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Job Security and Job Protection

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Abstract

We construct indicators of the perception of job security for various types of jobs in 12 European countries using individual data from the European Community Household Panel (ECHP). We then consider the relation between reported job security and OECD summary measures of Employment Protection Legislation (EPL) strictness on one hand, and Unemployment Insurance Benefit (UIB) generosity on the other. We find that, after controlling for selection into job types, workers feel most secure in permanent public sector jobs, least secure in temporary jobs, with permanent private sector jobs occupying an intermediate position. We also find that perceived job security in both permanent private and temporary jobs is positively correlated with UIB generosity, while the relationship with EPL strictness is negative: workers feel less secure in countries where jobs are more protected. These correlations are absent for permanent public jobs, suggesting that such jobs are perceived to be by and large insulated from labor market fluctuations.

Keywords: Perceived Job Security, Employment Protection Legislation, Unemployment Insurance Benefits.

JEL codes: J28, J65, I31.

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1 Introduction

The most common policies used to protect workers against labor market risk are Employment Protection Legislation (EPL) and Unemployment Insurance Benefits (UIB). The effect of EPL on indicators of labor market performance is an arguably rare example of agreement among economists. Greater employment protection is thought to discourage both firing and hiring, with an overall ambiguous effect on the unemployment rate. The main effect of EPL is to reduce the permeability of the barrier between work and unemployment. This conclusion, which results from most recent equilibrium labor market models, is largely borne out by empirical research.[1] UIB generosity, on the other hand, is less clearly related to aggregate labor market flows (or aggregate labor market indicators in general). As such, UIB are generally thought of as being more compatible with the demand for labor market flexibility.[2]

While there is apparent agreement on the macroeconomic impact of EPL and UIB, only very few studies have asked how these institutions affect workers'perceptions of their job security. Yet the question would seem to be of obvious importance, as it seems likely that policy makers are responsive to the expression of public sentiments of "social insecurity". The primary aim of this paper is to address that question.

The balance between labor market flexibility and security is a live policy issue. One of the campaign posters of Arlette Laguiller, the candidate for one of the far-left wing parties calling itself *Lutte Ouvrière* (literally: "Workers' Struggle"), in the 2002 French presidential election stated: "*Arlette Laguiller—Always on the workers' side—Ban layoffs*". The baseline argument behind the proposal to "ban layoffs" is that all-powerful shareholders use labor force adjustments to maximize their profits, and in so doing they let workers bear all the financial risk, thus creating social insecurity. Judging by the 2002 election results,[3] the idea to make layoffs illegal sounded

[1]See Addison and Texeira (2003), OECD (1999) or the excellent survey in Cahuc and Zylberberg (2004). Further, as it affects private decisions about job creation and destruction, EPL can obviously be thought of as serving a more general purpose than just to protect workers against layoff risks. See Blanchard and Tirole (2004).

[2]While raising other standard incentive-related problems. Here also, we refer the reader to the corresponding chapter in Cahuc and Zylberberg (2004).

[3]Arlette Laguiller received 5.72% of the votes. Yet the platfoms of at least three other left wing parties—the *Ligue Communiste Révolutionnaire* (Revolutionary Communist League), the *Parti Communiste Français* (French Communist Party) and the *Parti des Travailleurs* (Workers' Party)—also contained similar proposals concerning

appealing to a nontrivial fraction of French voters.

The more "official" view of the European Union on flexibility and security is somewhat different. The 2003 *Employment Guidelines for Member States*[4] recommend that "*Member States will facilitate the adaptability of workers and firms to change, taking account of the need for both flexibility and security* [...]. *Member States will review and, where appropriate, reform overly restrictive elements in employment legislation that affect labour market dynamics* [...]." While social insecurity is definitely a matter of concern in many official EU documents,[5] the current trend in addressing social insecurity seems to be toward institutions that are more friendly to labor market dynamics. In short, less EPL and, to an extent deemed reasonable, more UIB.

