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Youth emancipation and perceived job insecurity of parents and children

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Abstract We test whether job insecurity of parents and children affect children's moving-out decisions. Macroeconomic estimates for 13 European countries over 1983–2004 show that coresidence increases by 1.7 percentage points (PP) following a 10 PP rise in the share of youths perceiving their job to be insecure and declines by 1.1 PP following the same increment in insecurity for older workers. Microeconometric evidence for Italy in the mid-1990s shows that the probability of moving out 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 · Moving out · Job security

JEL Classifications 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, it was as high as 73% in Italy and Finland, and above 60% in Greece, Spain, and Portugal. 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 moving-out 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 Organisation for Economic Co-operation and Development (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: high internal migration makes unemployment rate disparities across states to be scarcely persistent in the USA (Blanchard and Katz 1992), whereas they are very persistent across regions in low internal migration countries like Italy and Spain (Decressin and Fatas 1995; Bentolila and Jimeno 1998, respectively).

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, the number of births per woman of reproductive age has 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).

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

In sum, there are huge disparities in coresidence rates across countries, and they matter for welfare. The economic literature on moving-out decisions has focused mainly on parental and children's income and on housing prices (see Section 2). Here, we study empirically one factor that has not received much attention so far, namely, the degree of job insecurity perceived by children and their parents. In particular, we test two hypotheses derived from the theoretical literature, as follows: that children's job insecurity lowers the probability of moving out, whereas parental job insecurity raises it (Fernandes et al. 2008).

We also add to a recent strand of the literature by using subjective measures of job insecurity, in addition to more standard, objective ones.

Looking first at macro data, we show in Section 3 that, in a panel of 13 EU countries from 1983 to 2004, after controlling for a host of factors, higher youth insecurity and lower parental insecurity are associated with higher coresidence. The results also 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 4, we exploit the panel data structure of the Italian Survey of Household Income and Wealth (SHIW) collected by the Bank of Italy, which contains high-quality data on individual-specific perceived job insecurity for fathers. For children, we consider only objective measures, since the data are not sufficiently informative to construct subjective ones. 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 of them 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 5, we present our conclusions.

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. They find 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 parental income is, since then, the child shares from a larger pie, unless parents have a strong taste for privacy (so that they are willing to make higher transfers to independent children). The effect of an increase in housing prices is ambiguous (Ermisch 1999). 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. However, if parents' housing demand is elastic, then the decline in housing may lead the child to leave.

For Anglo-Saxon countries, empirical findings indicate that higher child earnings reduce coresidence, whereas child unemployment raises it (Rosenzweig and Wolpin 1993; for the USA; Ermisch 1999, for the UK). At the aggregate level, Card and Lemieux (2000) find that stronger local demand conditions and higher wages induce young men to move out. The father's earnings are found to raise coresidence in Italy, the UK, and the USA (Manacorda and Moretti 2006; Ermisch 1999; and McElroy 1985, respectively; although Rosenzweig and Wolpin 1993 find a deterring effect if parents are divorced). Housing prices have been found to deter emancipation in Italy, Spain, and the UK (Martinez-Granado and Ruiz-Castillo 2002; Giannelli and Monfardini 2003; and Ermisch 1999, respectively), while Alessie et al. (2006) link late emancipation in Italy to high transaction costs in housing.

Another potential determinant of moving-out decisions is job insecurity. Fogli (2004) presented a model featuring low parental job insecurity as a cause of late nest leaving. In Fernandes et al. (2008), we extend the standard framework to analyze the effect on coresidence of uncertain income streams. In our model, the moving-out decision is taken before observing the realizations of the child's and parental income. We assume that moving back home is costly, which gives rise to an option value associated with waiting to see the income realizations before deciding whether to leave. 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.[1] 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 due to a higher chance of regret, whereas the opposite is true for the variance of parental income.

In this paper, we study empirically the effect on moving-out decisions of the degree of job insecurity perceived by children and their parents. Our key variable of interest is the perceived probability of being unemployed, labeled p. Drawing from Fernandes et al. (2008), we can infer its effect on moving-out decisions. Suppose that workers are either employed and receiving a wage or unemployed and receiving unemployment benefits. 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. In this case, and as long as transfers to independent children are relatively unimportant, the theory predicts that an increase in the child's p should make moving out less likely, whereas the opposite is true of an increase in the parents' p. These are the two key hypotheses we aim at testing.

Of course, we cannot be sure that the assumptions needed to obtain unambiguous predictions in our theoretical model strictly hold in reality. For instance, in the model, we assume that parents are altruistic and children are selfish. Instead, altruistic children might remain with their parents in the face of increased parental job insecurity in order to financially assist them. The predicted effects of child and parental insecurity are, therefore, an empirical issue.

