Youth Emancipation and Perceived Job Insecurity of Parents and Children*

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Abstract

In this paper, we test whether lower job insecurity of parents and higher job insecurity of children delay youth emancipation. Macroeconomic estimates for 13 European countries spanning the period 1983-2004 show that the coresidence rate increases by about 1.7 percentage points following a 10 percentage-point rise in the percentage of youths perceiving their job to be insecure and declines by about 1.1 points following the same increment in insecurity for workers aged 50-59. In the mid-1990s in Italy, 75% of youth aged 18-35 lived at home and had only a 4% probability of emancipation in the 3 subsequent years. Microeconometric evidence for this country shows that the probability of emancipation increases by about half a percentage point for a one-standard-deviation increase in paternal insecurity and by one-third of a percentage point for a one-standard-deviation decrease in children's insecurity.

Keywords: Coresidence, Youth emancipation, Job security.
JEL Classification: J1, J2.

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1 Introduction

The age at which children leave the parental home differs considerably across countries. In 2004 coresidence rates for men aged 25 to 29 years old ranged from 20 to 24% in France, the Netherlands, and the UK. On the opposite end, in Italy it was as high as 73%. Close to the Italian record are other southern European countries like Greece, Spain, and Portugal, with figures above 60%, but also a Nordic country like Finland (73%). Moreover, in the mid-1980s coresidence rates for that demographic group were around 50% in Italy, Greece, and Spain, and 38% in Portugal. Thus, there has been a sustained upward trend in these countries, with more stability in the remaining European Union (EU) countries.

The average emancipation age is well worth studying. To start with, it is negatively correlated with the interregional migration rate. This is around 0.5% in Italy, Portugal, or Spain, but between 1 and 2% in other OECD countries (OECD, 2000). Lower mobility entails higher equilibrium unemployment (Layard *et al.*, 1991). It also induces lower flexibility in responding to idiosyncratic regional shocks. Blanchard and Katz (1992) showed that high internal migration makes unemployment rate disparities across states to be scarcely persistent in the US, whereas they are very persistent in low internal migration countries like Italy and Spain (Decressin and Fatás, 1995, and Bentolila and Jimeno, 1998, respectively).

With a longer-term perspective, the emancipation age is also strongly related to fertility. In southern Europe young people most often leave home when they get married and, as noted by Giuliano (2007), these countries feature a very low incidence of out-of-wedlock births, e.g. 3% in Greece or 8% in Italy, vis-à-vis 37% in France or 54% in Sweden. Thus, household formation and procreation are being postponed. Indeed, total fertility rates (births per woman of reproductive age) have gone down dramatically in southern Europe between 1980 and 2000: from 2.2 to 1.2 in Spain, 2.2 to 1.3 in Greece, 2.2 to 1.5 in Portugal, and 1.6 to 1.2 in Italy. In contrast, other EU countries (bar Ireland), show stability or small declines in fertility over that period (World Bank World Tables, www.worldbank.org).

Low fertility has a crucial impact on many outcomes. It may be good: helping growth in less developed countries or alleviating congestion. But it may also cause problems. To take a prominent example, it hampers the sustainability of pension systems. By 2030, public pension payments are forecasted to reach 20.3% of GDP in Italy, 14.1% in Spain, and 13% in Portugal. These shares imply unsustainable paths for net financial liabilities and would require increases in the tax to GDP ratio of 11.4%, 7.4%, and 8.2%, respectively, just to keep net debt constant (Disney, 2000). These figures are on the upper side of the spectrum across Europe.

In sum, there are huge disparities in coresidence rates across countries and they matter for welfare. The economic literature on emancipation has focused mainly on parental and youth income and on housing prices (see Section 2). Here we focus on one factor which has not received much attention so far, namely the degree of job insecurity perceived by young people and their parents. Fogli (2004) presented a model featuring low parental job insecurity as a determinant of late youth emancipation, while Fernandes *et al.* (2008) study the roles played in the child's residential decision by parental and youth income expectations. As discussed in Section 2, the theory implies that, under some conditions, children's job insecurity lowers the probability of moving out, whereas parental job insecurity raises it. These are our hypotheses of interest here.

We also add to a recent strand of the literature by using subjective measures of job insecurity (see Section 3), in addition to more standard, objective measures. Looking first at macro data, we document, in Section 4, some stylized facts on coresidence and job insecurity in 13 EU countries from 1983 to 2004. We show that, after controlling for a host of factors, higher youth insecurity and lower parental insecurity are associated with higher coresidence. The results suggest that the rise in coresidence in the 1990s is related to the increase in the degree of job insecurity perceived by the young.

In Section 5 we exploit the panel data structure of the Italian Survey of Household Income and Wealth (SHIW), collected by the Bank of Italy, which, for fathers, contains high quality data on individual-specific perceived job insecurity. The data are unfortunately not sufficiently informative to construct subjective measures for children, for which we can consider only objective measures. We estimate probability models for whether children live independently after a given year, 1995, as a function of indicators of job insecurity of parents and children, and of a set of control variables measuring demographic, educational, and labor market characteristics. Few papers in this literature have exploited the panel structure of microeconomic datasets, indeed most papers present just cross-sectional evidence. Our microeconometric results are again consistent with the prediction that children's job insecurity lowers the probability of moving out, whereas parental job insecurity raises it. In Section 6 we present our conclusions.

It is important to remark that our macro and microeconomic pieces of evidence are complementary. In the former we have good information on youth perceived job insecurity but only approximate information on perceived parental insecurity, in a sample of 13 countries, through qualitative answers to a survey question, and we use it to address the determinants of the level of coresidence. In contrast, the microeconomic evidence is based on good information on perceived parental job insecurity but no information on perceived youth insecurity, for a single country (Italy), through the quantitative replies to a survey question, which we employ to examine the determinants of changes in coresidence. While the two sets of results are therefore not directly comparable, we interpret the consistency in the qualitative results we obtain from both analyses as providing robust evidence regarding the validity of the hypotheses of interest.

2 Job insecurity and coresidence

The economic analysis of moving-out decisions has been developed by McElroy (1985), Rosenzweig and Wolpin (1993), and Ermisch (1999), among others. Standard assumptions in this literature are: utility depends on consumption and housing, parents are altruistic but the child is selfish, individuals may exhibit a taste for privacy, and parents share income (consumption) and housing with the child when coresiding, and, if their income is high enough, they make transfers to the child when living apart.

It has been found that, under fairly general conditions, the higher the child's income, the higher the probability of living apart, since the child can avoid sharing her income with her parents and enjoy more privacy. Coresidence is more likely the higher is parental income, since then the child shares from a larger pie, unless parents have a strong taste for privacy—so that they give a larger transfer to their independent child. The effect of housing prices is ambiguous. Ermisch (1999) explains this results as follows. An increase in housing prices reduces the child's utility both at home and away. If parents do not respond by reducing their housing demand, then utility at home is constant, whereas it falls away from home, making the child less likely to leave. If parents' housing demand is elastic, however, then the decline in housing may lead the child to leave. In the Cobb-Douglas case of a unit elasticity of substitution between housing and consumption the price of housing does not affect coresidence.

