Job Insecurity and Children's Emancipation: The Italian Puzzle*

Sascha O. Becker
CESifo

Samuel Bentolila
CEMFI

Ana Fernandes
CEMFI

Andrea Ichino
EUI

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Abstract

Why are Italian grown-up children so reluctant to leave the parental home? In this paper we present a theoretical model predicting that this behavior can be induced by high job insecurity of children and low job insecurity of their parents. Both aggregate data for EU countries and microeconometric results, using data from the Italian Survey of Household Income and Wealth for the 1990s, are broadly supportive of this hypothesis.

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1 Introduction

One of the most striking features of Italian society today is that grown-up children are very reluctant to leave their parents' home and to begin to work. For instance, in 1995 a staggering 86% of children aged 20 to 24 lived with their parents, and a remarkable 56% of those aged 25 to 29 still did so. Furthermore, these shares have been rising over time: for instance, the 1987 shares were 81% and 39%, respectively.

Table 1 provides an international perspective for this fact, by showing coresidence rates for youths aged 15-24 and 25-29 years old. The second column makes it clear that late emancipation is shared by Portugal, Greece, and Spain, but it is infrequent in the largest EU countries –Germany, France, and the UK– and almost inexistent in Scandinavian nations.

What determines youth decisions to leave the parental home? 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.[1] More specifically, several potential determinants of the late emancipation age in southern European countries have been studied in the empirical literature. For Italy, Manacorda and Moretti (2002) focus on the income of the parents, who are seen as bribing their children to stay at home longer, while Giannelli and Monfardini (2000) emphasize housing costs. Within an international investigation, Ghidoni (2002) finds that in Italy the child's unemployment situation is important for males, whereas parents and children's income matters for females. Martinez-Granado and Ruiz-Castillo (2002) also find an effect of housing costs for Spain. The Spanish case inspired Díaz and Guilló (2002) in proposing a theoretical model which stresses the mother's housework as a public good inducing children to stay.

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 by their parents. Fogli (2000) has recently 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 and young people tend to postpone the age of emancipation and entry to the labor force. According to Fogli's model, children remain with their parents –whose jobs are very secure due to labor market regulations– to enjoy 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, we will present a different model, in Section 3. We will show that for forward-looking children who are considering emancipation, 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 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 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

[1] See, for instance, McElroy (1985), Rosenzweig and Wolpin (1993), or Ermish (1996).

independent children. We also show that the joint consideration of parental and child job security is an 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. This hypothesis is motivated in Section 2 by discussing some aggregate evidence on coresidence and job insecurity in European Union (EU) countries in the mid-1990's.

To test the job security hypothesis we rely on the panel data structure of the Italian Survey of Household Income and Wealth (SHIW) collected by the Bank of Italy. Our particular sample and our empirical strategy are described in detail in Section 4. In all cases we estimate probit models for whether children live independently after a given year (1991, 1993, and 1995) as a function of indicators of parental and child job insecurity, and of a set of control variables measuring demographic, educational and labor market characteristics of the children and their fathers, and other family background indicators. Very few papers in this literature have exploited the panel data structure of available microeconomic datasets.

We start with a measure of "experienced" job insecurity, which captures the actual unemployment status of the male parent and of the child. The impact of this measure is examined for all three waves mentioned above. But our core empirical analysis is devoted to the 1995 wave of the SHIW, in which we have a subjective measure of job insecurity. In that wave, individuals were asked about their perceived likelihood of having a job in the subsequent 12 months. We use the answers to this question to construct an indicator of perceived job insecurity. This question was asked only to a randomly chosen 50% of the households, however, which leaves us with a too small sample. In order to increase sample size, we impute perceived job insecurity to non-asked individuals. To improve the reliability of the results, we implement both the traditional conditional mean imputation and also Rubin's (1987) multiple imputation.

At this stage the evidence is very preliminary. Our empirical results, discussed in Section 5, are however broadly consistent with the theoretical predictions. Higher job insecurity of children is found to deter emancipation, while the father's insecurity favors it. In Section 6 we present our preliminary conclusions.

2 Job insecurity of parents and children in the EU

As stated in the Introduction, we wish to test, for the Italian case, the hypothesis that child job insecurity has a negative effect on the probability that children leave their parents' home while parental job insecurity has a positive effect. Before doing so, it is worth checking whether the aggregate international evidence supports this idea.

For this purpose, we will use a measure of job insecurity for the 15 member countries of the EU, given by the percentage of employees in each age group who reply that they do not strongly agree with the statement "my job is secure" in 1996 (from the European Commission's *Eurobarometer*). Figure 1 shows this measure for two groups: workers aged 15-24 and those aged 45 and over.[2] Young workers feel most insecure in Spain,

[2] Legenda for Figures 1 and 2: BE Belgium, DK Denmark, DE Germany, GR Greece, ES Spain, FR

France, Greece, Portugal, and Italy, while older workers do so in France, the UK, the Netherlands, Finland, and Sweden. In five countries there is not much difference in job insecurity across age groups (Austria, Belgium, Denmark, Ireland, and Luxembourg).

On the other hand, in most countries young workers feel less secure than older ones, though this is not so in three countries with non-negligible differences across groups (Finland, the Netherlands, and the UK). The highest young-old insecurity ratios are found in Spain, Italy, Greece, Portugal, and France. Quite strikingly, the first four countries also show the highest coresidence rates (recall Table 1). Figure 2 makes this point.[3]

To probe the point further, Table 2 presents simple correlation coefficients between these variables. Clearly, coresidence is positively correlated with the job insecurity perceived by young workers and negatively correlated with that of older workers, so that it is also positively correlated with the ratio of perceived insecurity of the young over that of the older group. Indeed, the 0.74 coefficient for 25-29 years old group is quite remarkable for just 15 observations.

One may think that unemployment rates should capture the same effect as job insecurity. The bottom half of the Table shows this is not fully so. The correlation coefficients of coresidence with unemployment rates are, except once, lower than those with job insecurity. This may be because unemployment is measured in a more heterogeneous way across countries than insecurity or because the age brackets available for insecurity are better suited to capture the group of parents of young workers than those for unemployment. But there are potentially meaningful reasons for this fact: perceived insecurity may matter more than actual unemployment. Moreover, insecurity is intimately related to the risk of job loss, which may be high even if unemployment is low, as can happen with the fixed-term contracts that prevail among young workers in several EU countries.

The risk of job loss is intimately related with job security provisions, which are common in European countries and tend to favor older workers, at the expense of the employment prospects of younger ones. For instance, in OECD panel data regressions employment protection legislation is found to raise the unemployment rate of young women in Jimeno and Rodriguez-Palenzuela (2002) and to lower the employment-population ratio of young women (vis-à-vis prime-aged women) in Bertola *et al.* (2002) (while for young men the effects are not statistically significant). Regarding our variables, related, but rougher, evidence is provided by the simple correlation coefficients between a measure of employment protection in 1990-1995 from Blanchard and Wolfers (2000)[4] and our various measures of perceived job insecurity. They take the following values: 0.51 for the young (16-24 y.o.), 0.39 for the prime aged (25-44 y.o.), -0.14 for the older group (45 y.o. and more), and 0.48 for the relative insecurity of the youngest to the oldest group. This suggests that young workers feel more insecure the more

France, IE Ireland, IT Italy, LU Luxembourg, NL Netherlands, AT Austria, PT Portugal, FI Finland, SE Sweden, and UK United Kingdom.

[3]Spain, which has the highest values for both variables and therefore fully conforms to the pattern, has been excluded from the graph because its relative job insecurity value (2.1) is an outlier.

[4]More precisely, their "new measure of employment protection" or *newep*.

protected are older workers. In this connection, it is useful to recall that Italy has a highly regulated labor market. For instance, Italy ranks second among EU countries in terms of the degree of stringency of the protection of regular workers against dismissals according to the OECD (Grubb and Wells, 1993) and ranks third in the Blanchard and Wolfers (2000) indicator just mentioned.

In sum, while being quite rough, the aggregate evidence for European countries suggests that youth job insecurity, in particular relative to their parents', might contribute to explain why youth coresidence is high in Italy. Before testing such hypothesis more rigorously, we proceed to derive it formally from a theoretical model.

3 A model of job insecurity and coresidence

The purpose of this section is to illustrate with a model why and how coresidence decisions are related to job insecurity of parents and children. With this aim, we present a dynamic, two-period model of residence choice.

3.1 The family

The family in our model has n_0 altruistic parents and $n_1 + 1$ children. We assume that a family has either one or two living parents and at least one child: $n_0 \in \{1, 2\}$ and $n_1 \geq 0$. Family size is denoted by n. Our focus is on the residence choice of one of the children, assuming that the remaining n_1 siblings remain with their parents.

Individuals in our model care only about consumption[5]. We assume that, in the parental home, all individuals pool income and consume an equal fraction of total family income. We root the latter assumption on the difficulty experienced by altruistic parents in excluding children from the consumption of public goods in the household.

Suppose that all n family members are coresiding. Then, consumption in the parental home, c_p^n, is given by:

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

where γ_p is the rent or the imputed cost of housing, p a price index of a representative basket of consumption goods, y_p the sum of parental incomes, and y_c the sum of children's incomes. Given our focus, we will in fact regard y_c as the income of the child who is contemplating to move out (the remaining children are therefore assumed to earn no income).

We denote the child's consumption by c_c. If she stays with the parents, she gets c_p^n. If she moves out, she will consume all of her income plus a non-negative transfer t from her parents. Her consumption while independent is therefore given by

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

[5] We will later address some of the implications of the value of leisure for our results.

where γ_c is rent or the imputed cost of housing for the child.

On the other hand, consumption at the parent's home, denoted c_p^i, is now:

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

We assume that parents value their direct utility from consumption more than their children's. This partial altruism is captured by a weight $\lambda \in (0.5, 1)$ on parents' direct utility, while their children's consumption gets weight $(1-\lambda)$. Parental utility U_p is then

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

where c_c is equal to c_p^n if the child coresides and to c_i if she becomes independent, while c_p is equal to c_p^n if the child coresides and to c_p^i if she becomes independent.

In what follows, we will in fact use a slightly modified functional form for parental utility:

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

Relative to (1), the functional form in (2) puts more weight on the utility of parents and the n_1 children who always remain at home. We adopt (2) since this transformation leaves all the results of the analysis qualitatively unchanged while it simplifies the algebra significantly.

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 $\alpha > 0$.

3.2 Timing

As stated earlier, our problem is cast in a two-period framework. We begin by describing the timing of events in period 1. First, parent and child observe their income realizations, y_{p1} and y_{c1}.[6] For convenience, we will assume that there is a lower bound on income realizations, corresponding to the housing costs, γ_p and γ_c. This ensures that consumption of family members is always non-negative, irrespective of residence choices. A positive income realization for the parent, interpreted as a draw of y_{p1} strictly exceeding γ_p, is equivalent to a job offer, and similarly for the child. Given that there is no disutility from work, job offers are always accepted[7]. The child then decides whether or not to move out. Finally, consumption takes place as a function of the residential choice of the child, as explained above.