How do we measure job insecurity? In empirical work, expectations are often replaced by outcomes—under rational expectations, forecast errors are

[1]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.

purely random. Recently, however, a burgeoning literature has shown the usefulness of survey-based individual expectations (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 and Straub (2000) 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, of the question in the European Community Household Panel on how satisfied respondents are with their job in terms of job security (Deloffre and Rioux 2004; Clark and Postel-Vinay 2008). Those authors also stress the importance of qualitative vs quantitative replies. For instance, the question in the US General Society Survey on the probability of job loss allows 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 following US Survey of Economic Expectations 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 microeconomic evidence relies on quantitative answers to a very similar question, while our macroeconomic evidence is based on qualitative answers.

3 Macroeconomic evidence for the EU

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

3.1 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 are 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 are 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 1983–1984 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 5 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 the 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-year-olds.

The table shows that job insecurity is negatively correlated with age and that older females feel slightly more insecure than men.[2] Across countries, young workers feel most insecure in France, Spain, Greece, Italy, the UK, and Finland, while older workers feel the 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.

[2]For analyses of the determinants of the job loss probability, see Manski and Straub (2000) for the USA, Green et al. (2001) for the UK, and Böckerman (2004) for 15 European countries.

Table 1 Descriptive statistics on coresidence and perceived job insecurity in Europe, 1983 to 2004 (%)

	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)
UK	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)

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

The OECD (1997) claimed that there was a widespread increase in perceived job insecurity from the 1980s to the 1990s in OECD countries. Nickell et al. (2002) consider a wider concept of insecurity, by including the chances of wage losses as a result of unemployment or in continuing jobs, finding an increase for British men from 1982 to 1997. We cannot discuss long-term trends, given the limited sample of countries available for the 1980s. However, 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.

3.2 Results

3.2.1 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* 50–59):

$$\begin{aligned}\text{Coresidence}_{ijt} = \sum_{i=1}^{13} \beta_{0i}\text{Country}_i + \beta_1 \text{Youth insecurity}_{ijt} + \beta_2 \text{Insecurity 50–59}_{it} \\ + \beta_3 \text{Age 25–29} + \beta_4 \text{Female} + \beta_5 \log(\text{Real GDPpc}_{it}) + e_{ijt}\end{aligned}$$

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 25–29 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 leave their parents' home earlier than males, 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 years old, respectively, by 2003. 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. Moving out and living in a couple are closely linked, especially for women. In 2005, 55% of independent 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%. These facts match well with the increase in the moving-out age over time and with the difference in the median age of leaving between men and women, which was equal to 2.5 years in 2005.[3]

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 adjusted R^2 drops from 0.95 to 0.64.

[3]Finland, Ireland, and Luxembourg are excluded due to lack of data. Sources: Eurostat (2002), Table 1, section 1.3.13, for age at marriage in 1980; Eurostat (2008), Statistical Annex Tables A.6, A.10, and A.12 for age at marriage in 2003, living in a couple, and age of moving out, respectively.

Table 2 Macroeconomic evidence on coresidence and perceived job insecurity

	(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

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 1-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 (see Staiger and Stock 1997) are all below 0.01. The statistical significance of the test that the underlying coefficient is equal to zero is denoted by: $p < 0.05 = *$, $p < 0.01 = **$

Conceptually, they control for country-specific factors affecting coresidence. 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 USA 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 once the *sexual revolution* of the late 1960s reduced the privacy cost from coresiding. The remaining variables, which are time-varying, cannot be ascribed to cultural differences across countries and are meant to capture economic effects.

Column 1 of Table 2 reveals that job insecurity perceived by young workers 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.[4]

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 by 10 percentage points, the coresidence rate would increase by 1.7 percentage points, while the same change in the share of insecure older workers would reduce coresidence by 1.1 points.

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 as low as 0.53 for young workers and 0.58 for older ones, while, for the share of temporary jobs in dependent employment, the correlation coefficients are, respectively, 0.50 and 0.30. This leaves scope for perceived and objective measures to play separate roles. We 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 (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 worker groups. We expect coresidence to be jointly determined with the youth unemployment rate, and so the model is estimated by instrumental variables.[5] To control for the possibility that youth insecurity may also be endogenously determined, it is instrumented as well. Due to the lags involved, the sample size drops

[4]Recalling from Table 1 that 20–24-year-old women perceive on average the same insecurity as men, whereas the 25–29-year-olds feel more insecure, we may expect women to emancipate at the same time as men or later. However, insecurity is lower for women than for men in 44% of all age–country–year cells. Thus, we should not expect a strong difference in coresidence by gender stemming from insecurity.