For Anglo-Saxon countries, empirical findings confirm that higher child earnings reduce coresidence, whereas child unemployment raises it (Rosenzweig and Wolpin, 1993, for the US or Ermisch, 1999, for the UK). At the aggregate level, Card and Lemieux (2000) find that in regions with stronger local demand conditions and higher wages, young men are more likely to emancipate. Results are more varied for the father's earnings: they are found to raise coresidence in Ermisch (1999) and McElroy (1985), whereas Rosenzweig and Wolpin (1993) estimate a significant, negative effect if parents are divorced. Housing prices are found to weakly deter emancipation in Ermisch (1999). Regarding southern European countries, Manacorda and Moretti (2006) emphasize the income of parents, who are portrayed as bribing their children to stay at home longer, as the source of late emancipation in Italy. Negative effects of housing costs on emancipation are found by Giannelli and Monfardini (2003) for Italy and by Martinez-Granado and Ruiz-Castillo (2002) for Spain.[1]

In Fernandes *et al.* (2008) we extend the standard framework to analyze the effect on coresidence of uncertain income streams. Our model shares the assumptions listed above, except that utility depends only on consumption. The model has two periods and the moving-out decision is taken before observing the realizations of the child's and parental income. We also assume that moving back home is costly.[2] This gives rise to an option

[1] Alessie *et al.* (2006) link late emancipation in Italy to high transaction costs in housing.

[2] In fact, we assume that moving out is irreversible but show that this assumption carries no loss of generality compared to the case of finite moving costs.

value associated with waiting to see the income realizations before deciding whether to leave. In fact, a child who moved out may come to regret she did. For instance, if she experiences a low income realization she will tend to be worse off than if she had remained at home both because overall resources are lower (due the child's additional housing costs) and because parents are partially altruistic (and so their transfers do not fully compensate for the child's lower income). The same is true for unexpected, positive parental income shocks. We show that, under some conditions, when the child's income distribution shifts to the right—in the first-order stochastic dominance sense—the child is more likely to move out. The reason is that the shift reduces the probability and disutility of future regret and therefore makes it less likely that she might wish to go back home. Conversely, the same kind of rightward shift in parental income makes coresidence more likely by raising the probability and disutility of regret.[3] We also show that a higher variance of the child's future income holding the mean constant—i.e. under second-order stochastic dominance—makes the child more reluctant to leave, whereas the opposite is true for the variance of parental income.[4] Again, shifts in the income distributions matter to the extent they affect expected regret.

In this paper we study a factor which has not been examined empirically so far, namely the degree of job insecurity perceived by youth and their parents. Our key variable of interest is the perceived probability of becoming unemployed, labeled p. Drawing from Fernandes *et al.* (2008) we can infer the effect of p on emancipation decisions. Suppose that workers are either employed and receiving a wage or unemployed and receiving unemployment benefits. Since benefits are usually a fraction of the previous wage, p measures the probability that the worker will get benefits as opposed to her full wage. For this two-point support distribution of income, a reduction in perceived job insecurity (lower p) exactly captures the notion of first-order stochastic dominance used in the model.[5] Under these conditions and as long as transfers to independent children are rel-

[3]The conditions for these results are a low incidence of transfers between parents and their independent children and/or a low degree of altruism. Regarding Italy, evidence from Guiso and Jappelli (2002) suggests that the first assumption holds.

[4]The latter result depends on more restrictive assumptions on the distribution of parental income.

[5]The effect of p on the variance of income is, however, ambiguously signed, see Guiso *et al.* (2002).

atively unimportant (parents are not too altruistic), the theory predicts that an increase in the child's p should make emancipation less likely and an increase in the parents' p should make it more likely. These are the two key hypotheses we aim at testing.

Though not in the model, in reality emancipation can be reversed at a cost, which differs across families—depending, for instance, on parental preferences or on the available space at home. The higher the probability that the child will become unemployed, the higher the probability that she has to return and pay the reversal cost, so that moving out will be less likely for this reason as well.

Of course, we cannot be sure that the assumptions needed to obtain unambiguous predictions in our theoretical model strictly hold in reality. For instance, altruistic children might remain with their parents in the face of increased parental job insecurity in order to save on housing costs and to financially assist them. The predicted effects of child and parental insecurity are therefore an empirical issue and our estimation exercise aims at letting the data speak about them.

3 The measurement of perceived job insecurity

In empirical work, expectations are often replaced by outcomes—under rational expectations forecast errors are purely random. Recently, however, a burgeoning literature has shown the usefulness of individual expectations as measured in surveys (Manski, 2004). In this paper we use two different surveys to measure the perceived probability of unemployment, in addition to more objective measures, and so it is worthwhile to briefly take stock of this literature.

Manski (1990) started the analysis of survey data on the probability of unemployment. Dominitz and Manski (1997) point out that in surveys the probability of job loss may be confounded with its subjective cost. This is the case, for example, with the question in the European Community Household Panel (ECHP) on how satisfied respondents are with their job in terms of job security (Deloffre and Rioux, 2004; Clark and Postel-Vinay, 2008). The answer set matters too. For instance, the question in the US General Society Survey on the probability of job loss features the following answers:

"Very likely, fairly likely, not too likely, or not at all likely". Dominitz and Manski (1997) note that respondents interpret these answers in different ways, providing only ordinal information. They argue in favor of the probabilistic elicitation of expectations, as in the US Survey of Economic Expectations (SEE) question "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?". Our macroeconomic evidence is based on qualitative answers, whereas our microeconomic evidence relies on quantitative answers to a question very similar to the SEE one.

As to their determinants, Manski and Straub (2000) found that, in the SEE, expectations of job loss tend to decrease with age and schooling, being slightly higher for women than for men and substantially higher for blacks than for whites. However, personal characteristics explain only a small part of sample variation in expectations. Green *et al.* (2001) examine a job-loss question in the British Household Panel Survey over 1996-97. The fear of unemployment is found to increase with age, past experience of unemployment, having a fixed-term or a part-time contract, and regional unemployment. Interestingly, the level of fear is positively correlated with actual individual unemployment experience over the subsequent year.

International comparisons are provided by Böckerman's (2004) analysis of the question "Do you worry about the security of your present work?" in a 1998 survey covering 15 European countries. This question should capture the anticipated utility implications of unemployment. He finds, surprisingly, that individuals on temporary jobs perceive lower job insecurity, perhaps because they expect to find another job easily. He also finds large differences across countries and argues that perceived insecurity is correlated positively with the stringency of employment protection legislation and negatively with the generosity of unemployment benefits. Clark and Postel-Vinay (2008) find the same correlations with the ECHP data on job satisfaction.

Regarding trends, the OECD (1997) claimed that there was a widespread increase in perceived job insecurity between the 1980s and the 1990s in OECD countries. In the aggregate data for the EU used in the next section we shall see that there is also an indication of increasing insecurity from the early to the late 1990s, but not in the 2000s.

4 Macroeconomic evidence for the EU

In this section we test the hypotheses that higher own insecurity delays emancipation, whereas higher parental insecurity hastens it. We first describe the macroeconomic data used and then present and discuss the empirical results.

4.1 Data description

We measure coresidence through the aggregate fractions of men and women, 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 in 1983-2005.

Data availability for perceived job insecurity is more limited. We construct it from the European Commission's Eurobarometer, which includes the following questions:

• 1983 and 1984: "During the last year, have you (or someone in your household) worried about losing a job or not finding a job?: a lot*, a little, not at all.

• 1992: "And in the future, how great a risk do you think there is that you will become unemployed?": no risk, quite a low risk, quite a high risk*, a very high risk*.

• 1997: "How likely do you think it is that you may lose your job in the next few years?": 0%, no risk at all; 25%, low risk; 50%, fifty-fifty*; 75%, high risk*; 100%, definitely will*.

• 2004: "Would you say that you are very confident, rather confident, rather not confident* or not at all confident* in your ability to keep your job in the coming months?

Since the questions do not coincide in all surveys we transform the set of individual responses into a 0-1 dummy variable for being insecure, with the asterisks above marking the answers considered as 1. We then compute the age- and gender-specific fractions of individuals who report being insecure. The data from the 1980s is available only for France, Germany, Italy, and the UK, whereas it is at hand for 13 countries thereafter. We also compute perceived job insecurity for people aged 50-59 years old, who are representative of parents. Note that the 1983-84 questions confound the probability with the costs of job loss, whereas later questions refer to the probability alone. They are also less clean, in that they include other household members. We check below whether this makes a difference for the impact of insecurity on coresidence.