The main difference across periods comes from assuming that once the child has left the parental home, she cannot move back in. We can justify this irreversibility

[6] The second subscript always refers to the period.

[7] A positive income threshold beyond housing costs would be required before accepting a job offer if individuals experienced disutility from work or if they were productive while unemployed (through household production, say). We are ignoring these cases.

assumption on the grounds that the costs of moving moving back in are typically very high. In any event, qualitatively similar results would emerge from considering finite moving costs instead.

For a child who stayed with her parents in period 1, the period 2 timing of events and choices are identical to those of period 1. The same is true if the child has moved out in period 1, but there is one fewer decision to be made: the child will remain independent in period 2.

3.3 Period 2

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

$$\max_{\tilde{t}_2 \geq 0} \left\{ \lambda (n-1)\, u \left(\frac{y_{p2} - \tilde{t}_2 - \gamma_p}{p (n-1)} \right) + (1-\lambda)\, u \left(\frac{y_{c2} + \tilde{t}_2 - \gamma_c}{p} \right) \right\}. \quad (3)$$

First-order conditions yield:

$$\lambda u' (c_{p2}) \geq (1-\lambda)\, u' (c_{i2}),$$

holding with equality when positive transfers are provided. Since $\lambda > 0.5$, this condition says that the child has lower *per capita* consumption than the remaining family members when she receives transfers. Under CRRA preferences, positive transfers are given by:

$$t_2 = \frac{y_{p2} - \gamma_p - (y_{c2} - \gamma_c)\, \Gamma (n-1)}{1 + \Gamma (n-1)}, \quad (4)$$

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

As is common in the altruism literature, transfers from altruistic parents are increasing in parental income and decreasing in the child's income (both net of housing costs). The preference specification in (2) also ensures that, for concave utility functions, transfers depend negatively on family size n.

Using the transfer function in (4), we can solve for the consumption of the child when independent and receiving transfers:

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

whereas at the parental home consumption of the remaining family members is:

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

We note that, given the chosen functional form for preferences, for y_{c2} values such that transfers would be given to an independent child, the child is better off not moving out, if she still has that option left:

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

We now address the moving out decision for those who decided to stay home in period 1. Let us define Δ_j as the excess utility level when independent relative to coresiding, for period j:

$$\Delta_j\equiv u\left(c_{ij}\right)-u\left(c_{pj}^{n}\right);\quad j=1,2.$$

The child moves out if $\Delta_2>0$. (In the indifference case, $\Delta_2=0$, we assume she stays.)

How does Δ_2 change as a function of the child's income? Let us define the following important values. First, $\tilde{y}_{c2}$ denotes the value of the child's income such that parental transfers are zero:[8] $t_2\left(\tilde{y}_{c2}\right)=0$. Second, $\bar{y}_{c2}$ is the child's income value that leaves her indifferent between staying home or moving out: $\Delta_2\left(\bar{y}_{c2}\right)=0$.

Under CRRA preferences,

$$\tilde{y}_{c2}=\frac{y_{p2}-\gamma_p}{\Gamma\left(n-1\right)}+\gamma_c. \tag{5}$$

Note that $\tilde{y}_{c2}$ and γ_c coincide when $y_{p2}=\gamma_p$ (the case of an unemployed parent).

Given the transfer function, we know that, for income values y_{c2} such that transfers would be positive, the child's consumption is lower if she were to become independent compared to staying home ($\Delta_2<0$). Consequently, the income level $\bar{y}_{c2}$ that sets Δ_2 to zero must exceed $\tilde{y}_{c2}$. It equates consumption at the parental home with consumption while independent:

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

or

$$\bar{y}_{c2}=\frac{y_{p2}-\gamma_p}{n-1}+\frac{n}{n-1}\gamma_c>\tilde{y}_{c2}. \tag{6}$$

We now characterize formally how Δ_2 depends on the child's income y_{c2}.

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

Proof. *See Appendix 1.*

[8] The notation $t_j\left(x\right)$ emphasizes the dependence of function t_j with respect to variable x while omitting, for simplicity, other arguments of the function. This type of simplification will be used throughout the section.

Lemma 1 describes how the excess utility from moving out depends on the child's income. We now discuss the intuition underlying these results. For a child who did not move out in period 1, her residence choice is clear: move out if and only if her income exceeds $\bar{y}_{c2}$. Lower income values discourage changing residence since coresidence allows the child to share family resources more productively (from the point of view of generating utility) than independence. Note that even though low y_{c2} values trigger transfers ($y_{c2} \in [\gamma_c, \tilde{y}_{c2}]$), the transfers are not enough to compensate the child for the loss in income that moving out and paying a new rent entails.[9]

A child who did move out in period 1 would like to reverse her decision when y_{c2} falls below $\bar{y}_{c2}$. In fact, the set of income values $[\gamma_c, \bar{y}_{c2}]$ can be thought of as a *regret region* for children who moved out in period 1: they would have been better off staying home rather than on their own.

Lastly, 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. To start with, income modifies consumption differently depending on the residence state. For example, for y_{c2} values such that no transfers would be provided to the child (i.e. to the right of $\tilde{y}_{c2}$), higher y_{c2} implies that c_{i2} is changing by $1/p$ whereas consumption at home goes up by $1/np$. Thus consumption while independent is increasing faster with y_{c2} than consumption at home. This is a *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} is greater than c^n_{p2}, the marginal utility of consumption at home is higher than that of consumption while independent. This *marginal utility effect* counteracts the higher change in c_{i2} relative to c^n_{p2} induced by changes in y_{c2}. 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, the marginal utility effect and the impact of y_{c2} on consumption go in the same direction, ensuring that Δ_2 is positively sloped.

The relationship $\Delta_2 (y_{p2})$, describing how the residence choice depends on parental income is also interesting. Before proceeding to characterize it, let us define $\bar{y}_{p2}$ as the parental income level that leaves the child indifferent between moving out and coresiding: $\Delta_2 (\bar{y}_{p2}) = 0$. Similarly, let $\tilde{y}_{p2}$ be the level of parental income such that $t_2 (\tilde{y}_{p2}) = 0$. It is straightforward to show that $\bar{y}_{p2} < \tilde{y}_{p2}$. Then,

Lemma 2 *The function $\Delta_2 (y_{c2})$ is strictly decreasing for $y_{p2} \in \left[\gamma_p, \tilde{y}_{p2}\right)$ 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.*

Proof. *Similar to previous one.*

Concerning $\Delta_2 (y_{p2})$, whether or not $\Delta_2 \left(\gamma_p\right)$ is positive depends on the parameter values (specifically, a large number of family members n and a small rental cost γ_c

[9] This result hinges partly on the preference specification in the sense that, with other utility function, parents might raise the independent child's consumption by more than the coresiding child's in response to changes in income. If so, it would be possible to have a positive value of Δ_2 for low values of y_{c2}.

for the independent child 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 preserved the option of staying home will not move out, whereas those who became independent in period 1 would now like to reverse their decision. Just as with $\Delta_2(y_{c2})$, it is also the case here that higher parental income does not necessarily raise the child's willingness to stay home. In fact, for $y_{p2} > \tilde{y}_{p2}$, the slope of $\Delta_2(y_{p2})$ depends on the preference specification. The intuition is the same as before: even though higher parental income causes consumption at home to increase by more than the consumption of an independent child, the utility gain from higher consumption could be larger in the counterfactual situation of independence (i.e. the sharing vs. the marginal utility effect).

In Figure 3 we plot the curves $\tilde{y}_{c2}$ and $\bar{y}_{c2}$ in (y_c, y_p) space. The schedule $\bar{y}_{c2}(y_{p2})$ separates the (y_{c2}, y_{p2}) plane into the moving-out area, to the right of $\bar{y}_{c2}(y_{p2})$, from the coresidence area, to the left of the curve. From the point of view of the moving-out decision taken in period 1, we can also divide the coresidence area in two parts. At income pairs to the left of $\tilde{y}_{c2}(y_{p2})$, children who became independent in period 1 will receive financial help from their parents, i.e. $t_2(y_{c2}, y_{p2}) > 0$. On the other hand, children belonging to families with income pairs to the right of $\tilde{y}_{c2}(y_{p2})$ (but still inside the regret area) will not be helped by their family. It is important to recall that from this point of view, the coresidence area is a *regret area:* in it, children who left would have preferred, ex-post, to have stayed in the parental home.

We end the discussion of Period 2 by characterizing how Δ_2 depends on family size. We want to compare Δ_2^n with $\Delta_2^{n^+}$, where $n^+ > n$. We have that:

Lemma 3 *$\tilde{y}_{c2}^{n+}(y_{p2}) < \tilde{y}_{c2}^{n}(y_{p2})$ for $y_{p2} > \gamma_c$ and $\bar{y}_{c2}^{n+}(y_{p2}) < \bar{y}_{c2}^{n}(y_{p2})$ for all y_{p2}. Further, $\Delta_2^{n+}(y_{c2}) > \Delta_2^{n}(y_{c2})$ for all y_{c_2}.*

Proof. *See Appendix 1.*

In (y_{c2}, y_{p2}) space, a larger family size makes the schedule $\tilde{y}_{c2}^{n+}(y_p)$ steeper relative to $\tilde{y}_{c2}^{n}(y_{p2})$. The schedule $\bar{y}_{c2}^{n+}(y_{p2})$ is also steeper relative to $\bar{y}_{c2}^{n}(y_{p2})$ and it has a smaller intercept. That is, $\bar{y}_{c2}^{n+}(y_{p2})$ lies to the left of $\bar{y}_{c2}^{n}(y_{p2})$.

Having characterized both the period 2 decision of children who stayed at home in period 1 as well as the consequences of becoming independent in period 1 in terms of period 2 utility, we now move to period 1.

3.4 Period 1

A simplified presentation of the model's structure is given in Figure 4, which makes it apparent that in period 1 the residence choice is more involved than in period 2. The reason is the irreversibility of the moving-out decision and the possibility of regret once period 2 is reached. Naturally, the latter depends on the likelihood that period 2 incomes will fall to the left of the schedule $\bar{y}_{c2}(y_{p2})$. So far, we have been silent about the distribution of period 2 income for either family member, since period 2 decisions are taken after y_{c2} and y_{p2} are observed. We now assume:

$$(y_{c2}, y_{p2}) \sim F(y_{c2}, y_{p2}),$$

where $F(\cdot)$ is the joint cumulative distribution function (cdf) of period 2 incomes (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 $[\gamma_c, \infty) \times [\gamma_p, \infty)$.

Given the stochastic nature of income in period 2, to the extent that positive probability is assigned by $F(\cdot)$ to the regret region, staying home with the parents in period 1 has an *option value*, the value associated with waiting to see the realization of the period 2 incomes and deciding then whether or not to move. Of course, just like with any real option, this waiting value has to be traded off against the potential gains from moving out early on.