[5]The instruments are the 1-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 footnote to Table 2).

to 108 observations. Youth unemployment shows the expected sign whereas older-worker unemployment shows an unexpected positive sign, but neither is significant. At the same time, the coefficient on youth insecurity doubles with respect to column 2, while that on older-worker insecurity jumps up even more, and both variables retain their significance. The income elasticity drops slightly.

3.2.2 Alternative specifications and discussion

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

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.

A few alternative specifications, not shown to save space, were also tried. In particular, we: (a) added the share of individuals in the gender–age cell who had received education or training during the previous 4 weeks (from the European Labor Force Survey), as a proxy for the share of individuals in full-time schooling; (b) interacted the insecurity variables with age and gender dummies, to capture potential differential effects of insecurity by demographic group; and (c) interacted the insecurity measures with a dummy variable for 1983–1984, to capture potential effects of the different wording of the insecurity survey questions vis-à-vis the later period. None of these variables was significant, nor did their inclusion qualitatively alter the preceding results.

In sum, the aggregate evidence for European countries we have uncovered indicates that, once persistent demographic and cross-country differences are controlled for, job insecurity of children and parents influences 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 4 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 one (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.

4 Microeconomic evidence for Italy

In this section, we present 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.

4.1 Data and sample design

We use a representative sample of Italian individuals of working age, between 18 and 35 years old, living with their parents. This sample has been extracted from the Italian 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.[6]

[6]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.

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 moving-out 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 asking individuals about the probability of having a job in the following year. This information 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 largely individuals who enjoy perfectly safe incomes. We do not do so because retired fathers are more likely to be at home all day, and this might affect the moving-out decisions of children for reasons different from the pure effect of job security. Moreover, 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.[7] Its characteristics are described in the first two columns of Table 3.

The second problem affecting the question on perceived insecurity is that, to limit the questionnaire length, not *all* households were asked this question but only those in which the (male) head was born in an odd year. Thus, our measure of 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 3, and moreover, we have an analogous measure for only 212 of these children. 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 2 years of data (see the last two columns of Table 3), this is not true for the 212 households for which the information is available for both fathers and children. The reason is that not

[7] 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.

Table 3 Descriptive statistics for the Italian micro sample, 1995[a]

Variable	Full sample Mean	Full sample Std. dev.	Restricted sample Mean	Restricted sample 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

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 fathers answered the question concerning perceived uncertainty. Wealth is in billions of Italian liras. Data source: Italian SHIW. The rental index is in thousand liras per square meter

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 job insecurity on coresidence, we have to rely on objective measures that will be described in the next section.[8] 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 3). Moreover, its timing structure is suitable for exploring the relationship between paternal job insecurity and the subsequent (rather than contemporaneous) decisions of children to leave home controlling for a large set of individual and family background characteristics.

4.2 The indicators of job insecurity and the outcome variable

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:[9]

> 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 nonrespondents are excluded. Those who expected to voluntarily retire or drop from the labor force are not included. The answers attest to the high degree of job security enjoyed by workers in Italy: the fourth decile is zero, the median is 30%, a 50% chance of unemployment is reached only in the eighth decile, and only 3% of individuals are certain of being unemployed in the year following the interview (though it is not clear if employed respondents reported only involuntary job losses or any change in employment status, including job mobility). The authors also compare this source, restricting the sample to those employed, with the US Survey of Economic Expectations.

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[8]The information on job insecurity is available only for 479 fathers and not for 571 (= 1,142/2), even though it was requested from those born in odd years, because some of them did not answer the question. The comparison of observables in Table 3 clearly shows, however, that nonresponses are randomly distributed in the data.

[9]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.

While in Italy, 59% of individuals report a zero chance of unemployment, in the USA, only 31% do so. The cumulated fraction of respondents for each probability of unemployment is systematically lower in the USA than in Italy up to a 10% probability (at the seventh decile), after which it becomes similar.

Table 4 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. 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. 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 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 degree of insecurity that they face.

Nevertheless, we also use objective (and arguably exogenous) measures of insecurity based on labor force information by age, gender, and province.

Table 4 The indicator of perceived job insecurity in the Italian micro sample, 1995

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	

Distribution of the indicator of job insecurity of fathers in the sample of 479 observations used in the econometric analysis (see Table 3). The indicator measures the probability assigned by the individual to the event that he does not work in the following year. Data source: Italian SHIW

These are also the only measures we can construct for children since, as noted in the previous section, the number of observations on perceived job insecurity of fathers *and* children is too small. For this purpose, we obtained from the Italian statistical office (Istat) information from the Quarterly Labor Force Statistics on the fractions of unemployed and of temporary employees within cells defined by age, gender, and provinces (105 administrative units, which correspond approximately to US counties). We were forced to construct our measures for the age brackets predefined by Istat, namely, 30–64 years old for fathers and 15–29 years old for children. For both brackets, we compute the change in the fraction of unemployed and temporary workers between 1995 and 1998 by gender and province. We expect that, for both fathers and children, 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 3. In our sample of children living with their parents in 1995, only 4% decided to leave home over the following 3 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. However, this is consistent with the aggregate Italian data: the coresidence rate for people aged 20–29—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.