The extent to which the reforms actually implemented conform with those broad recommendations varies across Member States. While the Dutch 1999 "Flexibility and Security Act" or the Danish agenda on "flexicurity" are clearly in line with the EU view, other (mostly Southern) countries are more hesitant. In fact, many authors have noticed that standard indicators of EPL strictness and UIB generosity are negatively correlated across European countries.

The origin of this apparent trade-off is a subject of active theoretical research. Saint-Paul (2000, 2002) analyzes political economy models of the choice of labor market institutions, in which EPL and UIB are treated separately. Boeri, Conde-Ruiz and Galasso (2003) offer a thorough theoretical exploration of the EPL-UIB trade-off, which they view as different realizations of stable politico-economic equilibria. One recurring point in this literature is that EPL is essentially championed by insiders—those who already have a job—who protect their associated rent, whereas UIB mostly favor outsiders. Labor market institutions then have a feedback effect on this conflict of interests, both because they impact on the composition of the labor force, and also because they directly

dismissals. In total (including Arlette Laguiller), these parties together received 13.71% of the votes.

[4]Published in the *Official Journal of the European Union*, and available online (English version) at http://europa.eu.int/comm/employment_social/employment_strategy/guidelines_en.htm. See also the UK Department for Work and Pensions' Report "Full employment in every region", Paragraph 2.12: "It is inevitable and desirable that people will change jobs during their working lives".

[5]One of the two parts of Priority 7 of the European Union's 6th Research Programme specifically mentions labor market insecurity. One of the four parts of the European Working Conditions Observatory's definition of quality of work is "ensuring career and employment security". Some of the recent projects funded by the European Union are entitled "Employment Precarity, Unemployment and Social Exclusion", "Social Exclusion and Social Protection—the future role for the EU", and "Precarious employment in Europe: a Comparative Study of Labour Market Related Risks in Flexible Economies".

affect insiders' rents.

While we do not claim to provide a complete empirical counterpart of that theoretical literature, in this paper we consider what comparative large scale survey data can teach us about the relation between workers' perceived job security and labor market institutions. We use data from the European Community Household Panel (ECHP) to explicitly address the issue of job security, as reported by workers in twelve European countries. We then consider the relation between this reported job security and standard OECD summary measures of EPL strictness on the one hand, and UIB generosity on the other. Our use of panel data allows us to explicitly model worker selection into four different types of labor market status (permanent private jobs, permanent public jobs, temporary jobs, and nonemployment). We are thus able to capture the endogeneity of job choice, and the relation between this allocation process and individual perceptions of job security.

We have five main findings. First, after controlling for selection, workers feel most secure in permanent public jobs, and least secure in temporary jobs. While this may sound intuitive, it nonetheless provides some support for the idea that subjective measures of job security predict "objective" job security. Put differently, workers' assessments about their job security are "correct".

Second, job security in permanent private and temporary jobs is positively correlated with UIB generosity across countries.

Third, in permanent private and temporary jobs, workers in countries with higher EPL actually feel *less* secure. While care needs to be taken in establishing the causality of these correlations, this third result suggests that job protection is not the best response to the problem—real or supposed (see OECD, 1997)—of job insecurity.

Fourth, public sector jobs are largely considered to be the most secure, and the correlation of job security with UIB or EPL in this sector is essentially zero. Public sector jobs are by and large perceived to be insulated from labor market fluctuations.

Fifth, the gap in perceived job security between public sector and other job types, which can be interpreted as the "job security returns to being an insider", increases with EPL strictness and falls with UIB generosity. This squares in well with the basic message of the political economy literature briefly mentioned above.

The paper is organized as follows. Section 2 discusses some issues relating to the measurement of job security perceptions using subjective data, and briefly describes the ECHP data that we use. Section 3 contains a brief review of the related empirical literature, and makes a first pass at examining the correlations between job security and labor market policy indicators. Section 4 presents our main statistical model, discusses endogeneity issues and explains the estimation protocol. In section 5 we present the estimation results in two parts: first we analyze the individual determinants of job insecurity and the selection of workers into various employment states; and, second, we focus specifically on the relationship between labor market institutions and job security. The interpretation of the results is discussed in Section 6. Finally, Section 7 offers some concluding remarks.