Table 1 summarizes the data for the five years (the Appendix gives details on definitions and sources). The overall coresidence rate is below 50%. It falls with age and it is lower for women than for men. Given the reduced number of countries observed in the 1980s, in the country breakdown of Panel C we only show data for 1992 to 2004. Coresidence is higher in Mediterranean, predominantly Catholic countries—Italy, Greece, Portugal, and Spain—and in Finland, than elsewhere. Over the full period, coresidence increases in the above five countries, as well as in France; Belgium and the Netherlands feature mild upward trends; Ireland, Germany, and the UK show some reductions among those aged 20-24, and stability in the 25-29 years-olds.

The table shows that job insecurity is negatively correlated with age and that older females feel slightly more insecure than men. Across countries, young workers feel most insecure in France, Spain, Greece, Italy, the UK, and Finland, while older workers feel least insecure in Luxembourg, Austria, Italy, Ireland, the Netherlands, and Germany. Thus, of the five countries with the highest coresidence rates, four also exhibit high insecurity among youth, while Italy also shows very low insecurity among older workers.

We cannot discuss long-term trends, given the limited sample of countries available for the 1980s. But it is interesting to observe that in 1992-1997 the coresidence rate as well as youth and older-worker insecurity rose; then over 1997-2004 coresidence rose and both measures of insecurity fell. Thus, while in each period one type of insecurity has the potential for explaining the evolution of coresidence, a more detailed, multivariate analysis is called for.

4.2 Results

4.2.1 Baseline specification

We test the hypotheses of interest with several empirical specifications. Our baseline equation runs $Coresidence$ rates on $Youth\ insecurity$ and insecurity perceived by the older age group ($Insecurity$ 5059):

$$\begin{aligned} Coresidence_{ijt} &= \sum_{i=1}^{12} \beta_{0i} Country_i + \beta_1 Age\ 2529 + \beta_2 Female + \beta_3 log(Real\ GDPpc_{it}) \\ &\quad + \beta_4 Youth\ insecurity_{ijt} + \beta_5 Insecurity\ 5059_{it} + e_{ijt} \end{aligned}$$

where i denotes countries, j age-gender cells, and $t = 1983, 1984, 1992, 1997, 2004$. $Country_i$ denotes a full set of country effects; $Age\ 2529$ and $Female$ are dummy variables for those groups; $log(Real\ GDPpc_{it})$ is the (log) national real GDP per capita at purchasing power parity, and e_{ijt} is random noise. We report standard errors with clustering for age-gender-country cells. See the Appendix for a description of the variables.

We start with an equation where, rather than including GDP per capita, we include year dummies. These will capture any aggregate effect, beyond that of GDP. Estimation results for this coresidence equation are shown in Table 2. The age dummies confirm that coresidence is lower for the 25-29 year-olds and the year dummies indicate an upward trend since 1992.

Females emancipate earlier than men, so that coresidence is on average 15 percentage points lower for them. Why? An important channel is marriage or living in a couple. In 1980, in our reference countries, the average age at first marriage was 23.5 years old for females and 26.2 years old for males, increasing to 27.6 and 30.0, respectively, by 2003.[6] In other words, females married 2.7 years earlier than men in 1980 and, although both groups have delayed the age of marriage, the difference between them has narrowed only slightly, to 2.4 years. Emancipation and living in a couple are closely linked, especially for women. In 2005, 55% of emancipated women aged 20-24 lived in a couple and a full 74% of those aged 25-29 did, while the respective figures for men were 42% and 64%.[7] These facts match well with the increase in the emancipation age over time and with the difference in the median age of emancipation between men and women, which in 2005 was equal to 2.5 years.[8]

The country dummies (not shown) also confirm the cross-country differences in Table 1: Finland, Greece, Italy, Portugal, and Spain show significantly higher rates—ranging from 20 to 30 extra percentage points—than the other countries. These dummies are quantitatively very important: when they are excluded from the regression the R^2 drops from 0.95 to 0.64. Conceptually, they control for country-specific factors affecting cores-

[6] The figures exclude Finland, Ireland, and Luxembourg due to lack of data for either year (Eurostat (2002), Table 1, section.1.3.13, for 1980, and Eurostat (2008), Statistical Annex Table A.6 for 2003).

[7] In the same countries as for the age of marriage (Eurostat, 2008, Statistical Annex Table A.10).

[8] Eurostat (2008), Statistical Annex Table A.12.

idence. In particular, they capture in part cross-country cultural differences, i.e. in preferences about coresidence—importance of family ties, attitudes regarding partnership formation, taste for independence, etc. Some evidence in favor of this interpretation is given by Giuliano (2007), who finds for 1994-2000 that second-generation immigrants in the US aged 18-33 whose parents came from Italy, Greece, and Portugal were more likely to coreside than those from other countries, whereas those with UK origin were less likely to coreside. These differences are identified as arising from culture by the finding that only immigrants with a background from the first three countries plus Spain experienced a significant increase in coresidence after the "sexual revolution" of the late 1960s reduced the privacy cost from coresiding. Similarly, Algan and Cahuc (2007) find that national attitudes regarding the roles of women and of young and older people in the family affect employment rates in a panel of OECD countries.

Column (1) of Table 2 reveals that youth job insecurity significantly raises coresidence, whereas the job insecurity of the older group lowers it. 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 1.6 points, whereas for insecurity of workers in their 50s the effect is a reduction of 1.2 points.

We now turn to the specification with GDP per capita. To the extent that it captures the current income of the parents, we expect a positive coefficient on this variable, though it could capture other effects as well. From column (2) of Table 2, at sample mean values, the estimated elasticity of GDP per capita is 0.35. The effects of the insecurity variables are little changed: if the percentage of youth feeling insecure rose from 0 to 100, the coresidence rate would increase by 17 percentage points, while the same change in the share of insecure older workers would reduce coresidence by 11 points. These are large changes, given that the average rate is 48%.

We also examine the impact of including objective measures of the state of the labor market, namely the unemployment and temporary employment rates. As in other data sets (Böckerman, 2004), in our sample the correlation between perceived job insecurity and the unemployment rate is far from perfect, 0.53 for young workers and 0.58 for older ones, while for temporary employment as a share of all employees the correlation coefficients are, respectively, 0.50 and 0.30. This leaves scope for perceived and objective measures to play separate roles.

We have tried several specifications. The first one adds unemployment. Our measure of insecurity is the workers' expectations of losing their jobs, corresponding to the inflow rate into unemployment. Since the same unemployment rate can be associated with different inflow and outflow rates (see e.g. Machin and Manning, 1999), having included insecurity in the regression, the unemployment rate will tend to pick up variations in the outflow rate. Column (3) of Table 2 presents the results from including insecurity and the unemployment rate for both the young and the older group. We expect coresidence to be jointly determined with the youth unemployment rate, and so the model is estimated by instrumental variables.[9] To control for the possibility that youth insecurity may also be endogenously determined, we instrument insecurity as well. Due to the lags involved, the sample size is reduced to 108 observations. Youth unemployment shows the expected sign whereas older-worker unemployment shows an unexpected positive sign, but neither variable is significant. At the same time, the coefficient on youth insecurity doubles with respect to the estimates in column (2), while that on older-worker insecurity jumps up even more, and both variables retain their significance.[10] The income elasticity drops slightly.

A further test involved including the temporary employment rate of young and older workers as additional objective insecurity proxies. Again the former group showed the expected negative sign and the second the opposite, but neither was significant. Lastly, we also run a "horse race" between our insecurity variables and alternative measures of the expected inflow rate into unemployment, namely future (one year ahead) changes in the youth and older-worker unemployment rates, both including and excluding the two temporary employment rates. None of these variables turned out to be significant.