We define as Δ_1 the expected excess utility from moving out relative to staying home, from the point of view of period 1:

$$\begin{aligned} \Delta_1 \equiv\ & u(c_{i1}) + \int_{\gamma_p} \int_{\gamma_c} u(c_{i2})\, dF(y_{c2}, y_{p2}) \\ & - \left\{ u\left(c_{p1}^n\right) + \int_{\gamma_p} \left[\int_{\gamma_c}^{\bar{y}_{c2}(y_{p2})} u\left(c_{p2}^n\right) dF(y_{c2}) + \int_{\bar{y}_{c2}(y_{p2})}^{\infty} u(c_{i2})\, dF(y_{c2}) \right] \right\}. \end{aligned}$$

The function Δ_1 is defined over feasible income values (y_{c1}, y_{p1}), the set $[\gamma_c, \infty) \times [\gamma_p, \infty)$, which is identical to the domain of (y_{c2}, y_{p2}), as stated.

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 as that of an independent person: $c_{c2} = c_{i2}$. The last two terms represent the expected utility from staying home in period 1. In this case the child retains the possibility of choosing the best possible residential arrangement in period 2. This is reflected in the fact that, given y_{p2}, for values of y_{c2} not exceeding $\bar{y}_{c2}(y_{p2})$, the child chooses to remain with her parents and $c_{c2} = c_{p2}^n$. Alternatively, when $y_{c2} > \bar{y}_c(y_{p2})$, she moves out and $c_{c2} = c_{i2}$.

As in period 2, the child will decide to move out if the expected utility from doing so exceeds the expected utility from staying home (i.e., when $\Delta_1 > 0$):

$$\begin{aligned} & u(c_{i1}) + \int_{\gamma_p} \int_{\gamma_c} u(c_{i2})\, dF(y_{c2}, y_{p2}) > \\ & u(c_{pn1}) + \int_{\gamma_p} \left[\int_{\gamma_c}^{\bar{y}_{c2}(y_{p2})} u\left(c_{p2}^n\right) dF(y_{c2}) + \int_{\bar{y}_{c2}(y_{p2})}^{\infty} u(c_{i2})\, dF_c(y_{c2}) \right] dF_p(y_{p2}). \end{aligned} \tag{7}$$

When the child's period 2 income is sufficiently high $(y_{c2} > \bar{y}_{c2}(y_{p2}))$, having moved out in period 1 does not carry any utility loss: the residential status that prevails in period 2 is the one that the child would have chosen anyway. Therefore, for $y_{c2} >$

$\bar{y}_{c2}(y_{p2})$, the terms concerning period 2 utility while independent cancel on both sides of (7). Thus, the child will move out in period 1 if:

$$u(c_{i1}) - u\left(c_{p1}^{n}\right) > \int_{\gamma_p} \left[\int_{\gamma_c}^{\bar{y}_{c2}(y_{p2})} \left(u\left(c_{p2}^{n}\right) - u(c_{i2})\right) dF_c(y_{c2}) \right] dF_p(y_{p2}). \qquad (8)$$

It is worth commenting on equation (8). First of all, the right-hand side is nonnegative. It contains the difference between coresidence and independence. Since this is the difference between the best outcome and a forced condition, it represents the gain in expected utility associated with waiting for period 2 before choosing whether to move out (the option value). It will be strictly positive if the cdf $F(\cdot)$ places positive mass on the regret region, over which the integration on the right-hand side of (8) is being performed. The left-hand side represents the difference in period 1 utility from being independent relative to moving out. Not surprisingly, the child will choose to 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, to ascertain how it depends on period 1 income of the parent, y_{p1}, as well as on expectations of the future incomes of both parent and child, as summarized in $F_p(\cdot)$ and $F_c(\cdot)$. As before, $\bar{y}_{c1}$ is defined as the root of Δ_1, namely: $\Delta_1(\bar{y}_{c1}) = 0$.

We note the following facts:

Lemma 4 *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 period 2 function $\tilde{y}_{c2}(y_{p2})$.*

Proof. *See Appendix 1.*

This lemma says that, on the space 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 (8) coincides with function Δ_j –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. These properties now carry over to Δ_1.

We now define formally the regret region, R, namely period 2 income pairs such that the child's preferred living arrangement is coresidence:

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

Then,

Lemma 5 *If $F(R) = 0$, the schedules $\bar{y}_{c1}(y_{p1})$ and $\bar{y}_{c2}(y_{p2})$ coincide.*

Proof. *See Appendix 1.*

Intuitively, the lemma says that, if the probability of regret is zero, then today's residence choice becomes static and depends only on the comparison of the current utility levels attained under the two residential arrangements. But if this is the case, the residential decision of period 1 must be identical to that of period 2, when there are no future implications to consider and therefore no possibility of regret.

Naturally, the scenario described in Lemma 5 is only a starting point towards the more interesting case when there is a positive probability of regret, the case we now consider. For $F(R) > 0$, we can show the following:

Proposition 1 *The period 1 moving-out threshold schedule $\bar{y}_{c1}(y_{p1})$, 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 $F(R) > 0$*

Proof. *See Appendix 1.*

This result is illustrated in Figure 5. In order to clarify the previous result, consider the following example. Suppose that parental income takes the same value across periods:

$$y_{p1} = y_{p2} = y.$$

Suppose further that the distribution $F(\cdot)$, now defined simply over the child's period 2 income y_{c2}, assigns positive probability to the regret region. Then, proposition (1) shows that

$$\bar{y}_{c1}(y) > \bar{y}_{c2}(y).$$

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

We are further interested in the impact of y_{p1} on the period 1 moving-out threshold, as well as in the implications that differently shaped distribution functions $F(\cdot)$ have on the same value. First, define the expected regret as:

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

$\bar{R}$ naturally depends on the distribution $F(\cdot)$. We now rewrite the condition $\Delta_1 = 0$ while specifying 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,

$$\Delta_1\left(\bar{y}_{c1}, y_{p1}\right) = \bar{R}. \tag{9}$$

We are now ready to state the following result:

Lemma 6 *Higher period 1 parental income y_{p1} raises the child's moving-out threshold $\bar{y}_{c1}(y_{p1})$ for $\alpha < 1$.*

Proof. *See Appendix 1.*

Lastly, let us consider the impact of expectations on the child's moving-out decision. How do different distribution functions $F(\cdot)$ affect $\bar{y}_{c1}$? For simplicity, we currently assume that the distributions of period 2 income of parent and child are independent. This implies:

$$F(y_{c2}, y_{p2}) = F_c(y_{c2}) F_p(y_{p2}).$$

The natural way to compare different income distributions is to rank them in terms of stochastic dominance. Let $F_p^1(y_{p2})$ first-order stochastically dominate $F_p^2(y_{p2})$. Then, $F_p^1(\cdot)$ and $F_p^2(\cdot)$ must satisfy $F_p^1(y_{p2}) \leq F_p^2(y_{p2})$, for all y_{p2} values. It is well-known that the expectation over non-decreasing functions of y_{p2} computed under the distribution F_p^1 exceeds the same expectation under F_p^2.

Intuition suggests that, if the parent received the good news that the distribution of his period 2 income had changed from F_p^2 to F_p^1, the child of a wealthier parent (in expectation) would be more reluctant to move out in period 1. In other words, we would expect a higher value of $\bar{y}_{c1}$ to be associated with F_p^1 relative to the moving-out threshold corresponding to F_p^2. As it turns out, this is not always the case. The reason is the ambiguous impact of higher parental income over Δ_2, for (y_{c2}, y_{p2}) values in the regret region, as described in Lemma 2. There is, however, a subset of distributions F_p for which that result holds.

To illustrate the problem, fix a particular value for the child's period 2 income, y_{c2}, with $y_{c2} > (n/(n-1))\gamma_c$. That is, y_{c2} is to the right of the schedule $\bar{y}_{c2}(y_{p2})$, when $y_{p2} = \gamma_p$ (recall equation (6)). Along the vertical line corresponding to this value of y_{c2}, consider increasing y_{p2} starting at $y_{p2} = \gamma_p$. As can be seen from Figure 3, for low values of y_{p2}, the income pairs (y_{c2}, y_{p2}) initially belong to the moving-out region. Then, the schedule $\bar{y}_{c2}(y_{p2})$ is reached and we enter the regret region. There is a set of y_{p2} values ranging between the two schedules $\bar{y}_{c2}(y_{p2})$ and $\tilde{y}_{c2}(y_{p2})$. These income pairs are such that the child who moved out would like to move back home and her parents are not wealthy enough to provide her with financial help. For this range of y_{p2} values –which induce regret but are associated with zero transfers– higher parental income unambiguously reduces Δ_2. Therefore, 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 expected result that F_p^1 induces a higher period 1 moving-out threshold relative to F_p^2 goes through.[10]

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 formally define a convenient subset $\mathcal{F}$ of Ω_F:

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

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 (y_{c2}, y_{p2}) pairs are such that transfers would be granted to children living on their own

We may now state the following result.

[10] Note that, given the assumption of independence between F_p and F_c, having a rectangle of (y_{c2}, y_{p2}) values that falls under the $\tilde{y}_{c2}(y_{p2})$ schedule necessarily requires a lower bound on y_{c2} strictly greater than $(n/(n-1))\gamma_c$.

Proposition 2 *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 *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* $\alpha < 1$, $\bar{y}_{c1}\left(F_p^1\right) \geq \bar{y}_{c1}\left(F_p^2\right)$

Proof. *See Appendix 1.*

The reasoning underlying proposition 2 applies to a comparison of distributions for the child's income. We therefore offer the next result without proof.

Proposition 3 *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* $\alpha < 1$, $\bar{y}_{c1}(F_c^1) \leq \bar{y}_{c1}(F_c^2)$*.*

This result states that, if the child expects to have higher income in period 2, she will be more eager to move out in period 1 as there is a lower expected loss from regret. As argued before, the impact of changes in the distributions F_p and F_c is ambiguous for a general case. Propositions 2 and 3 are derived by restricting the set of distributions to an appropriate domain which eliminates the ambiguity. Naturally, changes in F_p and F_c may happen simultaneously. What is relevant for the child's period 1 moving-out decision is the net outcome of those changes.

A final remark concerns the effects of family size on the period 1 moving-out decision. Since consumption when coresiding is subject to income pooling, quite intuitively, a larger family size unambiguously leads the child to leave their parents at a lower income level:

Proposition 4 *The first-period moving-out threshold for a child with family size* $n^+ > n$*, denoted* $\bar{y}_{c1}^{n+}$*, strictly exceeds the corresponding threshold for a child with family size* n*,* $\bar{y}_{c1}^{n}$*.*

Proof. *See Appendix 1.*

Let us end this section by summarizing the key results provided by the model as a guide for our 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 –hers and her parents'– 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 parents' 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 consumption less than theirs), there will be regret even if parents are wealthy enough to provide financial transfers to independent children.

