinces and ages, we are lucky to have a subjective measure of job insecurity, because unemployment rates would be less informative on the individual-specific degree of insecurity faced by prime-age males. In the case of children, provincial unemployment rates by age and gender have the potential to capture their labor prospects adequately. Moreover, in the case of children the subjective perception of the likelihood of being employed in the future may reflect a labor supply decision and would therefore be endogenous for the purposes of our analysis (i.e. it might capture the following sequence: "I have decided not to work and thus I expect that I will not have a job"). This is less likely to be a problem in the case of fathers, for whom the expectation of future unemployment is more likely to be due to a perceived exogenous shock more than to an increase in the preference for leisure.[21]

Finally, as far as the outcome variable is concerned, it is a dummy variable taking the value 1 if the child left the household between 1995 and 1998 and it is described in the last row of Table 3. Of the 1142 children living with their parents in 1995 on which our analysis is based, only 40 (4%) decided to leave home in the following three years (21 females and 19 males). Proportions are similar in the subset of 479 households for

[21] In other words, we do not think it is likely, given sample participation rates for fathers, that the father would stop working in order to make the child leave home.

which paternal perceived uncertainty is available.

4.3 Imputations of perceived job insecurity

It is unfortunate that the SHIW asks the question concerning job insecurity only to half of the potential sample, but at least this half is selected randomly on the basis of the year of birth (odd or even) of the household head. Thus, using the terminology of the literature on missing data (see, for example, Little and Rubin, 1987), we can safely say that the missing information is "missing completely at random". We are therefore in a relatively good position to increase the size of our sample, by imputing the missing data on the basis of the available information using the procedure known as "Multiple imputation".[22]

Multiple imputation is a procedure based on the replacement of each missing value with a vector of $M \geq 2$ simulated values. Within a Bayesian framework, these simulated values are obtained as random draws from the posterior distribution of the missing data given the observed data and some prior distribution. As a result, M imputed data sets are obtained, one for each of the multiple imputations. Therefore, in each of these datasets the observed information is the same while the imputed information differs. The main analysis of interest is then conducted in each of the M imputed datasets, producing M estimates of the coefficient of interest. Denoting by $\hat{\beta}_i$ the estimate obtained from the *i-th* dataset, Rubin (1987) shows that the M estimates can be aggregated into a single estimate, defined as

$$\bar{\beta}_M = \sum_i^M \frac{\hat{\beta}_i}{M}. \tag{12}$$

The variability associated with this estimate is the sum of two components: the average within-imputation variance,

$$\bar{W}_M = \sum_i^M \frac{\hat{W}_i}{M} \tag{13}$$

[22]See Rubin (1987) and more recently Rubin (1996).

where $\hat{W}_i$ is the estimated variance of $\hat{\beta}_i$, and the between-imputation component

$$\bar{B}_M = \sum_{i}^{M} \frac{(\hat{\beta}_i - \bar{\beta})^2}{M-1}. \tag{14}$$

The total variability associated with $\bar{\beta}_M$ is therefore

$$T_M = \bar{W}_M + \frac{M+1}{M}\bar{B}_M \tag{15}$$

where $(M+1)/M$ is an adjustment for finite M. For significance tests and interval estimates, the estimand β is distributed according to a t distribution with mean $\bar{\beta}_M$, variance T_M, and degrees of freedom approximated by

$$\nu = (M-1)\left[1 + \frac{1}{M+1}\frac{\bar{W}_M}{\bar{B}_M}\right] \tag{16}$$

The results reported in the next section are obtained with the strategy described above on the basis of $M = 5$ multiple imputations. As shown in Rubin (1987), p. 114, this small number of imputations is sufficient, under normal circumstances, to achieve efficiency.[23] The last two columns of Table 4 report the distribution of the indicator of perceived uncertainty for the full sample of 1142 fathers, where the indicator has been imputed with the procedure described above for the 663 observations for which it was missing at random. The imputed distribution displays a lower frequency of zero uncertainty but otherwise appears fairly similar to the observed distribution.