[9] We use the same instrumental variables in all specifications to avoid any dependence of the results on variations in the instrument set. The instruments are one-year lags of: the coresidence rate, the youth unemployment rate, the unemployment rate of workers aged 50-59 years old, the youth temporary employment rate, and log GDP per capita. The first step regression shows that the instruments are strongly correlated with the instrumented variables (see the note to Table 2).

[10] As is to be expected, the exclusion of the two unemployment rates from the equation hardly alters the coefficients on the insecurity variables.

Overall, the results indicate that controlling for the current state of the labor market does not alter the usefulness of insecurity measures in explaining coresidence.

Lastly, we should note that in the preceding specifications differences in housing costs are captured by the country dummies. Measuring them directly through housing prices is difficult. National housing price indices, available from the Bank for International Settlements, are not comparable across countries due to heterogeneity in definitions. For this reason we included the housing inflation rate rather than the levels in the equation. The data set was reduced to 120 observations. The estimated coefficient on housing price inflation was positive and equal to 0.050, but it was not significant (p-value: 0.19). Although in line with the ambiguity of theoretical predictions, this result is likely to be affected by measurement error.

4.2.2 Robustness checks

A potential challenge to the validity of the foregoing hypothesis regarding own perceived job insecurity stems from the fact that women emancipate earlier than men. Table 1 indicates that, on average, 20-24 year old women perceive the same insecurity as men, whereas the 25-29 year-olds feel more insecure. Thus, according to that hypothesis, women should emancipate at the same time in the first age group and later in the second. Nevertheless, while for all age-country-year cells the coresidence rate of women is lower than for men, there is substantial underlying variation in relative perceptions of insecurity across gender groups. Women in the 20-24 year-old group actually perceive less insecurity than men in 6 out of the 13 countries and in 5 out of the 13 countries in the 25-29 year-old group. Overall, insecurity is lower for women than for men in 44% of all age-country-year cells. Thus, this variable is essentially balanced across genders so that we should not expect a strong difference in coresidence stemming from insecurity.

However, the early arrival of marriage offers for women could also confound the effect of insecurity on emancipation. For instance, for women who emancipate to live in a couple, their partner's job security could matter more than their own, especially in countries where, for young workers, there is a gender wage gap in favor of men.[11] We

[11] Such gaps are pervasive (see, e.g., Blau and Kahn (1992).

shall try to capture this possibility below, but before showing how, let us first discuss age effects.

Coresidence is likely to be jointly determined with young people's schooling status. In many European countries, especially southern ones, people tend to stay at home while attending college. One way to tackle this issue is to estimate the coresidence equation including the share of individuals in the gender-age cell who are studying as a regressor. We can construct that share from a question in the European Labor Force Survey asking respondents whether they had received education or training during the previous four weeks. This is a proxy for a more adequate variable such as the (unobserved) share of individuals in full-time schooling. We estimate the extended equation instrumenting the schooling rate with its first lag and the sample is reduced to 112 observations. The schooling rate attracts a coefficient of 0.26 (p-value: 0.11) and the coefficients on youth and older worker insecurity remain highly significant. While we would expect a stronger association between schooling and coresidence, this weak result may also be due to measurement error in the schooling rate (given its proxy nature).[12] Since we expect young people who are studying to be less sensitive to job insecurity, we make another attempt at capturing this effect in a different way below.

Since, as just seen, there are reasons why perceived job insecurity may affect emancipation differently by gender and age, we investigate these issues within the same empirical specification, by interacting the insecurity variables with demographic group dummies.[13] The results, presented in column (1) of Table 3, suggest that own insecurity affects women less than men, in line with the reasoning above. It appears that parental insecurity affects women more than men, which could arise if in their emancipation decisions children take into account expected transfers from parents when living apart, if those transfers do not allow children to achieve the same level of utility as at home (as in Fernandes et *al.*, 2008), and, lastly, if parents are more likely to make transfers to daughters than to sons (an issue that has not been settled in the empirical literature).[14]

[12] Note that we are better able to control for schooling in the micro data in Section 5.

[13] We also tried including the interactions of GDP with age and gender, as well as the interaction of the latter two variables, but none was significant.

[14] Cox (2003) remarks that *inter-vivos* transfers do not seem to be shared equally among offspring and

On the other hand, the results indicate that job insecurity, both their own and their parents', affects the older group of children more strongly, which is consistent with the above-mentioned schooling effects.

While the estimated effects from the interactions are large, often leading to either doubling or halving the estimated effect of insecurity vis-à-vis the reference group—males aged 20-24—they are however not well determined (their p-values range from 0.2 to 0.4), possibly due to the low number of observations, so that testing of these channels must await further research.

Secondly, as previously indicated, the questions on job insecurity differ across surveys, especially those in 1983-84 with respect to the rest. The correlation coefficient for the indicator of job insecurity between 1984 and 1992 is 0.7, dropping to 0.45 between 1984 and 1997. Overall, the correlation across all years is around 0.4. But we should bear in mind that for 1983 and 1984 we only observe insecurity in 4 countries and that perceptions of insecurity are liable to change over time, due to both the business cycle and changes in labor market institutions (for example, there were important labor reforms in the UK in the second half of the 1980s and early 1990s, see Petrongolo and Pissarides, 2008).

To check whether question wording this matters, we interact the insecurity measures with a dummy variable for 1983-84. From column (2) of Table 3, there is no apparent effect from definitional changes on the impact of insecurity, since neither interaction is significant at all. Alternatively, we re-run the specification in column (1) on the data excluding the 1983-84 observations. The results are qualitatively the same as before, except that parental insecurity is no longer significant at 5% (the p-value is 0.12). But recall that parental insecurity is measured in a less direct way than own insecurity (fortunately, it is better captured in Section 5).

In sum, the aggregate evidence for European countries we have uncovered indicates

that parents appear to target transfers toward liquidity constrained children. For the US, Cox (1987) found that females living independently have a higher probability of receiving *inter-vivos* transfers from parents than males, *ceteris paribus*, and that they receive higher amounts, though the latter effect was not significant. In Cox and Rank (1992) both the probability and amount of transfers are found to depend negatively on being an unmarried female, but when distance is used as a proxy for children's services, that variable affects positively the probability of transfers and negatively their amount.

that, once persistent demographic and cross-country differences are controlled for, job insecurity of children and parents influence the coresidence choices of European youth. This can help us account for both differences in the levels of coresidence rates across European countries and trends in coresidence. Let us illustrate this for the country analyzed with microeconomic data in Section 5 below, namely Italy in the 1990s. Take the group with the highest coresidence rate overall, men aged 20-24 and start with the cross-section variation. In 1997, 92.4% of Italian men in that age group lived with their parents, and the model in column (2) of Table 2 provides an accurate prediction, 92.0%. Youth insecurity was moderate, at 28.8%, while parental insecurity was low, at 11.6%. If these young men instead had perceived the lowest insecurity rate of this age-gender cell across countries in that year (22.6%, in Germany) and their parents the highest (31.9%, in France), then their coresidence rate would have been 3.1 percentage points lower, *ceteris paribus*. If we instead use the IV estimates in column (3), the predicted fall is much larger, 12.5 percentage points.

Now take the change over time. In 1992 the coresidence rate was 90.3%, and youth and parental insecurity were, respectively, 13.8 and 8.2 points lower than in 1997. Again, the model in column (2) predicts a 2.4 percentage point increase in coresidence from 1992 to 1997 due to the increase in youth insecurity—5.3 points using the IV estimates in column (3)—though the predicted increase is actually lower due to the countervailing impact of the increase in parental insecurity. Taken together, these results imply that coresidence in this demographic group is relatively high in Italy partly because parents are very secure in their jobs—vis-à-vis other countries—and that the increase over the 1990s was related to a rise in youth insecurity, whose effect exceeded that of the increase in parental insecurity.