(ii) *Job insecurity.* In general, higher parental income and lower child income (current or future expected), deter children from leaving their parents house. Thus, we expect higher job insecurity of the parents to hasten emancipation and higher job insecurity of the children to delay emancipation. Nevertheless, there are cases where this result may not hold. The reason is that the level of consumption depends on the residential regime. For instance, a child living independently who finds she has a higher income than she expected will be happy she moved out (because she gets to consume all of this unexpected income rather than sharing it with her family). However, if consumption while independent is higher than at home, *at the margin* the increase in utility provided by the unexpected increase in consumption would be larger at home. In the ex-ante calculation of the moving-out decision, the higher *marginal utility effect* may dominate the *sharing effect.* This will depend, for instance, on the degree of risk aversion of the child. Typically, we expect the sharing effect to dominate, and we have provided examples where it unambiguously does. The model also predicts that a child with a higher number of siblings should leave her parents earlier, so that she does not have to share their income with them.

4 The data

In this section we describe our data set. We start with the sample design, which is crucial to our capacity to estimate the causal effects of interest. We then describe in turn the treatment and the outcome indicators we are interested in.

4.1 Sample design

To answer the questions addressed in this paper we need information on a representative sample of Italian individuals (hereafter the "children") of working age (i.e. older than 14) and living with their parents. Such information is offered by the Italian Survey of Household Income and Wealth, a rotating panel providing a representative sample of the Italian population. We use four waves: 1991-93-95-98, each of which contains data on 7,000 to 8,000 households composed of a total of around 20,000 to 25,000 individuals.[11]

Our goal is to explain the children's decision to leave home *after* a given year (1991, 1993 and 1995), as a function of household and individual characteristics measured in that year, among which the degree of job insecurity of children and their fathers are the *treatment* variables of major interest.[12] Indicators of these decisions of children (the *outcome* indicators) can only be obtained for the households belonging to the panel section of the SHIW. For instance, only 2,699 of the 8,135 households interviewed in 1995 were reinterviewed in 1998.[13] Since these panel households were randomly selected, they are still representative of the reference population.

[11] The respective numbers of households and individuals are as follows. 1991: 8,188/24,930; 1993: 8,089/24,013; 1995: 8,135/23,924; and 1998: 7,147/20,901.

[12] We focus on the fathers, 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).

In our empirical investigation we will be using two measures of job insecurity. *Experienced job insecurity* is a dummy taking the value of 1 if the individual is unemployed. This variable has the advantage that it can be used for all three years in the sample. The second and most interesting measure, *Perceived job insecurity*, which we will describe in detail below, is constructed from the answers to a survey question about the probability of having a job in the following year.

The problem, from our viewpoint, is that the intersection between the subsample of children belonging to the panel and the subsample of those for whom we have information on job insecurity is very small. The reason is that only individuals who are either working or unemployed are asked about their job prospects. This excludes the retired, "househusbands", and students. Retirement is by far the most important source of selection. In our analysis we prefer not to consider retired fathers as having a sort of perfectly secure job, since they are in large part individuals who enjoy perfectly safe incomes not through choice but due to their age and to legal provisions, and since being completely sure about having no unemployment in the subsequent year is not equivalent of being sure of no unemployment for life, as the retired are. Since we are interested in emancipation, we also restrict the sample to children aged up to 40 years old. These criteria, while required by the focus of our analysis, reduce the sample considerably, to 2,821 children, 923 of them in the 1995 sample. The characteristics of households in our working sample for 1995 are described in Table 3. The descriptive statistics for 1991 and 1993, not reported to save space, are very similar. Since the SHIW is a rotating panel, individuals may be observed repeatedly over time: out of the 2,821 observations, 1,687 correspond to different individuals.

A final criterion applied by the SHIW for asking household members about their job prospects is purely random: that the head be born in an odd year (those born in even years are asked an alternative set of questions). Out of the 923 observations in our 1995 sample, only 395 parents and 315 children were asked. Since this would leave us with a too small sample, we will resort below to methods that allow us impute perceived insecurity to all individuals.

As a result of the selections described above, imposed by the available data, we are forced to use a relatively small sample. But this sample is still representative of the population of interest and its timing structure is suitable for exploring the relationship between job insecurity and the subsequent decisions of children to leave home, controlling for a large set of individual and family background characteristics. Before presenting our results a description of the treatment and outcome indicators is needed.

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:[14]

[14]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.

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 4,799 individuals, which become 4,205 after non-respondents (8%) and 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.[15] The authors also compare this source, restricting the sample to those employed, with the Survey of Economic Expectations (SEE), which contains a similar question for employed workers in the US.[16] 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.

Our own cross-sectional sample is not very different from the above for the sample of fathers, as described in Table 4. The average perceived unemployment probability of fathers is 20% vis-à-vis 22% in the full sample of Guiso *et al.* (2002). 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. On the other hand, in our case the probability mass at zero expected unemployment is slightly lower, 53%. This fraction is large enough to induce us to use in the regressions, as the measure of perceived job insecurity a dummy taking the value 1 if the individual perceived a positive probability of unemployment (for both parents and children).[17]

As expected, the perceptions of children reveal much more uncertainty. Only the first quartile perceives no chances of unemployment, the median is 30%, and a 50% chance of unemployment is reached in the 6th decile. This is also coherent with the unemployment experience presented in Table 3: unemployment affects 8% of fathers but a staggering 57% of children.

The last row of Table 3 contains information on the outcome variable of interest. It is a dummy variable taking the value 1 if the child left the household between 1995 and

[15] 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).

[16] *"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).

[17] An alternative measure, in which the dummy was equal to 1 if the probability was above 0.5 yielded very similar results to those presented below.

1998.[18] Of the 923 children living with their parents in 1995 on which our analysis is based, only 48 (5%) decided to leave home in the following three years (26 females and 22 males). Like in the aggregate, as mentioned in the Introduction, emancipation rates have trended downwards: in our sample the rates for 1991 and 1993 (which correspond to having left the household by 1993 and 1995, respectively), were 17.6% and 10%.

4.3 Imputations of perceived job insecurity

Using the samples just described, we estimate probit regressions for the decision to live independently, controlling for a set of characteristics explained in the next section. In the models using the perceived probability of unemployment, the small sample size forces us to use imputations for this variable. 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. 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 impute the missing data on the basis of the available information.

We follow two strategies to impute perceived job insecurity when it is missing. The first strategy, sometimes called in the literature Conditional Mean Imputation or CMI (Little and Rubin, 1987), is based on a regression of perceived job insecurity on a set of observable characteristics for the individuals with full information. Missing values for the remaining individuals are then replaced by predicted values obtained using the estimated coefficients of the regression and the observable characteristics of these individuals. To improve the reliability of the imputation we use four different regressions for, respectively, working fathers, non-working fathers, working children, and non-working children. In this way for each of these groups the prediction is based on the widest possible set of regressors.[19]

The final rows of Table 3 show the statistics for observed and imputed values of perceived job insecurity, as they are used in the probit regressions, namely taking the value 1 if the individual perceives non-zero chances of unemployment. CMI produces a larger average degree of insecurity, specially for fathers. A full comparison across deciles is shown in Table 4.

Conditional mean imputation is simple to implement, but unfortunately it tends to generate a downward bias in the estimates of the effect of the imputed variable in the main equation of interest, as well as an upward bias in its standard error. This happens because the imputed values are estimates of the true missing values, and of course these estimates are subject to an error. Given the opposite direction of these biases, it is

[18] Note that children who are in jail, long-term hospitalized or dead in 1998 are excluded.

[19] Specifically, the list of regressors used in the four cases is as follows. Father not working: age, years of schooling, dummies for geographical area, dummy for homeownership, yearly income, and family wealth. Father working: as for father not working plus dummies for job qualification, sector and firm size. Children: as for the corresponding labor status for fathers, plus a dummy for single. Results of the four regressions are available from the authors upon request.

not simple to assess their consequences in terms of significance of the estimates in the equation of interest. Leaving to a future version of this paper an attempt to correct our results for these biases, here we also present estimates based on an alternative imputation strategy, known as Multiple Imputation.[20]

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

$$\bar{\beta}_M = \sum_i^M \frac{\hat{\beta}_i}{M}. \tag{10}$$

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

$$\bar{W}_M = \sum_i^M \frac{\hat{W}_i}{M} \tag{11}$$

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

$$\bar{B}_M = \sum_i^M \frac{(\hat{\beta}_i - \bar{\beta})^2}{M-1}. \tag{12}$$

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

$$T_M = \bar{W}_M + \frac{M+1}{M}\bar{B}_M \tag{13}$$

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

$$\nu = (M-1)\left[1 + \frac{1}{M+1}\frac{\bar{W}_M}{\bar{B}_M}\right] \tag{14}$$

Our preliminary results obtained with this strategy, reported in Section 5, are based on $M = 10$ multiple imputations.[21]

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

[21] As shown in Rubin (1987), p. 114, this small number of imputations is sufficient, under normal

5 Microeconomic evidence on emancipation decisions of Italian children

As already stated, we estimate probit regressions for the decision to live independently. Ideally, we wish to compare children who share all relevant personal and family characteristics potentially affecting the outcome, so that we can isolate the effect of the two treatments, job insecurity of the child and of the father.

The estimations share a set of control variables dated in the year where all children are observed coresiding. For children, we control for their age, gender (female = 1), and completed years of schooling. 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. We also control for family wealth, the income of members other than the child and the father, homeownership (owner-occupied = 1), and region of residence (through two dummy variables for Center and South, with North becoming the reference).

One of the implications of our theoretical model is that children will tend to leave earlier the higher the number of their siblings. In order to test this implication we add as a control the number of children in the household. Also, as mentioned in the Introduction, Díaz and Guilló's (2002) model predicts that children will emancipate later if they enjoy household services provided by their mother. We introduce two dummies, one for whether the mother is present and a second one where this dummy is interacted with a dummy for whether the mother works.

Descriptive statistics of all the control variables are presented in Table 3. Let us now present in turn the empirical results for experienced and perceived job insecurity.

5.1 Experienced job insecurity

Table 5 contains the estimates for the model where we use the state of unemployment as a measure of insecurity. We present separate regressions for years 1991, 1993, and 1995. The coefficients on all the dummy variables measure the change in the probability due to a discrete change from 0 to 1.

As predicted by the model in Section 3, the probability that an unemployed child will leave her parents' home is lower than that for an employed child, an effect estimated to be between 4 and 6 percentage points. While the coefficients are similar across periods, the effect only attains statistical significance in 1995. This is even more so for the father being unemployed, which has a strictly positive effect on the child's emancipation only for the 1995 sample, that is similar in size to that of the child's own unemployment.

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circumstances, to achieve efficiency. We will, however, try with a larger number of imputations in future versions of this paper to increase efficiency.

Free-ware software written by Joe Schafer (Department of Statistics, The 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.

As expected, emancipation is more likely the older the child and also more likely for females than for males. The remaining control variables do not attain statistical significance, including the number of children and the presence and labor status of the mother.