4.4 Results

Table 5 reports estimates of the marginal effects of our measures of job insecurity for fathers and children on the probability that children leave home within three years from the baseline. Ideally, we would like to base these estimates on a comparison of children who share all relevant personal and family characteristics potentially affecting the outcome, so that we can isolate the effect of the treatment, job insecurity. We try

[23] A free-ware software package written by Joe Schafer (Department of Statistics, Pennsylvania State University) available at http://www.stat.psu.edu/jls/misoftwa.html has been used to generate the M imputed datasets and to aggregate the estimates. This software assumes a normally distributed model.

to approximate this ideal condition by controlling for a large set of variables dated in 1995, when all children are observed coresiding, or before. For children, we control for their age, gender, and completed years of schooling.[24] Emancipation decisions are likely to be affected by both family traits and the current situation in the household. So, we condition on the father's age and completed years of schooling.[25] We also control for family wealth, home-ownership (owner-occupied = 1), number of children present in the household, and region of residence (through 5 dummy variables for North-West, North-East, Center, South and Islands, with the North-West becoming the reference). Finally, since the availability of affordable housing plays a crucial role in the decision to move out, we include as controls also housing price indicators at the province level.[26] Table 5 does not report estimates for all these controls, however, in order to save space and focus attention on the marginal effects of interest.[27]

The first row of Table 5 displays estimates based on the restricted sample of 479 observations for which the uncertainty perceived by fathers is observed. Using a linear probability model (first column), if the job insecurity perceived by a father goes from zero (sure to be employed next year) to 1 (sure to be unemployed next year), the probability that the child leaves home increases by 9.3 percentage points (remember that the average probability of emancipation in the sample is 0.04). Despite the small sample size, this estimate is significantly different from zero and relatively large in absolute size. A probit model (third column) produces a more conservative estimate of the marginal effect of

[24]The child's schooling is likely to be endogenous with respect to her residential choice, and so our preferred regression excludes schooling. Including this variable, however, would allow us to control in part for the child's expectations concerning the level of her future income, a variable the model predicts as relevant in the residence choice. For this reason, we performed robustness checks by including schooling as a regressor, with unchanged results.

[25]Notice that, to the extent that father's age and schooling control for the father's income level when employed, and since unemployment benefits are proportional to previous wages in Italy, the degree of perceived job insecurity measures (the complement of) the probability that the parent will get his full wages, as opposed to the corresponding unemployment benefits. For this two-point support distribution of parental income (employment wages versus unemployment benefits), a reduction in perceived job insecurity exactly captures the notion of first-order stochastic dominance used in the model.

[26]These data have been purchased from Nomisma S.p.a. We also tried including a dummy capturing whether the household owned a second house, but it was not significant.

[27]The full set of results is of course available from the authors.

Table 5: The effect of the father's job insecurity on the child leaving

	OLS	Probit
Non-imputed	0.093	0.036
	(0.033)	(0.017)
Imputed sample 1	0.073	0.046
	(0.023)	(0.015)
Imputed sample 2	0.092	0.054
	(0.022)	(0.014)
Imputed sample 3	0.131	0.064
	(0.023)	(0.015)
Imputed sample 4	0.107	0.056
	(0.023)	(0.015)
Imputed sample 5	0.099	0.057
	(0.023)	(0.014)
Aggregate	0.100	0.055
	(0.033)	(0.016)

The first column displays the coefficient of a linear probability model. The second column shows the marginal effects of a probit model. The first row shows results on the non-imputed subsample (479 observations). Rows 2-6 show the results on the five multiple imputation samples described in Table (4) (1142 observations). Row 7 shows aggregate estimates and standard errors based on Rubin's formula as described in the main text. Numbers in brackets are standard errors.

a similar change of paternal insecurity (3.6 percentage points), but the estimate is still significantly different from zero and large with respect to the starting point.

The next five rows of the Table display the results obtained in five different samples in which the insecurity perceived by parents has been imputed using the procedure described in Section 4.3. In all these samples, the estimated marginal effect is large and statistically significant for both the linear probability and the probit models. When these five estimates are aggregated using equations (12) and (15), the results obtained in the restricted sample are confirmed and, actually, the estimated marginal effects are slightly higher, particularly in the case of the probit model.