5 Microeconomic evidence for Italy

In this section we extend the macro evidence with micro evidence based on Italy in the second half of the 1990s. Among the European countries in which youth coresidence has recently reached very high rates, Italy is the only one for which we could find household panel data containing information on individual-specific perceived job insecurity in addition to objective measures from local labor markets.

5.1 Data and sample design

We use a representative sample of Italian individuals of working age, between 18 and 35 years old, and living with their parents. This sample has been extracted from the Italian Survey of Household Income and Wealth (SHIW). We use its 1995 wave, which contains information on 8,135 households and 23,924 individuals, to select our baseline sample. We then use the 1998 wave to obtain information on whether a child left home between 1995 and 1998. Our primary goal is to test whether measures of job insecurity of the father and the child affect this decision, controlling for observable confounding factors measured in 1995.[15]

A first reduction of the initial sample is due to the fact that the SHIW is a rotating panel (Banca d'Italia, 1997,2000). The emancipation decision of children (the *outcome*) can only be observed for the 2,699 households (out of 8,135) that were interviewed in both 1995 and 1998. Note, however, that since these panel households were randomly selected, 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 use a measure of perceived job insecurity, described in detail below, constructed from the answers to a survey question in which individuals were asked about the probability of having a job in the following year. This question was designed to measure various dimensions of uncertainty and it 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 were either working or unemployed were 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,

[15] 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.

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 being 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 1,142 children; note, however, that this sample is still representative of the population of children living in households where fathers were not retired in 1995.[16] Its characteristics are described in the first two columns of Table 4.

The second problem related to the question on perceived uncertainty is that, to limit the questionnaire length, labor force members in *all* households were not asked this question but only those in households in which the (male) head was born in an odd year.[17] Thus, our measure of perceived insecurity 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, we observe a measure of perceived job insecurity for the fathers of only 479 of the children described in the first two columns of Table 4 and moreover we have an analogous measure for only 212 of these 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 1,142 households for whom we have two years of data (see the last two columns of Table 4), this is not true for the 212 households for which the information is available for both fathers and children. The reason is that not all individuals born in odd years were asked the question, but only individuals belonging to households in which *the father* was born in an odd year. For this reason we cannot construct a reliable and representative measure of perceived insecurity for children, and to assess the effect of their job insecurity on coresidence we have to rely on objective measures that will be described in the next section.[18] Thus we

[16] 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.

[17] Different questions were asked in households where the head was born in an even year.

[18] A careful reader may wonder why the information on job insecurity is available only for 479 fathers

cannot build a reliable measure of perceived job insecurity for children from the SHIW. Finally, due to the presence of siblings, the 479 children that constitute the restricted sample originate from 298 families. For this reason, the standard errors of our estimates are clustered by family.

Before describing in detail our indicators of job insecurity and emancipation, let us note again that while data limitations force us to use a relatively small sample, it is still representative of the population of interest (see Table 4). Moreover, its timing structure is suitable for exploring the relationship between paternal job insecurity and the subsequent (not contemporaneous, as in most of the related empirical literature) decisions of children to leave home controlling for a large set of individual and family background characteristics.

5.2 The indicators of job insecurity and the outcome variable

In what follows we explore the effects of both subjective and objective job insecurity measures. For the first our key variable is the reply to the following question, posed to employed and unemployed individuals:[19]

> 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, that of unemployment. Note that this question aims at eliciting the probability of job loss, not its costs.

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

and not for $571 = 1,142/2$ 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 4 clearly shows, however, that the cases of non-response are randomly distributed in the data.

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

are excluded. Those who expected to voluntarily retire or drop from the labor force are excluded. The 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 of being unemployed in the year following the interview.[20] The authors also compare this source, restricting the sample to those employed, with the above-mentioned Survey of Economic Expectations. 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.

Table 5 reports the distribution of the perceived insecurity indicator for fathers in our sample of 479 households. As expected given the sample design, our sample is not very different from the full sample 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.

It could be argued that such an indicator of perceived insecurity is endogenous and less informative than measures of local unemployment by age and gender. We think that neither claim is correct, for the following reasons. First, for fathers, it is unlikely that the subjective perception of the likelihood of being employed in the future reflects a labor supply decision.[21] In other words, it is unlikely that it might capture a situation in which the father has decided not to work and thus expects not to have a job. Moreover, it does not seem plausible, given the observed very high participation rates for fathers, that they would stop working in order to make their children leave home. Thus, the expectation of future unemployment by fathers is likely due to a truly exogenous factor more than to an increase in the preference for leisure, and can therefore be considered exogenous for our purposes. As to the second claim, in the case of fathers, who are

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

[21] Note that such a claim would instead be more plausible for children.

typically characterized by employment rates which are very high and constant across provinces and age groups, it is most likely that a subjective measure of job insecurity is better than local unemployment as an indicator of the individual-specific degree of insecurity faced by prime-age males.

Nevertheless, we also construct objective (and arguably exogenous) measures of insecurity based on labor force information by age, gender, and province. These measures are also the only ones we can construct for children since, as argued in the previous section, the number of observations for which we have perceived job insecurity of fathers *and* children is too small.

To construct objective measures of insecurity we have obtained from the Italian statistical office (Istat) information from the Quarterly Labor Force Statistics on the fraction of unemployed and of temporary employees within cells defined by age brackets, gender, and provinces.[22] We were forced to construct our measures for the age brackets pre-defined by Istat, which are not exactly the ones we would have preferred. As a result, the measures for fathers are based on information from the 30-64 age bracket, while the measures for children are based on information from the 15-29 age bracket. For both brackets we compute the change in the fraction of unemployed and temporary workers between 1995 and 1998, by gender and province. This amounts to saying that for both fathers and children's job insecurity increases when unemployment and temporary jobs grow in their local labor market. In this case, as opposed to the macroeconomic specification, the levels of these variables are captured by geographic area dummies.

Lastly, the outcome variable 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 4. In our sample of children living with their parents in 1995, only 4% decided to leave home over the following three years. When matched with the 1995 wave, the 1998 wave of the SHIW features an apparently low moving-out rate. Preceding waves had larger panel-sample rates: 14% from 1991 to 1993 and 8% from 1993 to 1995. But this is consistent with the aggregate Italian data: the coresidence rate for people aged 20-29—

[22] The Italian territory is divided into 105 administrative units called provinces, which correspond approximately to US counties.

which represent 75% of our sample—rose by 2.9 percentage points from 1991 to 1995 and by another 2 points from 1995 to 1998.

5.3 Results

The first column of Table 6 reports estimates based on a probit model of the marginal effect of job insecurity measures for fathers and children on the probability that children leave home within three years from the baseline.

Ideally, we would like to base our estimates on a comparison of children who are identical with respect to all relevant personal and family characteristics potentially affecting the outcome in order to identify convincingly the effect of job insecurity. We try to approximate this ideal condition by controlling for a large set of variables dated in 1995, when all children are observed coresiding.

Emancipation decisions are likely to be affected by both family traits and the current situation in the household. Thus, we condition on the father's age and completed years of schooling. Note 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—perceived job insecurity is measuring the probability that the father will get unemployment benefits as opposed to his full wage. As indicated in Section 2, in this setting a reduction in job insecurity exactly captures the notion of first-order stochastic dominance used in Fernandes *et al.* (2008).

We also control for family wealth, home-ownership (owner-occupied = 1), number of children in the household, employment status of parents, and rental prices at the province level.[23] Local conditions are further controlled for by the inclusion of five geographical area dummies. As far as the characteristics of children are concerned, we control for age, gender, schooling, and employment status.