5.2 Perceived job insecurity

Table 6 shows the estimates obtained with the indicator on perceived job insecurity, with values for individuals who were not asked imputed using the conditional mean method. This estimation can only be implemented with the 1995 sample. There is an evident change with respect to experienced insecurity. The effect of the child's own insecurity doubles, to around -11%, while the father's halves to 2% and loses significance.

The results obtained with the multiple imputation strategy are reported in Table 7, which includes only the coefficients of the two measures of job instability of interest for the ten different imputations obtained from this method. Other variables exhibit coefficients which are similar to those in the preceding two tables. Both job insecurity variables have the respective signs predicted by our model in all cases (except in one case, for the father's insecurity), and they are statistically significant in five out of the ten cases for child insecurity and three out of ten cases for parental insecurity. The average of these marginal effects is equal to -3.9 percentage points when the child goes from perceiving none to some insecurity and to 2.3 points in the case of the father. The p-values for the one-sided t-ratios for a negative and a positive effect, respectively, are 0.08 and 0.20. Thus, although the pattern of signs clearly confirms our story, aggregate significance is low. More work with multiple imputation will be needed to see whether the efficiency of these estimates can be improved.

It is worth noting that multiple imputation brings the estimated effect of child job insecurity much closer to the one found for experienced insecurity based on unemployment status, while the estimated effect of parental insecurity is in line with the value estimated from conditional mean imputation.

6 Conclusions

In this paper we explore one potential cause for a most striking feature of the Italian society today, namely that grown-up children are very reluctant to leave home. We have tested to what extent this fact depends on the high job insecurity suffered by young Italians and the high job security enjoyed by their fathers.

We have shown that aggregate data for the 15 European Union member countries around 1995, on emancipation rates for the 15-24 and 24-29 years old brackets and on perceived job insecurity, are consistent with this hypothesis. Thus, we go on and test it with microeconomic panel data from the Italian Survey of Household Income and Wealth (SHIW) collected by the Bank of Italy.

Taken together, our preliminary empirical results indicate that the likelihood that young Italians aged 15 to 37 left the parental home in the 1990's was negatively affected by the fact that they were unemployed and positively affected by their father's unemployment. The magnitude of the effects was, in both cases, around 5 percentage points.

The same pattern obtains, for the move out of the parents household from 1995 to 1998, for job insecurity measured in terms of the individuals' own perceptions of their chances of having a job over the subsequent 12 months. These results are found estimating simple probit models controlling for children's, father's and family characteristics, and imputing a perception for a fraction of the individuals in the sample. The simplest method, conditional mean imputation, leads to a doubling, to -11 percentage points, of the effect of the child's insecurity, and a halving of the father's, to 2 points. A more sophisticated method, Rubin's multiple imputation confirms the pattern of signs found, although it lowers the coefficients' significance. In this case the estimates are around -4 and 2 percentage points for children and fathers, respectively.

At this stage we are working on strengthening the results by enlarging the sample and improving on the implementation of the imputation methods.

Appendix 1. Proofs of results in Section 3

Proof of Lemma 1

We first note that:

$$\Delta_2 > 0 \Leftrightarrow c_{i2} > c^n_{p2}.$$

In words, the child prefers to move out if and only if her consumption while independent exceeds the consumption she would enjoy if she did not move out. When parental income is such that $\tilde{y}_{c2} > \gamma_c$ (that is, the parent would provide positive transfers to a child who moved out while earning income $y_{c2} < \tilde{y}_{c2}$), the difference in the child's consumption across residence states is:

$$c_{i2} - c^n_{p2} = \frac{y_{p2} + y_{c2} - \gamma_p - \gamma_c}{(\Gamma(n-1)+1)\,p} - \frac{y_{p2} + y_{c2} - \gamma_p}{np},$$

which is negative, since $\Gamma(n-1)+1 > n$. This shows that $\Delta_2(y_{c2}) < 0$ for $y_{c2} \in [\gamma_c, \tilde{y}_{c2}]$. For $y_{c2} \in (\tilde{y}_{c2}, \bar{y}_{c2}]$, the child would not receive any transfers if she moved out. In this case,

$$c_{i2} - c^n_{p2} = \frac{y_{c2} - \gamma_c}{p} - \frac{y_{p2} + y_{c2} - \gamma_p}{np}.$$

It is straightforward to show that the previous inequality is negative for income values y_{c2} such that $y_{c2} < \bar{y}_{c2}$, and positive for $y_{c2} > \bar{y}_{c2}$. This proves that $\Delta_2(y_{c2}) < 0$ for $y_{c2} \in (\tilde{y}_{c2}, \bar{y}_{c2})$ and $\Delta_2(y_{c2}) > 0$ for $y_{c2} > \bar{y}_{c2}$.

The derivative of Δ_2 with respect to y_{c2} is:

$$\frac{\partial u(c_{i2})}{\partial c_{i2}} \frac{\partial c_{i2}}{\partial y_{c2}} - \frac{\partial u\left(c^n_{p2}\right)}{\partial c^n_{p2}} \frac{\partial c^n_{p2}}{\partial y_{c2}}.$$

For $y_{c2} \in (\tilde{y}_{c2}, \bar{y}_{c2})$, $\partial c_{i2}/\partial y_{c2} = 1/p$, and $\partial c^n_{p2}/\partial y_{c2} = 1/np$. Also, since $\Delta_2 < 0$ in this range, $\partial u(c_{i2})/\partial c_{i2}$ exceeds $\partial u\left(c^n_{p2}\right)/\partial c^n_{p2}$. This implies that $\partial \Delta_2/\partial y_{c2} > 0$ in this interval.

The expression for $\partial \Delta_2/\partial y_{c2}$, for $y_{c2} \in (\gamma_c, \tilde{y}_{c2})$ is given by:

$$\frac{\partial \Delta_2}{\partial y_{c2}} = (c_{i2})^{-\alpha} \frac{1}{(\Gamma(n-1)+1)\,p} - \left(c^n_{p2}\right)^{-\alpha} \frac{1}{np},$$

and

$$\begin{aligned} \frac{\partial \Delta_2}{\partial y_{c2}} &< 0 \Leftrightarrow \left(\frac{c^n_{p2}}{c_{i2}}\right)^{\alpha} < \frac{\Gamma(n-1)+1}{n} \\ &\Leftrightarrow \left(\underbrace{\frac{y_{p2} + y_{c2} - \gamma_p}{y_{p2} + y_{c2} - \gamma_p - \gamma_c}}_{>1}\right)^{\alpha} > \left(\underbrace{\frac{(\Gamma(n-1)+1)}{n}}_{>1}\right)^{1-\alpha}. \end{aligned}$$

Since $\Gamma > 1$, the last term in brackets on the right exceeds unity. When $\alpha > 1$, this term will be smaller than 1 and, since the left-hand side of the inequality is greater than 1, the inequality will be satisfied for all values of y_{p2} and y_{c2}. This proves that $\Delta(y_{c2})$ is strictly decreasing for $y_{c2} \in (\gamma_c, \tilde{y}_{c2})$ when $\alpha > 1$.

The expression for $\partial\Delta_2/\partial y_{c2}$, for $y_{c2} > \bar{y}_c$ is given by:

$$\frac{\partial \Delta_2}{\partial y_{c2}} = (c_{i2})^{-\alpha}\frac{1}{p} - \left(c_{p2}^n\right)^{-\alpha}\frac{1}{np},$$

and

$$\frac{\partial \Delta_2}{\partial y_{c2}} > 0 \Leftrightarrow \left(\frac{c_{p2}^n}{c_{i2}}\right)^{\alpha} > \frac{1}{n} \Leftrightarrow \left(\underbrace{\frac{y_{p2}+y_{c2}-\gamma_p}{y_{c2}-\gamma_c}}_{>1}\right)^{\alpha} > \left(\underbrace{\frac{1}{n}}_{<1}\right)^{1-\alpha}.$$

Since $y_{p2} \geq \gamma_p$, the fraction on the left-hand side of the last inequality exceeds unity and, since families have at least two persons, the fraction on the right-hand side is smaller than 1. For $\alpha < 1$, therefore, the inequality above is satisfied for all values of y_{c2} and y_{p2}. This proves that $\Delta_2(y_{c2})$ is strictly increasing for $y_{c2} > \bar{y}_{c2}$ when $\alpha < 1$. ■

Proof of Lemma 3

The relationship between $\tilde{y}_{c2}^{n+}$ and $\tilde{y}_{c2}^{n}$ follows directly from (5). Likewise, equation (6) establishes the result $\bar{y}_{c2}^{n+}(y_{p2}) < \bar{y}_{c2}^{n}(y_{p2})$. The inequality $\Delta_2^{n+}(y_{c2}) > \Delta_2^{n}(y_{c2})$ can be shown as follows. For $y_{c2} \leq \tilde{y}_{c2}^{n+}$, the child is receiving transfers under either family size. Since $c_{i2}(t_2 > 0)$ is increasing in n whereas c_{p2}^n is decreasing in n, it follows that $\Delta_2^{n+}(y_{c2}) > \Delta_2^{n}(y_{c2})$. For the range $\left(\tilde{y}_{c2}^{n+}, \tilde{y}_{c2}^{n}\right)$, $\Delta_2^{n+}(y_{c2})$ is increasing whereas $\Delta_2^{n}(y_{c2})$ is decreasing. Since Δ_2 is continuous under a given family size, this further establishes $\Delta_2^{n+}(y_{c2}) > \Delta_2^{n}(y_{c2})$ for the latter range of incomes. Finally, for $y_{c2} > \tilde{y}_{c2}^{n}$, both functions Δ_2 are increasing. Since Δ_2 is the difference in utility in the state of independence versus coresidence, the utility while independent is the same under either family size in this range, as no transfers are given. However, $c_{p2}^{n+} < c_{p2}^{n}$, implying

$$\Delta_2^{n+}(y_{c2}) = u(c_{i2}) - u\left(c_{p2}^{n+}\right) > u(c_{i2}) - u\left(c_{p2}^{n}\right) = \Delta_2^{n}(y_{c2}).$$ ■

Proof of Lemma 4

Period 1 transfers, when positive, are chosen by parents to solve:

$$\max_{\tilde{t}_1 \geq 0}\left\{\lambda(n-1)\,u\left(\frac{y_{p1}-\tilde{t}_1-\gamma_p}{p(n-1)}\right) + (1-\lambda)\,u\left(\frac{y_{c1}+\tilde{t}_1-\gamma_c}{p}\right)\right\}.$$

This is exactly the same problem stated for the choice of period 2 transfers, given in (3), except that now the income values of parent and child bear period 1 subscripts. Consequently, the schedule $\tilde{y}_{c1}(y_{p1})$ coincides with the schedule $\tilde{y}_{c2}(y_{p2})$. ■