We conclude that, as far as paternal job insecurity is concerned, the insight of our theoretical model is fully confirmed. A child is significantly more likely to leave home when his/her father perceives that the possibility of not having a job in the future is high.

The variable controlling for labor market tightness facing children, namely the provincial unemployment rate, yields the following estimated coefficients (standard errors): 0.121 (0.083) for the linear probability model, and 0.075 (0.047) for the probit. Thus the marginal effects are estimated to be positive but not significantly different from zero. This result probably reflects the fact that provincial unemployment rates, even if disaggregated by age and gender, are not sufficiently informative on the degree of job insecurity faced or perceived by young adults. In other words, more than rejecting the prediction our model concerning job security of children, we see this result as inviting us to suspend the verdict until an informative measure of perceived job insecurity for children is available.[28]

[28]For completeness, we report that the number of children in the household is a statistically insignificant regressor. The model predicted an ambiguous effect of this variable on the residence choice of the child, as shown in Becker *et al.* (2004).

5 Conclusions

In this paper, we have explored one potential determinant of the youth's decision to leave their parents' home. Our key insight was that it may depend on the degree of job security of parents and children. To address this issue, we presented a theoretical model. It predicted, under certain conditions, that higher own expected income induces children to leave the parental home earlier, while higher expected parental income has the opposite effect.

We have shown that the aggregate evidence for 12 European Union member countries in the 1980s and 1990s on emancipation rates, for the 20-24 and 24-29 year-old brackets, and on perceived job insecurity is consistent with this hypothesis. According to our estimates, if the percentage of youths feeling that their job is insecure went from 0 to 100, the coresidence rate would have increases by 13 percentage points, while the same change in the percentage of workers aged 50-59 y.o. would have reduced coresidence by 9 percentage points. We read this evidence as suggesting that job security is a consistent explanatory variable of coresidence decisions across countries, once adequate control for differences in institutions and culture is performed.

We were further able to validate the job security hypothesis using microeconomic panel data from the Italian Survey of Household Income and Wealth (SHIW), collected by the Bank of Italy. We estimated linear probability and probit models controlling for children's, father's and family characteristics, and imputing perceived job insecurity for a fraction of the individuals in the sample. With this sample, however, we could not use a measure of perceived job insecurity for children. Our empirical results indicate that the likelihood young Italians aged 18 to 35 years old left the parental home between 1995 and 1998 is positively related to parental job insecurity, measured in terms of their father's own perception of his chances of having a job over the subsequent 12 months. Going from a perception of a fully secure job to a perception of a certain job loss (or of the inability to find a new job) raised the likelihood of the child leaving by 4 to 10 percentage points, depending on the sample and estimation method used. This is a quite substantial increase given that the average probability of emancipation observed in the sample is 4%.

Having established the quantitative importance of the effects of perceived job security on coresidence, and given that labor market institutions are important determinants of the relative job insecurity of parents and children, our results uncover an empirically significant link between labor market institutions –labor market rigidity in particular– and family demographics.

What are the policy implications of our analysis? The main direct effects of late youth emancipation are low fertility, which for instance puts in jeopardy pension systems in southern European countries, and low geographical mobility, leading to a lower capacity to react to idiosyncratic regional shocks. Both of these problems are constantly debated in southern European countries but we are the first to link them to the problem of coresidence through the effect of job security provisions. Coresidence also has benefitial implications: society as a whole may gain from it if parents can monitor the job search activities of their children better than public employment agencies, and thus decide on the size of the provision of "unemployment benefits" within the family. In conclusion, what is socially desirable as far as these outcomes are concerned is debatable, but, whatever the result of this debate, our analysis shows that it should also take into account the effects of job security provisions for parents and children on youth emancipation.

A The effect of income distribution functions on the period 1 moving-out threshold

Generalization of the results on first-order stochastic dominance. In this section, we discuss how to generalize Propositions 6 and 7. We begin by defining the set $\mathcal{F}$ more formally.