2 Measuring perceptions of job security

In this first section we argue that our understanding of the way in which individuals are affected by labor market institutions such as employment protection or unemployment insurance can be enhanced by subjective data on job security, which appear in a number of different large-scale surveys. We first discuss the forms that these questions most commonly take, then we describe the particular data used in this paper.

2.1 The wording of job security questions

Survey questions on job security typically appear in two broad forms. Most commonly, individuals are asked to report their degree of satisfaction with respect to their job security. A typical "satisfaction" formulation would be: "*How satisfied are you with your present job or business in terms of job security?*" followed by a verbal scale such as "Very satisfied, somewhat satisfied, ..." and so on. This formulation renders the interpretation of the resulting measure of job (in)security somewhat problematic. First, it contains an important subjective element (the meaning of "satisfied" or even "job security" may vary from one person to another). As such, it is not immediately obvious that they can be usefully compared across individuals or countries. Second, it confounds the respondent's perception of at least two very different components of job security, namely the *probability* of job loss and the *cost* of job loss.

An alternative to the above "satisfaction" formulation is the use of a "probabilistic" question, i.e. to ask individuals about the probability of losing their job. Here a typical wording would be:[6] "*What do you think is the percent chance that you will lose your job during the next 12 months?*". Probabilistic questions are more immune to the "confounded issues" criticism. As such, their use is advocated in a number of recent contributions (Dominitz and Manski, 1996; Manski and Straub, 1999).

Our primary objective in this paper is to explore the relationship between perceived job security and labor market institutions; for the latter we specifically consider Employment Protection Legislation (EPL) or Unemployment Insurance Benefits (UIB). Since these institutions are typically defined at the national level (and are measured by indicators showing little if any time-series variation), we obviously need a multi-country data set. In the following we thus use a subsample of data from the European Community Household Panel survey (ECHP), which is a panel of individual data gathered by EUROSTAT covering fifteen EU countries in total. One decisive advantage of the ECHP data is that there is ex ante harmonization of the questionnaire between countries. Apart from the traditional variables found in national household surveys (demographic characteristics, income, health, housing, and so on), the ECHP contains a number of "sociological" questions regarding personal relationships and outside work activities, as well as a number of satisfaction questions. Included in these latter is a question on satisfaction with job security. The exact wording is as follows:[7]

> **Question:** *"How satisfied are you with your present job or business in terms of job security? Using the scale 1 to 6, please indicate your degree of satisfaction. Position 1 means that you are not satisfied at all, and 6 that you are fully satisfied."*

[6]As used in the US *Survey of Economic Expectations* (Dominitz and Manski,1996; Manski and Straub, 1999) and in the US *General Social Survey* (Blanchflower and Oswald, 1999).

[7]The ECHP User Data Base manual only provides the wording in English. At this point we have no way of assessing possible differences in the question's formulation across countries resulting from translation.

Clearly, this is not a probabilistic question and is therefore exposed to both the "interpersonal comparability" and the "confounded issues" criticisms discussed above. We shall try to (at least partially) deal with the former by allowing for unobserved individual heterogeneity in our statistical analysis below. It is important to bear the latter in mind when interpreting the results. Specifically, our general approach will be to interpret the replies to the question above as proxy measures of *the workers' subjective assessment of the expected change in utility associated with job loss multiplied by the (subjective) probability of that loss.*[8] This interpretation has the advantage of explicitly acknowledging the "confounded issues" issue in a natural and convenient way. The use of a pure probabilistic question would be a useful complement to the current analysis; unfortunately, we know of no multi-country panel including this information.

2.2 (Brief) sample description

Returning to our sample, due to missing data we are only able to use twelve of the fifteen countries and the last five (out of eight) ECHP waves. Moreover, for reasons that we shall briefly discuss below, we focus on men. Our final sample consists of male workers aged between 20 and 55 in 1997, who are observed to be either wage earners or nonemployed at every annual interview between 1997 and 2001: we have information on 12,091 individuals $\times$ 5 waves; the country distribution of observations is described in data Appendix A.