The estimate in the first row and first column of Table 6 indicates that insecurity perceived by fathers has a positive effect on the probability of emancipation. The magnitude of this effect can be inferred from the observation that if the father goes from being

[23] We also tried including a dummy variable capturing whether the household owned a second house, but it was not significant.

sure to be employed next year to being sure of being unemployed, the probability that the child leaves home increases by 1.7 percentage points. Despite the small sample size, this estimate is significantly different from zero and large, since the average probability of emancipation in the sample is 4%.[24]

Objective measures of paternal insecurity, in the second and third row of column (1) in Table 6 confirm the conclusion based on perceived insecurity. A one percentage point increase in the fraction of male unemployed workers aged 30-64 in the province increases the probability of emancipation by one third of a percentage point. This effect is also statistically significant and it should be evaluated with respect to the benchmark average probability of emancipation in the sample of 4%. An increase in the fraction of temporary workers has a positive effect but the estimate is not significant. This is what we would have expected given that temporary jobs are relatively infrequent at old ages, whereas local unemployment is a better measure of the degree of objective job insecurity faced by fathers.

Estimates of the effects of objective measures of job insecurity for children are reported in the fourth and fifth row of column (1) in Table 6. In this case both indicators of insecurity have a negative effect on emancipation, although only the increase in temporary employment is significant. The estimate indicates that a one-percentage-point increase in the fraction of temporary jobs for 15-29 year-old workers in the gender-province cell reduces the probability of emancipation by two thirds of a percentage point.

To get a sense of how these estimated effects compare in terms of size, we compute the change in the probability of coresidence induced by a one standard deviation (SD) change in the measures of insecurity associated with each estimate, limiting ourselves to statistically significant results.[25] Starting with fathers, a one-SD change in perceived insecurity is equal to 0.28 and it would induce a change of half a percentage point (0.48

[24] It is interesting to observe that our setting is ideal for using the Multiple Imputation method proposed by Rubin (1987,1996) to impute perceived job insecurity to the fathers for whom this information is missing. The reason is that, by construction, information is missing "completely at random", which is a necessary condition for Rubin's method. As expected, when we use this method our results remain unchanged in terms of the size of the coefficients but they gain remarkably in terms of efficiency. These results are omitted to save space but are available from the authors upon request.

[25] See Table 4 for descriptive statistics on the standard deviations of the insecurity measures.

$= 0.28\times0.017\times100$) in the probability of coresidence. Interestingly a one-SD change of objective insecurity for fathers (measured by the increase in unemployment) has exactly the same effect ($0.48 = 0.016\times0.30\times100$). The effect of a one SD change of objective insecurity of children (measured by the increase in temporary employment) is smaller, amounting to slightly less than one third of a percentage point ($0.31=0.051\times0.06\times100$).

It may be argued that in the case of subjective measures of insecurity causality runs in the opposite direction. For example, a father may decide not to take a more uncertain job because he expects the child not to leave home. Note that our controlling for a large set of relevant family and personal characteristics does not exclude this interpretation. Even if these controls allowed us to take care of all potential confounding factors, which is a necessary condition for identification, they would not exclude the possibility of reverse causation implicit in the alternative explanation of the evidence proposed above. However, this alternative explanation is clearly not compatible with the evidence based on objective measures of insecurity, for which reverse causality is out of the question. We conclude from the joint consideration of the evidence that our hypotheses are not rejected: the probability of emancipation increases with fathers' insecurity and decreases with children's insecurity, and the causal interpretation of these effects is plausible.

As for other relevant variables, the probability of emancipation is statistically unrelated to the child's age, schooling, and gender, whereas it is significantly lower for children not working in 1995. Father's age, education, and current employment status have no statistically significant effect on emancipation. Higher family wealth and home ownership increase the probability of emancipation while higher rental prices in the province seem to discourage it.

In columns (2) and (3) of Table 6 we change the measure of perceived insecurity of fathers looking at the effect of uncertainty regarding future paternal income on the probability of emancipation. All the other regressors are unchanged. As discussed above, we conjecture that when parental income uncertainty increases, the probability of moving out increases, holding expected income constant.[26] We can say something on this conjecture because the SHIW allows us to approximate the expected earnings

[26] See also Fernandes *et al.* (2008).

distribution of fathers. It asked participants the minimum, y_m, and the maximum, y_M, income they expected to earn if employed, and the probability of earning less than the midpoint of the support of the distribution, $Prob(y \leq (y_m + y_M)/2) = \pi$. Guiso *et al.* (2002) construct measures of income uncertainty by assuming two alternative distribution functions for earnings: uniform over the intervals $[y_m, (y_m + y_M)/2]$ and $[(y_m+y_M)/2, y_M]$, and triangular over the same two intervals. They then assume a point expectation for unemployment income, impute it for each individual, and compute the coefficient of variation, i.e. the ratio of the standard deviation to the expected value, for each individual in their sample. We use these computations to estimate a probit model like the one of the column (1) of the table, in which we replace job insecurity with the coefficient of variation of future expected income of the father.

Independently of the distributional assumption, in columns (2) and (3) of Table 6 we obtain positive and significant estimates of the effect of paternal uncertainty on the probability that the child leaves home. These estimates provide favorable evidence for the prediction that children will tend to move out more often when their father's income is perceived as being more uncertain. The effect of objective measures is basically unchanged with respect to the estimates of column (1).

6 Conclusions

In this paper, we have explored the determinants of the youth's decision to leave the parental home. Our key insight is that this decision may depend on the degree of job insecurity experienced by parents and children. Specifically, we have tested the conjecture that higher own insecurity induces children to leave the parental home later, whereas higher expected parental insecurity has the opposite effect.

The aggregate evidence for 13 European Union member countries since the 1980s on coresidence rates and perceived job insecurity is consistent with these hypotheses. According to our estimates, for every 10 percentage-point rise in the percentage of youths feeling that their job is insecure the coresidence rate increases by about 1.7 percentage points, whereas the same increment for workers aged 50-59 reduces coresidence by about

1.1 points. These results still hold, though estimated effects become more variable, especially for the latter variable, once we control for objective measures of insecurity, such as levels and changes of the unemployment and temporary employment rates. We read this evidence as indicating that perceived job insecurity is a relevant explanatory variable of coresidence decisions across countries, once differences in institutions, culture, and the state of the labor market are controlled for. The model implies, for instance, that the high Italian coresidence rate is the result of high parental job security and, in the 1990s, of rising youth insecurity.

We have been able to further validate these hypotheses using microeconomic panel data from the Italian Survey of Household Income and Wealth, collected by the Bank of Italy. Our empirical results indicate that the likelihood that young Italians aged 18 to 35 left the parental home between 1995 and 1998 is positively related to parental job insecurity and negatively related to children's job insecurity. More specifically, the probability of emancipation would have increased approximately by half a percentage point for a one-standard-deviation increase in paternal insecurity, and by one third of a percentage point for a one-standard-deviation decrease in children insecurity. These effects are again estimated beyond those found for changes in the local unemployment and temporary employment rates.

What are the policy implications of our analysis? 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 and family demographics. Employment protection legislation usually protects older workers vis-à-vis young ones, raising job security for the former and reducing it for the latter. Thus, our results imply that one of its unintended effects is that young people will emancipate later.

The main direct effects of late youth emancipation are low geographical mobility, reducing an economy's capacity to react to idiosyncratic regional shocks, and low fertility, which is already putting in jeopardy pension systems in southern European countries. Both of these problems are constantly debated but we are the first to provide quantitative evidence linking them to coresidence through the effect of job security. Coresidence also has beneficial 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. While what is socially desirable as far as these outcomes are concerned is debatable, our analysis shows that the effects of job security provisions for parents and children on youth emancipation should not be disregarded.