Let Ω_F denote the set of all pairs of independent distribution functions (F_c, F_p) such that F_c has support over $[\gamma_c, \infty)$ and F_p has support over $[\gamma_p, \infty)$. We now define a convenient subset $\mathcal{F}$ of Ω_F:

$$\mathcal{F} = \left\{ \begin{array}{c} (F_c, F_p) \in \Omega_F : F_c\left(y_{c2}\right) F_p\left(y_{p2}\right) = 0 \text{ for } \left(y_{c2}, y_{p2}\right) \\ \text{such that } y_{c2} < \tilde{y}_{c2}\left(y_{p2}\right) \end{array} \right\}.$$

In words, the distributions F_c and F_p in $\mathcal{F}$ are independent and assign mass zero to the strict subset of the regret region where transfers would be granted to children living on their own.

The condition $\alpha < 1$ invoked in propositions 6 and 7 could be generalized by imposing that the sharing effect dominate the marginal utility effect for income values such that the second-period residential choice would be independence. This would ensure that the function $\Delta_2\left(y_{c2}\right)$ is monotonically increasing to the right of $\bar{y}_{c2}$; consequently, the income threshold $\bar{y}_{c1}$ for which the child is exactly indifferent between moving out and staying at home in period 1 would be the unique solution to $\Delta_2\left(y_{c1}\right) = \bar{R}$; further, it would also ensure that $\bar{y}_{c1}$ is monotonic in $\bar{R}$.

How could propositions 6 and 7 be generalized, allowing F_p and F_c to place mass on the entire $\left(y_{c2}, y_{p2}\right)$ plane? Consider first the child's income. As argued in Lemma 1, the ambiguity of the slope of $\Delta_2\left(y_c\right)$ when transfers are positive hinges on the dominance of the sharing versus the marginal utility effect. If, as argued then, we assume that the marginal utility effect dominates for $y_{c2} \in \left[\gamma_c, \tilde{y}_{c2}\left(y_p\right)\right]$, then $\Delta_2\left(y_{c2}\right)$ is positively sloped and strictly increasing for y_{c2} anywhere in the regret region.[29] The next lemma makes this point formally.

Lemma 8 *Let (F_p, F_c^1) and (F_p, F_c^2) be two elements of Ω_F, and assume that F_c^1 first-order stochastically dominates F_c^2. Let the period 1 moving-out threshold corresponding to F_c^j be denoted $\bar{y}_{c1}\left(F_c^j\right)$. Then, if the sharing effect dominates the marginal utility effect for $y_{c2} \leq \tilde{y}_{c2}\left(y_p\right)$ and $\alpha < 1$, $\bar{y}_{c1}\left(F_c^1\right) \leq \bar{y}_{c1}\left(F_c^2\right)$.*

While the prevalence of the marginal utility effect will not allow a similar generalization for the parent's income distribution,[30] a continuity argument helps us analyze the

[29] Recall that one instance where the marginal utility effect prevails over sharing is the case where the parent values child consumption just as much as his, $\lambda = 0.5$.

[30] Altruism imposes the equality of $\partial\Delta_2/\partial y_{c2}$ and $\partial\Delta_2/\partial y_{p2}$ in the range of positive transfers. However, for the subset of income values such that the parent does not provide financial help but the child still prefers to be at home ($y_{c2} \in [\tilde{y}_{c2}, \bar{y}_{c2}]$), $\partial\Delta_2/\partial y_{c2}$ and $\partial\Delta_2/\partial y_{p2}$ must have opposite signs. The lack of transfers causes y_{p2} to benefit the child only if she is at home ($\partial\Delta_2/\partial y_{p2} < 0$), whereas higher y_{c2} is always more beneficial to an independent child ($\partial\Delta_2/\partial y_{c2} > 0$). This implies that, when $\Delta_2\left(y_{c2}\right)$ is increasing everywhere in the regret region, $\Delta_2\left(y_{p2}\right)$ must have an increasing and a decreasing range in that region.

case when F_p places mass over $[\gamma_p, \infty)$. What is required is that, for given values of the child's period 2 income, the changes in $\bar{R}$ associated with different distributions F_p be driven by values of Δ_2 in the no-transfer region. In other words, the density placed by different distributions F_p on y_{p2} values above $\tilde{y}_{p2}(y_{c2})$ cannot be too large.