Proof of Lemma 5

We rewrite the definition of Δ_1 as follows:

$$\begin{aligned}\Delta_1 &= u(c_{i1}) - u\left(c_{p1}^n\right) + \int_{\gamma_p}\left[\int_{\gamma_c}^{\bar{y}_{c2}(y_{p2})}\left(u\left(c_{p2}^n\right) - u(c_{i2})\right) dF_c(y_{c2})\right] dF_p(y_{p2}) \\ &= u(c_{i1}) - u\left(c_{p1}^n\right) + \int_R \left(u\left(c_{p2}^n\right) - u(c_{i2})\right) dF(y_{c2}, y_{p2}) \\ &= u(c_{i1}) - u\left(c_{p1}^n\right) = \Delta_2.\end{aligned}$$

Let the schedule $\bar{y}_{c2}(y_{p2})$ be denoted by $\bar{y}_c(y_p)$, interpreted now as a function relating incomes of parent and child for any period. Given that Δ_1 coincides with Δ_2, the relationship between child and parent income, $\bar{y}_{c2}(y_{p2})$, that solves

$$\Delta(\bar{y}_{c2}(y_{p2})) = \Delta(\bar{y}_c(y_{p2})) = 0,$$

must also solve:

$$\Delta_1(\bar{y}_c(y_{p1})) = \Delta_1(\bar{y}_{c1}(y_{p1})) = 0. \blacksquare$$

Proof of Proposition 1

If $F(R) > 0$, the right-hand side of (8) is strictly positive. The moving-out income threshold $\bar{y}_{c1}$, such that $\Delta_1(\bar{y}_{c1}) = 0$, solves:

$$\begin{aligned}& u(c_{i1}(\bar{y}_{c1})) - u\left(c_{p1}^n(\bar{y}_{c1})\right) = \int_R \left(u\left(c_{p2}^n\right) - u(c_{i2})\right) dF(y_{c2}, y_{p2}) > 0 \\ \Leftrightarrow \quad & \Delta(\bar{y}_{c1}) = \int_R \left(u\left(c_{p2}^n\right) - u(c_{i2})\right) dF(y_{c2}, y_{p2}) > 0.\end{aligned}$$

Applying Lemma 1 to Δ_1, we know that it is strictly negative for $y_{c1} < \bar{y}_{c2}$, and strictly positive for $y_{c1} > \bar{y}_{c2}$. Further, from the properties of the utility function $u(\cdot)$, Δ_1 is continuous. Since $\Delta_2(\bar{y}_{c2}) = 0$, it follows that, for identical values of parental income across periods, $y_{p1} = y_{p2}$, the value of $\bar{y}_{c1}(y_{p1})$ that solves the previous equation must strictly exceed $\bar{y}_{c2}(y_{p1})$.$\blacksquare$

Proof of Lemma 7

Since $\bar{R} \geq 0$, we know that $\bar{y}_{c1} \geq \bar{y}_{c2}$. From Lemma 1, we know that $\partial\Delta_1/\partial y_{c1} > 0$ for $y_{c1} > \bar{y}_{c2}$, when $\alpha < 1$. Simple algebra shows that, for $y_{c1} = \bar{y}_{c2}$,

$$\frac{\partial \Delta_1}{\partial y_{c1}}|_{y_{c1}=\bar{y}_{c2}} = \frac{1}{p} - \frac{1}{np} > 0.$$

Further, $\bar{y}_{c1} \geq \bar{y}_{c2}$ implies that $y_{p1} < \tilde{y}_{p2}$. From Lemma 2, we know that Δ_1 is strictly decreasing in y_{p1} for $y_{p1} \in \left[\gamma_p, \tilde{y}_{p2}\right)$. Therefore, fully differentiating (9), we get:

$$\frac{\partial \Delta_1}{\partial \bar{y}_{c1}} d\bar{y}_{c1} + \frac{\partial \Delta_1}{\partial y_{p1}} dy_{p1} = 0 \Leftrightarrow \frac{d\bar{y}_{c1}}{dy_{p1}} = -\frac{\partial\Delta_1/\partial y_{p1}}{\partial\Delta_1/\partial \bar{y}_{c1}} > 0,$$

and the result follows. ■

Proof of Proposition 2

$\bar{y}_{c1}\left(F_p^1\right)$ is the solution to the first line of the following equation:

$$\Delta\left(\bar{y}_{c1}\left(F_p^1\right), y_{p1}\right) = \int_{\gamma_p}\left[\int_{\gamma_c}^{\bar{y}_{c2}(y_{p2})}\left(u\left(c_{pn2}\right) - u\left(c_{i2}\right)\right) dF_c\left(y_{c2}\right)\right] dF_p\left(y_{p2}\right)$$

$$\begin{aligned} &= \int_{\gamma_c}\int_{\gamma_p}\left[-\Delta\left(y_{c2}, y_{p2}\right)\right] dF_{p2}^1 dF_{c2} \geq \int_{\gamma_c}\int_{\gamma_p}\left[-\Delta\left(y_{c2}, y_{p2}\right)\right] dF_{p2}^2 dF_c \\ &= \Delta\left(\bar{y}_{c1}\left(F_p^2\right), y_{p1}\right) \end{aligned}$$

where the inequality follows from the first-order stochastic dominance of F_p^1 over F_p^2 and the fact that Δ is a decreasing function of y_{c2} in the range $\left[\gamma_p, \tilde{y}_{p2}\left(y_{c2}\right)\right)$, as shown in Lemma 2. Since $\alpha < 1$, from Lemma 1 it follows that

$$\Delta\left(\bar{y}_{c1}\left(F_p^1\right), y_{p1}\right) \geq \Delta\left(\bar{y}_{c1}\left(F_p^2\right), y_{p1}\right) \Leftrightarrow \bar{y}_{c1}\left(F_p^1\right) \geq \bar{y}_{c1}\left(F_p^2\right). \blacksquare$$

Proof of Proposition 4

As seen in Lemma 3, the function $\Delta_2^{n+}\left(y_{c2}\right)$ is everywhere above the corresponding schedule $\bar{y}_{c2}^n\left(y_{p2}\right)$ for a smaller family. The moving-out threshold $\bar{y}_1^{n+}$ is now the solution to:

$$\Delta^{n+}\left(\bar{y}_{c1}^{n+}, y_{p1}\right) = \bar{R}^{n+}.$$

A larger family size unambiguously implies that $\bar{R}^{n+} < \bar{R}$ since $\Delta_2^{n+}\left(y_{c2}\right) > \Delta_2^n\left(y_{c2}\right)$ also holds over the regret region. Therefore, assuming $\alpha < 1$, this increases $\bar{y}_1$. Also, Δ_1^{n+} is everywhere above Δ_1^n, further requiring a lower root $\bar{y}_1^{n+}$ to $\Delta^{n+}\left(\bar{y}_{c1}^{n+}, y_{p1}\right)$ for a fixed value of $\bar{R}$ on the right-hand side. Therefore, a larger family size unambiguously reduces the period 1 moving-out threshold.

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Table 1: Youth coresidence rates in EU countries (1995)

	15 to 24 years old	25 to 29 years old
Belgium	75	18
Denmark	57	8
Germany	73	21
Greece	82	49
Spain	90	60
France	71	18
Ireland	74	32
Italy	90	56
Luxembourg	76	29
Netherlands	70	15
Austria	75	29
Portugal	86	50
Finland	50	6
Sweden	60	5
United Kingdom	67	18
European Union	76	30

Definition: Percentage of the population in each age group living at parental home in 1995.
Source: European Communities (1999), Tables B-15/16/20/21.

Table 2: Coresidence rate, job insecurity, and unemployment (1995-96) (Simple correlation coefficients)

	Coresidence rate	
	15 to 24 y.o.	25 to 29 y.o.
Job insecurity		
16 to 24 years old	0,60	0,63
25 to 44 years old	0,39	0,31
45+ years old	-0,25	-0,31
16-24 relative to 45+ y.o.	0,66	0,74
Unemployment rate		
15 to 24 years old	0,39	0,53
25 to 54 years old	0,09	0,26
55 to 64 years old	-0,51	-0,29
15-24 relative to 25-54 y.o.	0,49	0,52
15-24 relative to 55-64 y.o.	0,32	0,27

Definitions and sources:
(1) Coresidence rate: Percentage of the population in each age group living at parental home in 1995, from European Communities (1999). See Table 1.
(2) Job insecurity: Percentage of employees in each age group not strongly agreeing that "my job is secure" in 1996, from OECD (1997), Table 5.2, itself drawn from Eurobarometer 44.3, Eurostat (1996).
(3) Unemployment: Age-specific standardized unemployment rates in 1995 from OECD (1997).
All coefficients are computed for the 15 member countries of the EU shown in Table 1.

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

Variable	Mean	Std. Dev.	Min.	Max.
Age	23.32	4.14	15	37
Female	0.42	0.49	0	1
Years of schooling	11.09	3.33	3	18
North	0.33	0.47	0	1
Center	0.17	0.38	0	1
South	0.50	0.50	0	1
Father's age	52.58	5.80	33	72
Father's years of schooling	8.19	3.79	3	21
Mother present	0.98	0.15	0	1
Mother works	0.28	0.45	0	1
Number of children	2.41	1.11	1	7
Homeownership	0.67	0.47	0	1
Family wealth	0.31	0.49	-0.07	6.47
Income of other members	0.02	0.02	-0.01	0.12
Child unemployed	0.57	0.50	0	1
Father unemployed	0.08	0.28	0	1
Child's perceived job insecurity	0.73	0.44	0	1
Child's imputed perceived job insecurity	0.87	0.33	0	1
Father's perceived job insecurity	0.47	0.50	0	1
Father's imputed perceived job insecurity	0.65	0.48	0	1
Out in 1998	0.05	0.22	0	1

Descriptive statistics of variables measured in 1995 for the 923 children who:

- live with their father in 1995,
- belong to households interviewed in both 1995 and 1998,
- are aged between 15 and 40 years old in 1995,
- are not studying,
- have a father who is either employed or unemployed, and
- are still alive, not in jail and non long-term hospitalized in 1998.

This is the sample used in the econometric analysis. Imputed job insecurity was estimated using the conditional mean method. Monetary variables are in millions of Italian Liras.

Table 4: The indicator of perceived job insecurity: observed and imputed with the conditional mean method. 1995 sample

	Observed			Observed and imputed		
	Frequency	Percent	Cumul.	Frequency	Percent	Cumul.
Children						
0.0	84	26.67	26.67	116	12.57	12.57
0.1	32	10.16	36.83	85	9.21	21.78
0.2	32	10.16	46.98	91	9.86	31.64
0.3	10	3.17	50.16	63	6.83	38.46
0.4	15	4.76	54.92	88	9.53	48.00
0.5	48	15.24	70.16	99	10.73	58.72
0.6	19	6.03	76.19	171	18.53	77.25
0.7	23	7.30	83.49	156	16.9	94.15
0.8	18	5.71	89.21	20	2.17	96.32
0.9	21	6.67	95.87	21	2.28	98.59
1.0	13	4.13	100.00	13	1.41	100.00
Total	315	100.00		923	100	
Fathers						
0.0	209	52.91	52.91	326	35.32	35.32
0.1	49	12.41	65.32	176	19.07	54.39
0.2	35	8.86	74.18	193	20.91	75.30
0.3	12	3.04	77.22	105	11.38	86.67
0.4	15	3.80	81.01	46	4.98	91.66
0.5	26	6.58	87.59	40	4.33	95.99
0.6	2	0.51	88.1	8	0.87	96.86
0.7	7	1.77	89.87	6	0.65	97.51
0.8	14	3.54	93.42	7	0.76	98.27
0.9	7	1.77	95.19	3	0.33	98.59
1.0	19	4.81	100.00	13	1.41	100.00
Total	395	100.00		923	100.00	

Distribution of the indicator of job insecurity of children and parents in the restricted 1995 panel 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 conditional mean method.