A Appendix: Description of macroeconomic data

Coresidence rate. Fraction of population living at parental home. Countries: all in the EU-15. Years: 1983-2005, though data start later for new EU members. Source: Eurostat, European Labour Force Survey (epp.eurostat.ec.europa.eu).

Perceived job insecurity. 0-1 dummy variable constructed from answers to questions asked in the Eurobarometers 19 (1983), 20 (1984), 37.1 (1992), 47.1 (1997), and 62.1 (2004) (See the text for wording of questions). Data are available for the following countries. 1983 and 1984: West Germany, France, Italy, and United Kingdom (Belgium and Ireland had missing values and had to be excluded). 1992 and 1997: Belgium, Western Germany, Greece, Spain, France, Ireland, Italy, Luxembourg, Netherlands, Austria, Portugal, United Kingdom and Finland. We construct the job insecurity variable from 27,659 individual observations for: 4 countries in 1983 and 1984 (3,839 and 3,853 observations, respectively) and 13 countries in 1992, 1997, and 2004 (6,433, 6,634, and 6,900 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 180 observations (16, 16, 44, 52, and 52, for the five years, respectively). Source: European Commission, Eurobarometer (ec.europa.eu/public_opinion).

Real GDP per capita. Measured in 1996 US dollars at purchasing power parity, Laspeyres index. Source: A. Heston, R. Summers and B. Aten, Penn World Table Version 6.2, Center for International Comparisons of Production, Income and Prices at the University of Pennsylvania, September 2006 (pwt.econ.upenn.edu). For West(ern) Germany the source is: Groningen Growth and Development Centre and The Conference Board (www.ggdc.net), Total Economy Database, January 2007, rescaled to 1996 as the base year.

Real house prices. House price index deflated by Consumer Price index. Available for Belgium, Germany, Spain, France, Ireland, Italy, Netherlands, United Kingdom, and Finland. Years: 1983-2004. Source: Bank for International Settlements Data Bank, national sources.

Unemployment rate. Source: Organisation for Economic Cooperation and Development, Corporate Data Environment, Labour Market Statistics (www.oecd.org). Data for 2005 and for the Netherlands in 1983-1986 were completed using International Labour Office, LABORSTA Internet, Yearly Data (laborsta.ilo.org).

Youth temporary employment rate. Years: 1983-2004. Source: Eurostat, European Labour Force Survey.

Fraction of youth studying. Years: 1983-2004. Fraction of youth who have received education or training during the previous four weeks. Source: Eurostat, European Labour Force Survey.

Table A.1 gives presents descriptive statistics of the variables used in Tables 2 and 3.

Table A.1
Descriptive statistics of macroeconomic data

	Mean	Standard deviation	Minimum	Maximum
Coresidence rate	0.477	0.242	0.055	0.924
Youth insecurity	0.274	0.146	0.000	0.617
Insecurity 50-59 years old	0.194	0.095	0.000	0.446
Log real GDP per capita	9.948	0.280	9.413	10.835
Youth unemployment rate	0.137	0.084	0.020	0.423
Unemployment rate 50-59 years old	0.060	0.029	0.013	0.139
Youth temporary employment rate	0.191	0.157	0.018	0.732

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Table 1: Descriptive statistics on coresidence and perceived job insecurity in Europe. 1983 to 2004 (%)[a]

		Coresidence rate		Perceived job insecurity			
				Youth		50-59 years old	
		Mean	Standard deviation	Mean	Standard deviation	Mean	Standard deviation
A. All		47.7	(24.2)	27.4	(14.6)	19.4	(9.5)
B. Age and gender:							
20-24 years old	Male	71.9	(12.4)	30.2	(15.6)		
	Female	56.6	(16.9)	30.4	(14.8)		
25-29 years old	Male	39.0	(18.6)	23.9	(12.8)		
	Female	23.2	(15.4)	25.3	(14.6)		
C. Country:							
Belgium		39.9	(20.2)	22.8	(9.0)	21.0	(2.6)
Germany		37.2	(20.4)	22.8	(10.0)	18.5	(4.5)
Greece		61.2	(15.7)	35.9	(10.3)	24.1	(5.9)
Spain		71.4	(17.6)	40.0	(18.9)	27.8	(6.6)
France		34.3	(20.6)	41.0	(11.2)	25.6	(4.6)
Ireland		47.0	(16.5)	27.7	(15.6)	15.6	(7.0)
Italy		72.1	(18.2)	29.1	(11.4)	12.5	(8.1)
Luxembourg		46.6	(22.2)	13.4	(16.0)	5.7	(6.2)
Netherlands		31.4	(22.2)	18.0	(12.4)	16.3	(3.2)
Austria		48.5	(22.0)	21.7	(6.7)	9.9	(7.3)
Portugal		62.8	(19.3)	22.7	(13.4)	20.2	(5.1)
U. Kingdom		31.8	(17.1)	28.4	(16.4)	26.2	(16.8)
Finland		70.5	(17.0)	28.3	(15.6)	19.6	(9.6)
D. Year:							
1983		37.6	(23.6)	30.1	(11.5)	26.2	(7.4)
1984		37.6	(23.4)	26.6	(8.7)	21.1	(7.3)
1992		46.1	(23.8)	29.0	(16.6)	20.1	(12.0)
1997		51.0	(24.3)	36.5	(10.9)	22.1	(6.4)
2004		51.9	(24.0)	26.5	(11.4)	13.7	(8.1)

[a] Coresidence rate: percentage of youth population living at parental home (Eurostat: Labor Force Survey). Perceived job insecurity: percentage of respondents who think that their job is at risk (Eurostat: Eurobarometers). Data are available for 4 countries in 1983 and 1984, and 13 countries in 1992, 1997, and 2004. Only the last three years' data are presented for country breakdown (Panel C). See the Appendix.

Table 2: Macroeconomic evidence on coresidence and perceived job insecurity[a]

	(1) Year dummies	(2) Real GDP per capita	(3) Youth unemployment
Youth insecurity	0.156	0.171	0.340
	(0.043)**	(0.045)**	(0.122)**
Insecurity 50-59 years old	-0.120	-0.107	-0.483
	(0.044)**	(0.045)**	(0.176)**
Age 25-29	-0.322	-0.322	-0.317
	(0.011)**	(0.011)**	(0.016)**
Female	-0.156	-0.156	-0.151
	(0.010)**	(0.010)**	(0.012)**
1984	-0.001		
	(0.005)		
1992	0.030		
	(0.016)		
1997	0.055		
	(0.019)**		
2004	0.086		
	(0.021)**		
Log real GDP per capita		0.165	0.130
		(0.044)**	(0.044)**
Youth unemployment			0.021
			(0.185)
Unemployment 50-59 years old			0.255
			(0.619)
Adjusted R^2	0.96	0.95	0.95
No. of observations	180	180	108

[a] OLS and IV (column (3)) estimates of the coresidence equation in 1983-2004 (selected years, see text). Standard errors in parentheses. The dependent variable is the coresidence rate. GDP per capita is in thousand 1996 US dollars (PPP). See data sources and descriptive statistics in Table 1 and in the Appendix. The reference cell is that of males aged 20-24 living in Belgium in 1983. A constant and country dummies are included in all regressions. In column (3) youth insecurity and youth unemployment are instrumented. The instruments are one-year lags of: the coresidence rate, the youth unemployment rate, the unemployment rate of workers aged 50-59 years old, the youth temporary employment rate, and log GDP per capita. The p-values of the F tests for the inclusion of the instrumental variables in the first stage are all below 0.01 (see Staiger and Stock, 1997). The statistical significance of the test that the underlying coefficient is equal to zero is denoted by: $p < 0.05 = *, p < 0.01 = **$.