Second-order stochastic dominance. Let us address the impact of higher variance of the future income streams of parent and child on the moving-out decision. We compare the variance of income under different distributions while holding the expected value constant. That is, using the child's income as an example, we say:

Definition 9 *$F_c^1(y)$ dominates distribution $F_c^2(y)$ in the second-order stochastic sense if:*

i) $\int yF^1(y)\,dy = \int yF^2(y)\,dy$, and

ii) $\int_{\gamma_c}^{y_c}\left[F^1(z) - F^2(z)\right]dz \leq 0$, with the inequality holding for all y_c in the domain of the child's income.

Proposition 10 *Let (F_p, F_c^1) and (F_p, F_c^2) be two elements of $\mathcal{F}$ such that F_c^1 dominates F_c^2 in the second-order stochastic sense. Then, $\bar{y}_{c1}(F_c^1) \leq \bar{y}_{c1}(F_c^2)$.*

A similar result can be obtained for parental income. However, since changes in the distribution of income affect the period 1 moving out threshold only to the extent that they affect period 2 income values within the regret area, for the parent's income we need to ensure that the distribution of his income shifts so as to raise the variance of income values in that region specifically, as opposed to the requirement that it becomes more volatile over its global range. That is, if F_p^1 dominates F_p^2 in the second-order stochastic sense, we need additionally to impose that, for all values of the child's income y_{c2},

$$\int_{\gamma_p}^{\bar{y}_{p2}(y_{c2})} F_p^1(y_{p2})\,dy_{p2} = \int_{\gamma_p}^{\bar{y}_{p2}(y_{c2})} F_p^2(y_{p2})\,dy_{p2} = 0, \tag{17}$$

which, together with the conditions that ensure the second-order stochastic dominance of F_p^1 over F_p^2, implies:

$$\int_{\bar{y}_{p2}(y_{c2})}^{y_{p2}} \left[F_p^1 - F_p^2\right] dy_{p2} \leq 0,$$

for all $y_{p2} \geq \bar{y}_{p2}(y_{c2})$.

Proposition 11 *Let $\left(F_p^1, F_c\right)$ and $\left(F_p^2, F_c\right)$ be two elements of $\mathcal{F}$ satisfying equation (17), such that F_p^1 dominates F_p^2 in the second-order stochastic sense. Then, $\bar{y}_{c1}\left(F_p^1\right) \geq \bar{y}_{c1}\left(F_p^2\right)$.*

Proposition 10 says that a lower variance in the child's future income stream makes her less reluctant to leave the parental home. We get the opposite reaction concerning parental income: the child demands a higher income level in period 1 in order to move out if the parent faces a lower variance income stream.

B Description of macroeconomic data

Coresidence rate. Fraction of population living at parental home. Source: European Labour Force Survey. Countries: all in the EU. Years: 1983-2002, though data start later for new EU members.
Perceived job insecurity. 0-1 dummy variable constructed from answers to questions asked in the Eurobarometers (asterisks mark the answers considered as 1 in the dummy variable):

- April 1984: "During the last year, have you (or someone in your household) worried about losing a job or not finding a job?". Replies: 1. A lot (*). 2. A little. 3. Not at all.

- April-May 1992: "And in the future, how great a risk do you think there is that you will become unemployed?". Replies: 1. No risk. 2. Quite a low risk. 3. Quite a high risk (*). 4. A very high risk (*).

- March-April 1997: "How likely do you think it is that you may lose your job in the next few years?". Replies: 1. 0%, no risk at all. 2. 25%, low risk. 3. 50%, fifty-fifty (*). 4. 75%, high risk (*). 5. 100%, definitely will (*).

Countries available:

- 1984: West Germany, France, Italy, and United Kingdom. Data on Belgium and Ireland had missing values and had to be excluded.

- 1992 and 1997: Belgium, Western Germany, Greece, Spain, France, Ireland, Italy (not in 1992), Netherlands, Austria, Portugal, Finland, and United Kingdom.