Obviously, the above job satisfaction question was only asked of currently-employed individuals. Figure 1 shows per-country histogram plots of the distribution of replies to the job security question (among employed wage earners).

Concerning those distributions, note first that, as is often the case with such satisfaction scales, the responses at the bottom of the distribution (1 and 2) were given only infrequently. This is

[8] Formally, consider some employed worker i answering the job security question at date t. Denote worker i's expected lifetime utility from job continuation at date t by V_{it}^{E}, the same worker's expected lifetime utility from being dismissed at date t by V_{it}^{U}, and finally denote worker i's perceived probability of job loss at date t (say, within the year following the interview) by q_{it}. Then our proposed interpretation of worker i's response to the job insecurity question is as a measure of $q_{it} \cdot \left(V_{it}^{E} - V_{it}^{U}\right)$.

a standard and well-documented feature of job satisfaction data. Second, it clearly appears that individuals' feelings about job security differ from country to country. It will be our purpose for the rest of this paper to describe those differences.

3 Job security and job protection: a first pass

In this section we highlight simple correlations between reported job security and a number of individual characteristics and indicators of labor market institutions. In particular, it seems natural to ask whether EPL on one hand, and UIB on the other, play a role in attenuating feelings of insecurity. We begin with basic bivariate correlations between job security and job protection; we then investigate the role of observed individual heterogeneity.

3.1 A raw measure of job security

A first naive indicator of job security in each country can be constructed from the country-level mean response to the question described above.[9] In this paragraph we ask the following two simple questions. First, how do our 12 countries compare in terms of this raw indicator? Second, is the level of job insecurity revealed by these indices correlated with the labor market institutions in the different countries in our sample?

The answers to both of these questions are contained in Figures 2 and 3. Figure 2 plots the 1998 OECD indicator of employment protection (x-axis) against our measure of job security (y-axis); Figure 3 repeats the exercise with the 2000 OECD index of Unemployment Insurance generosity.[10]

[9] An alternative would be to use the per-country percentage of replies above a certain cutoff (e.g. the percentage of respondents saying they are either "satisfied" or "very satisfied".) Such alternative indicators lead to the same qualitative conclusions as the country-level mean that we are using in this paragraph.

[10] The OECD has various indicators of UIB generosity. The one that we are using takes the form of an average net replacement rate combining a variety of typical individual cases. One important drawback of this indicator is its failure to take into account the criteria governing eligibility for Unemployment Insurance. Since these criteria vary widely across countries, this is potentially problematic. Alternatively, we could use average UI expenditures per unemployed person, which are arguably more complete measures of UIB generosity. We choose average replacement rates because there is a mechanical negative correlation between mean UI expenditures per unemployed and the unemployment rate, which is in turn likely to be negatively correlated with job security for reasons that we discuss below. This mechanical correlation may thus cause an artificial positive correlation between mean UI expenditures and job security. Nevertheless, the following analysis can be carried out using either measure. The results obtained

First, looking at the vertical scales on both Figures, we obtain a job security ranking of the 12 countries in our sample. The basic picture is of workers in "Southern" countries (Portugal, Italy, France, Spain and Greece) feeling less secure than their counterparts in Northern countries (the Netherlands, Denmark and Ireland, with the most secure country being Austria). Workers in the United Kingdom also occupy a fairly low position in this ranking.

Second, Figure 2 strongly suggests a negative correlation between job security and job protection: at first blush, workers in countries with stricter EPL feel *less* secure in their jobs. Conversely, Figure 3 suggests—somewhat less strongly[11]—that workers in countries with more generous UIB feel *more* secure in their jobs.

Those two bivariate conclusions are however *a priori* fragile, as job security doubtless depends on any number of (observed or unobserved) individual, job or labor market characteristics. Such differences are unlikely to be orthogonal to the degree of EPL or to UIB generosity: for example, it is well-known that countries with stricter EPL have a greater proportion of temporary jobs. Holders of such jobs likely feel less secure (so that, across countries, EPL and insecurity are positively correlated), but they are not necessarily insecure because of the stricter EPL. Moreover, it is well-known that UIB generosity and EPL strictness are negatively correlated across European countries.