4.3 Results

The first column of Table 5 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 3 years from the baseline.

Ideally, we would like to base our estimates on a comparison of children who are identical with respect to all 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.

Moving-out 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 these variables control for his 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).

Table 5 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	0.303	0.364	0.363
unemployed of age 30–64 in gender-province cell	(0.163)**	(0.180)**	(0.179)**
Change between t and t+1 in fraction of temporary	0.036	0.042	0.045
jobs of 30–64 year old in gender-province cell	(0.105)	(0.122)	(0.122)
Objective measures of job insecurity for children			
Change between t and t+1 in fraction of	−0.033	−0.035	−0.034
unemployed of age 15–29 in gender-province cell	(0.046)	(0.051)	(0.051)
Change between t and t+1 in fraction of temporary	−0.063	−0.074	−0.074
jobs of 15–29 year old in gender-province cell	(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 SHIW. The statistical significance of the test that the underlying coefficient is zero is denoted by: $p < 0.05 = *$, $p < 0.01 = **$

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. Local conditions are further controlled for by the inclusion of five geographical area dummies. As far as children are concerned, we control for age, gender, schooling, and employment status.

The estimate in the first row and first column of Table 5 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 sure of being 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%.[10]

Objective measures of paternal insecurity, in the second and third rows of column 1 in Table 5, 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 leaving by one-third of a percentage point. This effect is also statistically significant and it should be evaluated with respect to the average probability of leaving 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 expected given that temporary jobs are relatively infrequent at old ages, whereas local unemployment is a better measure of objective job insecurity for fathers.

Estimates of the effects of objective measures of job insecurity for children are reported in the fourth and fifth rows of column 1 in Table 5. In this case, both indicators of insecurity have a negative effect on moving out, 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 moving out by two thirds of a percentage point.[11]

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 (see the descriptive statistics on the SDs of the insecurity measures in Table 3). 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 ($0.48 = 0.28 \times 0.017 \times 100$)

[10] We also used Rubin's (1987, 1996) multiple imputation method to impute perceived job insecurity to fathers for whom this information is missing. The coefficients were unchanged but efficiency was remarkably higher. The results are available from the authors upon request.

[11] In contrast, García-Ferreira and Villanueva (2007), who use legal changes in firing-cost regulations on fixed-term contracts in Spain to estimate the effect of employment risk on new household formation, do not find significant effects. However, Maeso and Mendez (2008) do find significant effects of job tenure and labor contract type—as objective measures for job insecurity—on the moving-out decisions of Spanish university graduates, while the significance of perceived insecurity measures depends on the estimation method.

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 ($0.48 = 0.016 \times 0.30 \times 100$). 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 ($0.31 = 0.051 \times 0.06 \times 100$).

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. 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 nest leaving 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 leaving 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 moving-out decisions. Higher family wealth and home ownership increase the probability of leaving, while higher rental prices in the province seem to discourage it.

In columns 2 and 3 of Table 5, we replace perceived insecurity of fathers by a measure of uncertainty regarding future paternal income. All 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. We can say something on this conjecture because the SHIW allows us to approximate the expected earnings 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 SD to the expected value, for each individual in their sample. We use these computations to estimate a probit model like the one of 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 5, 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.

5 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 EU 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. 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 SHIW, 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 estimated once changes in the local unemployment and temporary employment rates are controlled for.

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, which we have used to analyze 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 have employed to examine the determinants of changes in coresidence. The macro evidence is more descriptive, while the micro data allow us to control for a more exhaustive set of variables potentially determining moving-out decisions. While the two sets of results are therefore not directly comparable, we believe that the consistency of the qualitative results obtained from both analyses provides robust evidence regarding the validity of the hypotheses of interest.

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 leave the parental home later.

The main direct effects of late nest leaving 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 moving-out decisions should not be disregarded.

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Appendix: Description of macroeconomic data

Coresidence rate. Fraction of population living in the 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 UK (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, UK, 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 5 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, 2007, rescaled to 1996 as base year.

Real house prices. House price index deflated by consumer price index. Available for Belgium, Germany, Spain, France, Ireland, Italy, Netherlands, UK, and Finland. Years: 1983–2004. Source: Bank for International Settlements Data Bank.

Unemployment rate. Source: OECD, 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 6 presents descriptive statistics of the variables used in Tables 1 and 2.

Table 6 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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