Data on Luxembourg were excluded because several cells showed zero insecurity, due to the small number of observations included in the cells. We also exclude Italy in 1992 because the data show a very sharp drop in insecurity which then bounces back in 1997, again presumably due to the small number of observations.

We construct the job insecurity variable from 16483 individual observations for 4 countries in 1984 (3197 individual observations) and 12 countries in 1992 and 1997 (4982 and 5067 observations, respectively). They are constructed for cells by gender, age group (20-24, 25-29, and 50-59 years old), country, and year. To construct the cells, each individual observation is weighted by its population weight as given by the survey. We end up with 100 observations (16, 36, and 48, for the 3 years, respectively).

References

Banca d'Italia (1997), "I Bilanci delle Famiglie Italiane nell'anno 1995", Supplemento al Bollettino Statistico, Anno VII, Numero 14.

Banca d'Italia (2000), "I Bilanci delle Famiglie Italiane nell'anno 1998", Supplemento al Bollettino Statistico, Anno X, Numero 22.

Becker, S., S. Bentolila, A. Fernandes, and A. Ichino (2004) "Proofs of the Results in: 'Job Insecurity and Children's Emancipation: The Italian Puzzle"', ftp://ftp.cemfi.es/pdf/papers/sb/coresidence.pdf.

Böckerman, P. (2002), "Perception of Job Instability in Europe", mimeo, Labour Institute for Economic Research.

Díaz, A. and M.D. Guilló (2003), "Family Ties and Labor Supply", mimeo, Universidad Carlos III de Madrid.

Ermisch, J. (1996) "Analysis of Leaving the Parental Home and Returning to it Using Panel Data", ISER Working Paper 96-1.

Ermish, J. (1999), "Prices, Parents, and Young People's Household Formation", *Journal of Urban Economics* 45, 47-71.

European Communities (1999), *Demographic Statistics: Data 1995-98*, Luxembourg, Office for Official Publications of the European Communities.

Eurostat (2003), Eurobarometer (europa.eu.int/comm/public_opinion/).

Fogli, A. (2000), "Endogenous Market Rigidities and Family Ties", mimeo, New York University.

Ghidoni, M. (2002), "Determinants of Young Europeans' Decisions to Leave the Parental Household", mimeo, University College London.

Giannelli, G.C. and C. Monfardini (2003), "Joint Decisions on Household Membership and Human Capital Accumulation of Youths. The Role of Expected Earnings and Labour Market Rationing", *Journal of Population Economics* 16, 265-285.

Green, F., A. Felstead, and B. Burchell (2000), "Job Insecurity and the Difficulty of Regaining Employment: An Empirical Study of Unemployment Expectations", *Oxford Bulletin of Economics and Statistics* 62, 855-83.

Guiso, L., T. Jappelli and L. Pistaferri (2002) "An Empirical Analysis of Earnings and Employment Risk", *Journal of Business and Economic Statistics* 20, 241-53.

Little, R.J.A. and D.B. Rubin (1987), *Statistical Analysis with Missing Data*, Wiley.

Manacorda M. and E. Moretti (2003), “Why Do Most Italian Youths Live With Their Parents? Intergenerational Transfers and Household Structure”, mimeo, London School of Economics.

Manski, Ch.F. and J.D. Straub (2000), “Worker Perceptions of Job Insecurity in the Mid-1990s: Evidence from the Survey of Economic Expectations”, *Journal of Human Resources* 35, 447-79.

Martinez-Granado, M. and J. Ruiz-Castillo (2002), “The Decisions of Spanish Youth: A Cross-Section Study”, *Journal of Population Economics* 15, 305-330.

McElroy, M.B. (1985), “The Joint Determination of Household Membership and Market Work: The Case of Young Men”, *Journal of Labor Economics* 3, 293-316.

Organization for Economic Cooperation and Development (1997), *Employment Outlook*, Paris.

Rosenzweig, M.R. and K.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 Labour Economics* 11, 84-112.

Rubin D.B. (1987), *Multiple Imputation for Non-response in Surveys*, Wiley.

Rubin D.B. (1996), “Multiple Imputation after 18 years”, *Journal of the American Statistical Association* 91, 473-489.