In the following subsection we partly deal with this objection by controlling for a set of observed individual and labor market characteristics. The role of unobserved heterogeneity will be explored in section 4.

3.2 Observed individual heterogeneity

Personal characteristics, labor market conditions and job type. Our first step beyond simple bivariate analysis is to regress reported job security on a variety of controls, including the OECD indicators of EPL strictness and UIB generosity. Specifically, we consider the following

using average UI expenditures per unemployed (which are available upon request) are qualitatively similar to those presented here.

[11]The correlation is positive, but not statistically significant in a cross-country regression (whereas the slope of the EPL-security relationship is statistically significant). Concerning Figure 3, the corresponding scatterplot using average UI expenditures per unemployed as an indicator of UIB generosity is much more impressive. Yet this may be artificial to some extent—see the preceding footnote.

personal characteristics: age and age-squared, education (3 dummies[12]), marital status, the presence of children under 15 in the household, an indicator of foreign citizenship, and an indicator of the existence of a long-term unemployment spell in the recent past.[13] We also include the 5-year average local unemployment rate as an indicator of local labor market conditions.[14]

Finally, we want to allow for the possibility that holders of different job types have fundamentally different perceptions of their job security. To this end we distinguish 3 different job types:[15]

- $e =$ ppriv: employed under a permanent contract in the private sector;
- $e =$ ppub: employed under a permanent contract in the public sector;
- $e =$ temp: employed under a temporary contract.

The observed distribution of individuals across job types and the state-to-state transition matrix are shown in Appendix A.

Results. Table 1 displays the results from Ordered Probit regressions of perceived job security on the set of controls that we just described, separately for the three job types defined in the previous paragraph. The sample is the initial wave (1997) of our panel.[16]

Table 1 first reports a constant, which is normalized to zero for permanent private sector employees.[17] The constant is not statistically significant for public employees, suggesting that

[12] Third level education (ISCED 5-7), Second stage of secondary level education (ISCED 3) and less than second stage of secondary level education (ISCED 0-2). Those dummies are based on the ECHP variable indicating the "highest level of general or higher education completed" (PT022). The quality and cross-country comparability of this variable is questionable, but this is the only general education variable available in the ECHP user database.

[13] In practice we use an indicator of whether the individual has had an unemployment spell of over one year in the five years prior to 1997 (the first year in our observation window).

[14] The "local" unemployment rate is constructed using the ECHP data as the proportion of those active in the labor market who are unemployed (ILO definition) at the NUTS1 regional level.

[15] One of the main reasons why we focus on males is to limit the number of job states. As expected, a significant fraction (around 22%) of the female workers present in our intial sample work in part-time jobs (while the corresponding male share is less than 3%). Since part-time jobs have notoriously different "stability" characteristics than full-time jobs, they should count as distinct job types. Taking them into account would have led us to double the number of job states, which at this point is computationally very costly.

[16] Results from subsequent waves are very similar.

[17] Any *ceteris paribus* differences in job security are subsumed into this constant. That is, we have imposed equality of the ancillary cutoff parameters of the Ordered Probit across job types. While admittedly restrictive, this method facilitates the comparison of job security between job types.

there is no systematic *ceteris paribus* difference in perceived job security between employees of the private and public sectors, at least among permanent job holders. The point estimate for temporary workers is negative, larger in absolute value and of borderline statistical significance. Hence there is some mitigated evidence that temporary workers feel less secure than holders of permanent jobs.

The estimated cohort effects for all job types suggest that job security is decreasing and convex (U-shaped) in age, as is often found in the analysis of subjective well-being measures (Clark, Oswald and Warr, 1996). We note that, even though the magnitude of the point estimates are roughly similar across job types, the age variables are statistically insignificant for public sector employees.