Table 5: Experienced job insecurity and the decision of children to leave their parents' home. 1991, 1993 and 1995

	1991	1993	1995
Child's experienced job insecurity	-0.063	-0.039	-0.052**
	(0.036)	(0.022)	(0.018)
Father's experienced job insecurity	-0.075	0.003	0.057*
	(0.085)	(0.045)	(0.036)
Female	0.028	0.055**	0.029*
	(0.027)	(0.018)	(0.014)
Age	0.019**	0.017**	0.004*
	(0.004)	(0.003)	(0.002)
Years of schooling	0.003	-0.004	-0.001
	(0.005)	(0.003)	(0.002)
Father's age	-0.001	-0.004*	-0.001
	(0.003)	(0.002)	(0.001)
Father's years of schooling	-0.002	0.003	-0.002
	(0.004)	(0.003)	(0.002)
North	-0.028	-0.040	0.001
	(0.037)	(0.020)	(0.019)
South	0.030	0.008	0.012
	(0.033)	(0.022)	(0.016)
Homeownership	-0.011	-0.001	0.028*
	(0.030)	(0.020)	(0.012)
Family wealth	0.074	-0.034	0.008
	(0.045)	(0.032)	(0.011)
Income of other members	-0.266	-0.401	-0.189
	(1.099)	(0.673)	(0.397)
Number of children	-0.013	-0.006	0.005
	(0.014)	(0.009)	(0.006)
Mother present	0.056	0.012	0.012
	(0.071)	(0.050)	(0.031)
Mother works	-0.034	0.020	-0.013
	(0.031)	(0.024)	(0.013)
R^2	0.064	0.082	0.094
Number of observations:	841	1057	923

Marginal effects from probit estimations. The dependent variable is equal to 1 if the child is not living with parents in the subsequent survey year (1993, 1995, and 1998, respectively). All children live with their parents in the year shown in each column. Standard errors are reported in parentheses. The statistical significance of the test that the underlying coefficient is zero is denoted by: $p < 0.05 = *, p < 0.01 = **$.

Table 6: Perceived job insecurity with mean conditional imputation and the decision of children to leave their parents' home. 1995

Child's perceived job insecurity	-0.111*
	(0.041)
Father's perceived job insecurity	0.020
	(0.014)
Female	0.025
	(0.013)
Age	0.003
	(0.002)
Years of schooling	-0.001
	(0.002)
Father's age	-0.001
	(0.001)
Father's years of schooling	-0.002
	(0.002)
Center	0.002
	(0.019)
South	0.011
	(0.016)
Homeownership	0.029
	(0.013)
Family wealth	0.011
	(0.011)
Income of other members	0.015
	(0.40)
Number of children	0.005
	(0.006)
Mother present	0.005
	(0.038)
Mother works	-0.013
	(0.014)
R^2	0.095
N. obs:	923

N. of obs.: 923. Marginal effects from probit estimations. The dependent variable is equal to 1 if the child is not living with his/her parents in 1998). All children live with their parents in 1995. Perceived job insecurity is imputed with the mean conditional method. Standard errors are reported in parentheses. The statistical significance of the test that the underlying coefficient is zero is denoted by: $p < 0.05 = *$, $p < 0.01 = **$.

Table 7: Perceived job insecurity with multiple imputation and the decision of children to leave their parents' home. 1995

Child's perceived job insecurity	Father's perceived job insecurity	R^2
-0.012	0.018	0.061
(0.020)	(0.014)	
-0.025	0.021	0.065
(0.021)	(0.013)	
-0.037	0.017	0.069
(0.023)	(0.013)	
-0.059**	0.035*	0.095
(0.024)	(0.013)	
-0.022	0.025	0.066
(0.021)	(0.014)	
-0.050*	0.046**	0.102
(0.026)	(0.013)	
-0.017	0.021	0.063
(0.020)	(0.013)	
-0.073**	0.030*	0.096
(0.028)	(0.013)	
-0.042*	0.016	0.069
(0.024)	(0.014)	
-0.052*	-0.001	0.071
(0.026)	(0.014)	

N. of obs.: 923. Marginal effects from probit estimations. The dependent variable is equal to 1 if the child is not living with parents in 1998. All children live with their parents in 1995. Perceived job insecurity is imputed with the multiple imputation method. Only the coefficients on child's and the father's perceived job insecurity are reported. Other variables in the regresssion are as in Table 6. Standard errors are reported in parentheses. The statistical significance of the test that the is zero is denoted by: $p < 0.05 = *$, $p < 0.01 = **$.