Table 3: Macroeconomic evidence on coresidence and perceived job insecurity. Robustness checks[a]

	(1) Age/gender interactions	(2) Period interactions	(3) Short period & interactions
Youth insecurity	0.183	0.173	0.140
	(0.054)**	(0.044)**	(0.053)*
Youth insecurity × Age 25-29	0.092		0.050
	(0.079)		(0.074)
Youth insecurity × Female	-0.086		-0.076
	(0.071)		(0.066)
Insecurity 50-59 y.o.	-0.200	-0.117	-0.133
	(0.076)*	(0.049)*	(0.086)
Insecurity 50-59 y.o. × Age 25-29	0.077		0.127
	(0.090)		(0.089)
Insecurity 50-59 y.o. × Female	0.070		0.040
	(0.082)		(0.083)
Youth insecurity × Dummy 1983-84		-0.065	
		(0.119)	
Insecurity 50-59 y.o. × Dummy 1983-84		0.051	
		(0.149)	
Age 25-29 y.o.	-0.361	-0.322	-0.358
	(0.024)**	(0.011)**	(0.024)**
Female	-0.146	-0.156	-0.137
	(0.024)**	(0.010)**	(0.024)**
Log real GDP per capita	0.168	0.154	0.142
	(0.045)**	(0.048)**	(0.052)**
Adjusted R^2	0.96	0.95	0.96
No. of observations	180	180	148

[a] OLS estimates of the coresidence equation in 1983-2004 (selected years, see text). Standard errors in parentheses. The dependent variable is the coresidence rate. GDP per capita is in thousand 1996 US dollars (PPP). See data sources and descriptive statistics in Table 1 and in the Appendix. The reference cell is that of males aged 20-24 living in Belgium in 1983. A constant and country dummies are included in all regressions. The statistical significance of the test that the underlying coefficient is equal to zero is denoted by: $p < 0.05 = *, p < 0.01 = **$.

Table 4: Descriptive statistics for the Italian micro sample, 1995[a]

	Full sample		Restricted sample	
Variable	Mean	Std. dev.	Mean	Std. dev.
Outcome				
Out in 1998	0.04	0.18	0.04	0.20
Subjective measures of job insecurity of fathers				
Father's perceived job insecurity	–	–	0.16	0.28
Father's income uncertainty (uniform distr.)	–	–	0.25	0.50
Father's income uncertainty (triangular distr.)	–	–	0.24	0.50
Objective measures of job insecurity of fathers				
Change between t and t+1 in fraction of unemployed aged 30-64 in gender-province cell	0.005	0.016	0.005	0.016
Change between t and t+1 in fraction of temp. jobs of 30-64 year-olds in gender-province cell	0.008	0.017	0.009	0.017
Objective measures of job insecurity of children				
Change between t and t+1 in fraction of unemployed aged 15-29 in gender-province cell	0.004	0.035	0.004	0.035
Change between t and t+1 in fraction of temp. jobs of 15-29 year-olds in gender-province cell	0.021	0.047	0.027	0.051
Control variables				
Age	22.57	3.61	22.59	3.54
Female	0.45	0.50	0.44	0.50
Child not employed	0.70	0.46	0.70	0.45
Years of schooling	11.49	2.95	11.54	2.99
Father's age	52.23	5.37	51.85	5.11
Father's years of schooling	9.24	4.12	9.40	4.11
Father not employed	0.06	0.23	0.06	0.23
Home-ownership	0.71	0.45	0.72	0.45
Wealth	0.34	0.50	0.34	0.55
Number of children	2.37	1.04	2.36	0.87
Mother employed	0.35	0.48	0.36	0.48
Home rental index in province	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

[a] Descriptive statistics of variables measured in 1995 for the full sample of 1,142 children who: (a) lived with both of their parents in 1995, (b) belonged to households interviewed in both 1995 and 1998, (c) were aged between 18 and 35 years old in 1995, (d) had a father who was either employed or unemployed (i.e. not retired), and (e) were still alive, not in jail and not long-term hospitalized in 1998. Also for the restricted sample of 479 children (from 298 households) whose father answered the question concerning perceived uncertainty. Wealth is in billions of Italian liras. Data source: Italian Survey of Household Income and Wealth (SHIW). The rental index is in thousand liras per square meter.

Table 5: The indicator of perceived job insecurity in the Italian micro sample, 1995[a]

Value of the indicator	Percent	Cumulative
0.0	60.13	60.13
0.1	11.27	71.40
0.2	7.93	79.33
0.3	1.88	81.21
0.4	2.71	83.92
0.5	4.80	88.73
0.6	0.63	89.35
0.7	2.09	91.44
0.8	2.71	94.15
0.9	1.88	96.03
1.0	3.97	100.00
Total	100.00	

[a] Distribution of the indicator of job insecurity of fathers in the sample of 479 observations used in the econometric analysis (see Table 4). The indicator measures the probability assigned by the individual to the event that he does not work in the following year. Data source: Italian Survey of Household Income and Wealth (SHIW).

Table 6: Job insecurity and the probability of children's emancipation in the Italian micro sample between 1995 and 1998 - Marginal effects from probit models

	(1)	(2)	(3)
Subjective measures of job insecurity for fathers			
Father's perceived job insecurity	0.017		
	(0.010)**		
Father's income uncertainty (uniform distr.)		0.012	
		(0.007)**	
Father's income uncertainty (triangular distr.)			0.012
			(0.007)**
Objective measures of job insecurity for fathers:			
Change between t and t+1 in fraction of unemployed of age 30-64 in gender-province cell	0.303	0.364	0.363
	(0.163)**	(0.180)**	(0.179)**
Change between t and t+1 in fraction of temporary jobs of 30-64 year old in gender-province cell	0.036	0.042	0.045
	(0.105)	(0.122)	(0.122)
Objective measures of job insecurity for children			
Change between t and t+1 in fraction of unemployed of age 15-29 in gender-province cell	-0.033	-0.035	-0.034
	(0.046)	(0.051)	(0.051)
Change between t and t+1 in fraction of temporary jobs of 15-29 year old in gender-province cell	-0.063	-0.074	-0.074
	(0.052)*	(0.057)*	(0.057)*
Control variables			
Age	-0.0003	0.00002	0.00002
	(0.0005)	(0.0005)	(0.0005)
Female	0.006	0.007	0.007
	(0.005)	(0.005)	(0.005)
Child not employed	-0.048	-0.037	-0.037
	(0.020)**	(0.017)**	(0.017)**
Years of schooling	0.0002	0.0001	0.0001
	(0.0006)	(0.0007)	(0.0007)
Father's age	0.0004	0.0004	0.0004
	(0.0005)	(0.0005)	(0.0005)
Father's years of schooling	-0.0008	-0.0008	-0.0008
	(0.0007)	(0.0007)	(0.0007)
Father not employed	-0.003	-0.005	-0.005
	(0.003)	(0.004)	(0.004)
Home-ownership	0.007	0.008	0.008
	(0.005)*	(0.005)*	(0.005)*
Wealth	0.006	0.006	0.006
	(0.003)**	(0.003)**	(0.003)**
Number of children	-0.005	-0.006	-0.006
	(0.003)**	(0.003)**	(0.003)**
Mother employed	-0.002	-0.002	-0.002
	(0.003)	(0.004)	(0.004)
Home rental index in province	-0.002	-0.002	-0.002
	(0.002)	(0.002)	(0.002)
Controls for geographic area	yes	yes	yes
$Pseudo-R^2$	0.372	0.387	0.387

Note: Marginal effects from probit models that include a constant term. In each model the number of observations is 479 from 298 households. Standard errors (clustered by household) are reported in parentheses. Source: Italian Survey of Household Income and Wealth (SHIW). The statistical significance of the test that the underlying coefficient is zero is denoted by: $p < 0.05 = *, p < 0.01 = **$.