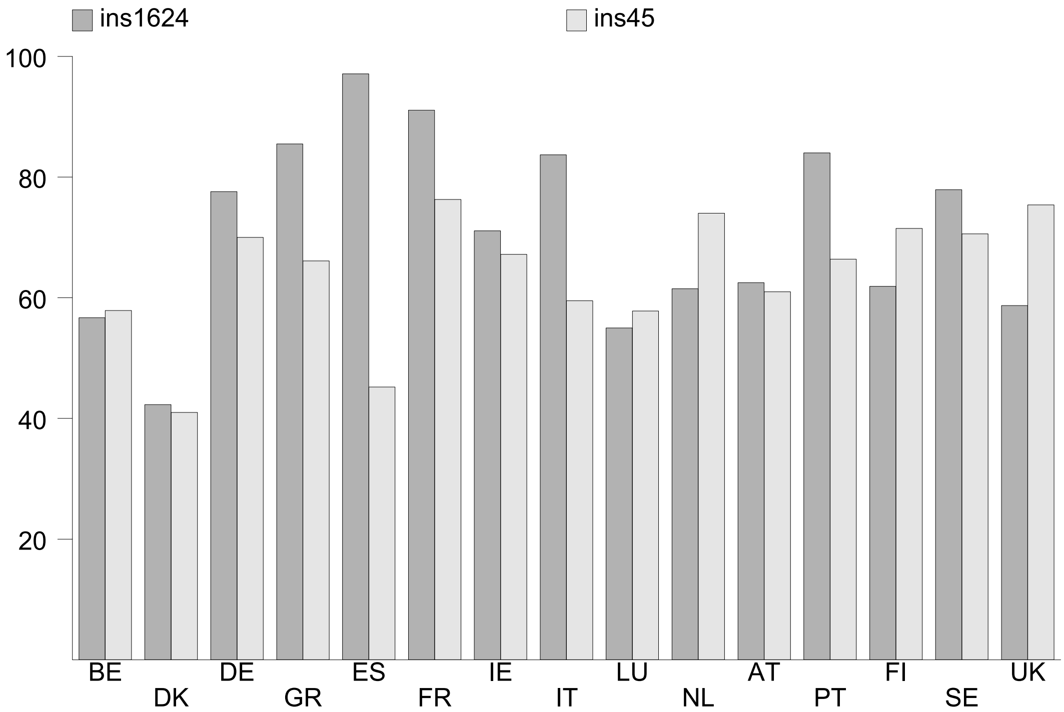

Figure 1. Perceived job insecurity in EU countries. Young and Over 45

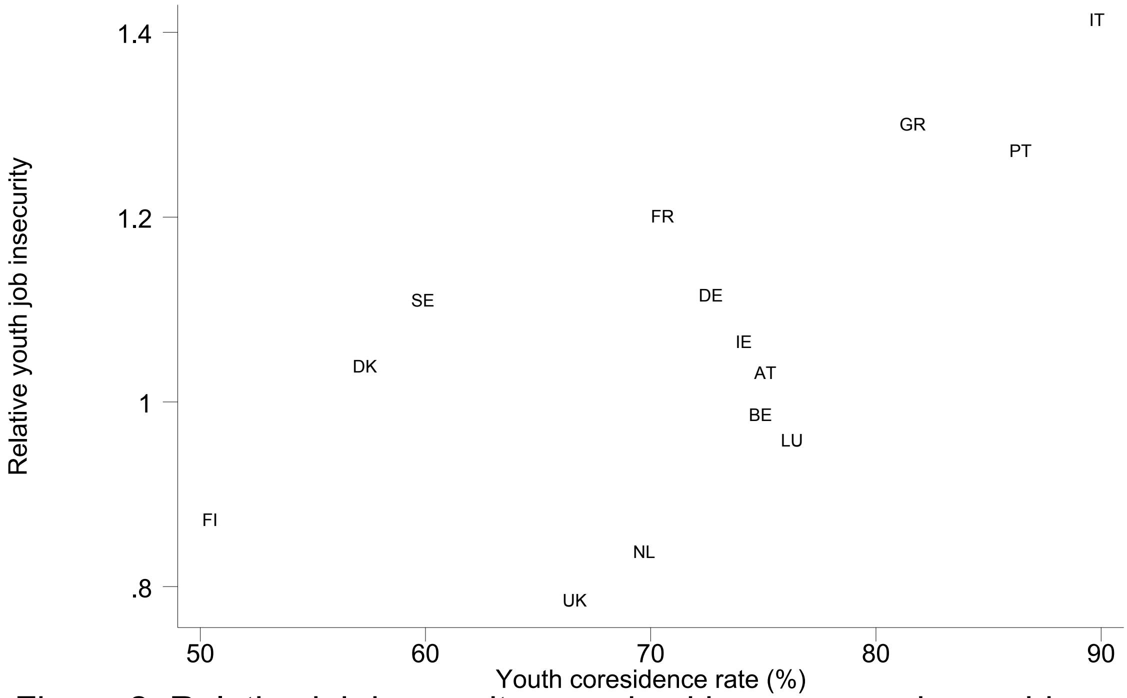

Figure 2. Relative job insecurity perceived by young and coresidence

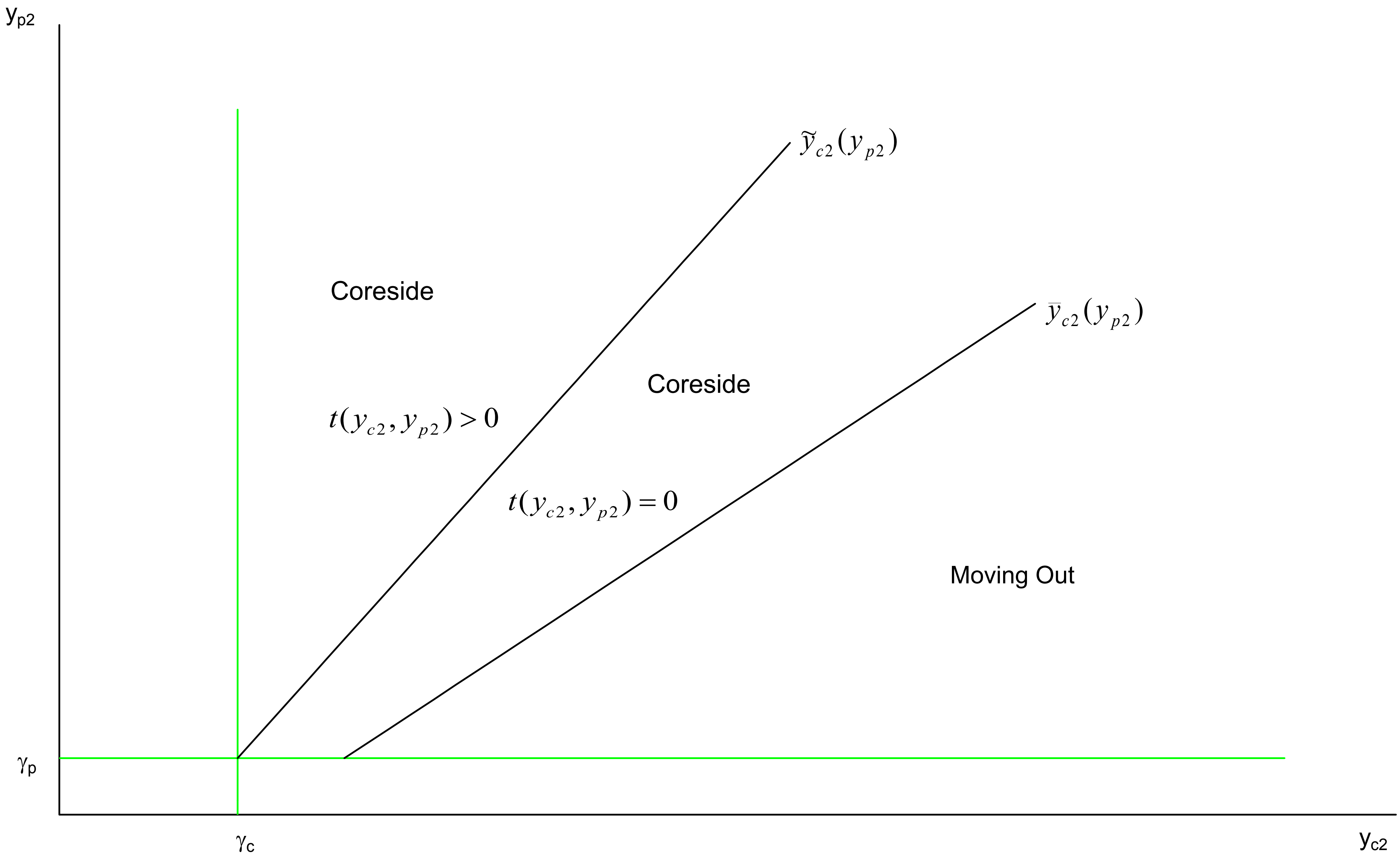

Figure 3. Period 2 residential regimes

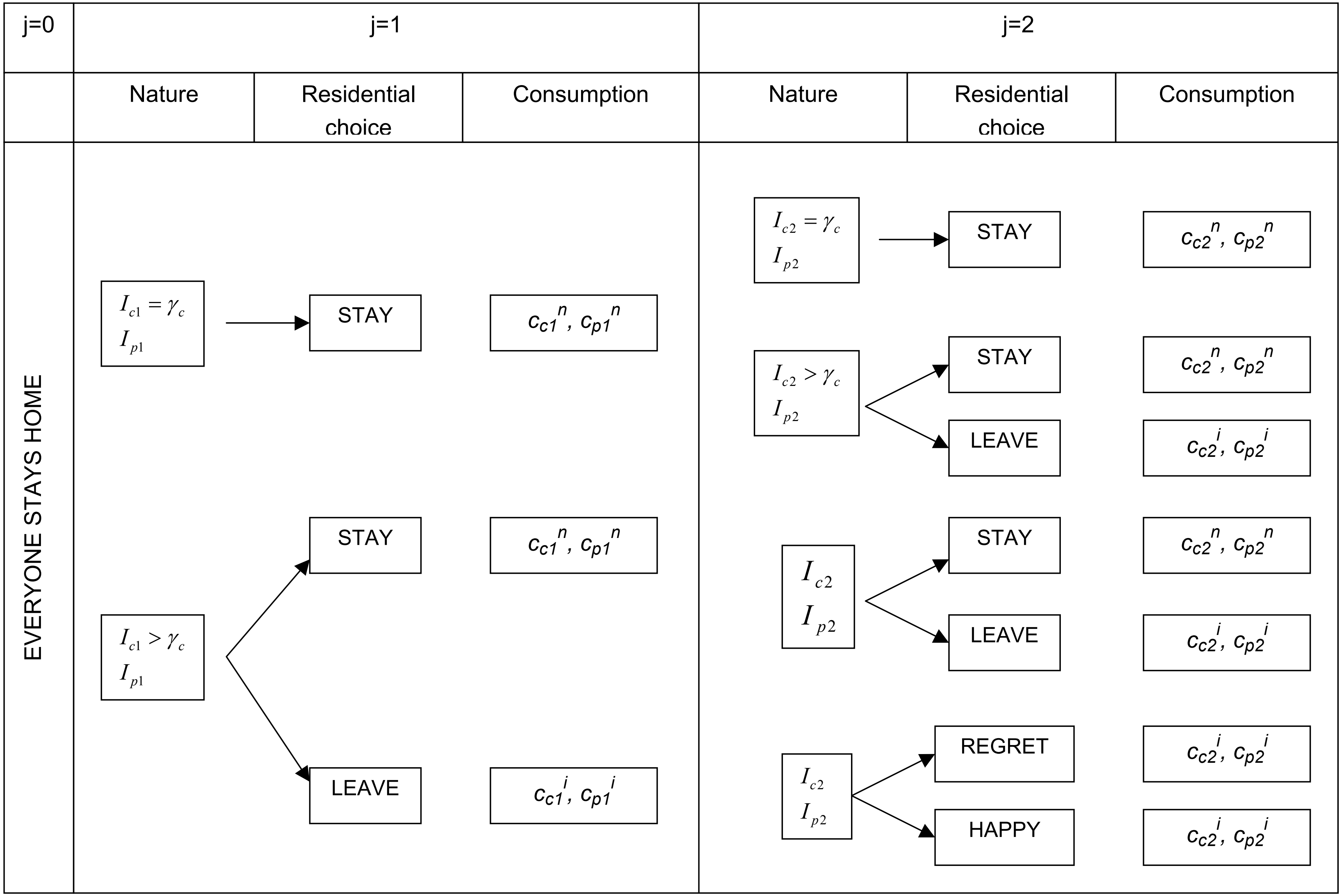

Figure 4. Structure of the model

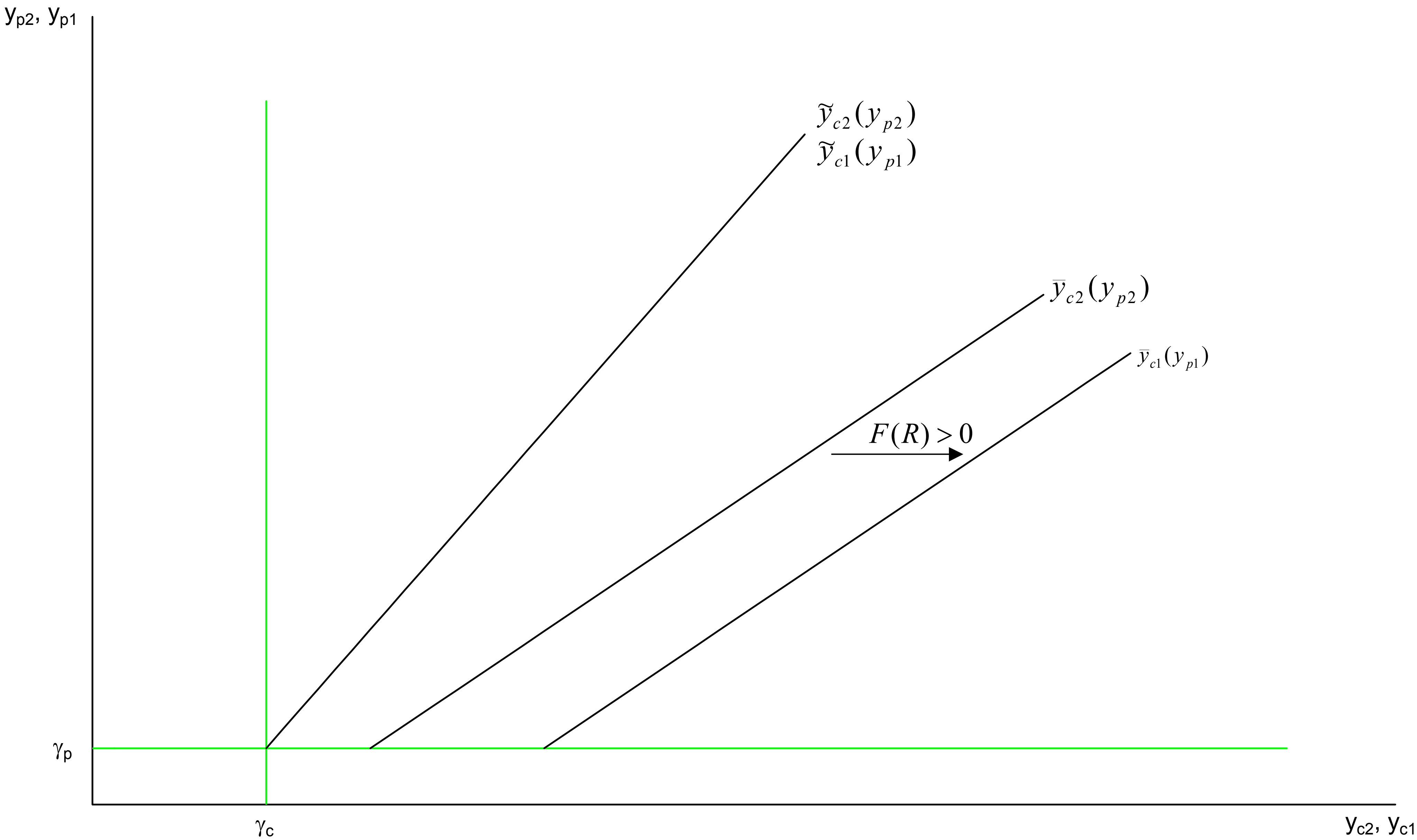

Figure 5. The effect of regret on the moving-out threshold