Lower-educated workers feel somewhat less secure than their higher-educated counterparts in all permanent job types, whereas education plays no role in terms of job security for temporary workers. This may be taken to reflect the generally less favorable conditions of low-skilled labor markets. Yet these particular results should be taken with some caution, given the arguably poor quality of the education variable (see footnote 12).

There is evidence that foreign-born workers feel more insecure than natives in permanent private jobs, while foreign citizenship is uncorrelated with job security among public or temporary employees. Neither cohabitation nor the presence of children in the household affect job security in any systematic way. Conversely, previous experience of long-term unemployment reduces perceived job security in all types of jobs (although with borderline significance for permanent private sector employees).

The next result is more striking: as one would expect, the average local unemployment rate sharply reduces perceived job security in temporary jobs, which are most exposed to the risk of layoff. The corresponding effect on perceived job security in permanent jobs (public or private) is *positive* (albeit insignificant for private jobs). A possible interpretation here is that workers in a depressed labor market tend to aspire to more "insulated" jobs, which is what permanent jobs, especially in the public sector, are perceived to be.

Finally, the last two rows in Table 1 report the estimated effects of our policy indices. These pertain to the main objective of this paper, which is to explore the link between job security and labor market policy. The effect of UIB generosity on perceived job security is positive for workers in all job types. Moreover, all three point estimates are roughly similar in magnitude (slightly larger for temporary workers). This corroborates the visual impression given by Figure 2. However, regarding EPL strictness, we find a negative correlation with perceived job security for permanent private employees (as Figure 1 suggests), essentially zero for temporary job holders, and positive for permanent public sector employees. This mitigates the conclusion drawn from the negative bivariate correlation shown in Figure 1, which only survives in multivariate analysis for permanent, private sector employees.

Discussion. While some of the regression results gathered in Table 1 are either standard or intuitive, others are puzzling. First, temporary work being the only job type for which education is uncorrelated with job security is difficult to explain. With temporary workers being objectively more exposed to the risk of losing their jobs, and with the chances of finding a replacement job being arguably lower for less-educated workers, one would have expected, if anything, the correlation between education and job security to be stronger among temporary than permanent workers. Second, the ranking of job types in terms of the strength of the correlation between job security and EPL strictness is also somewhat surprising. As temporary jobs are more exposed to the risk of job loss, we might have expected temporary workers to be most responsive (one way or the other) to differences in job protection regulations. Last, there is only a very small *ceteris paribus* job security difference between job types: the constant term in Table 1 is only negative and weakly significant for temporary job holders. Given the large differences across job types in "objective" job security (as measured, for instance, by the year-to-year probability of becoming non-employed—see Table A2 in the Appendix A), taking this result at face value would suggest that workers are biased in the assessments they make of their own job security.

While there are likely many potential explanations for those results, we explore in this paper the role of unobserved individual heterogeneity. Our prior is that the selection of workers into job types may not be independent of the workers' (unobserved) general attitude toward job security. For example, workers who are "worried"—an unobserved psychological trait that partly determines the sentiment of job security—may strive to select themselves into objectively "safer" jobs, e.g. permanent public jobs. Conversely, one could imagine that "low self-confidence"—another potential unobserved determinant of job security— is a handicap in job search. Consequently, workers with this trait may end up in "undesirable" job states—typically, temporary jobs or nonemployment—more often than their more self-confident counterparts. In either case, the results presented in Table 1 will partly pick up selection effects. Finally, because the job security question was only asked to workers holding a job at the time of the interview, the sample on which Table 1 estimates are based is itself a selected sample. It excludes all the nonemployed, a population likely to have different attitudes toward job security.

Our task in the remainder of this paper will be to exploit the longitudinal dimension of the data to try to account for all those selection problems. Before we proceed, however, we conclude this section with a quick review of the related results obtained by earlier contributions.

3.3 Related literature

This paper is not the first to consider the relationship between subjective measures of job security on the one hand, and institutional features of the labor market on the other. The bivariate analysis in OECD (1997) reveals no correlation between insecurity and EPL, but a negative correlation between insecurity and the replacement rate. More recent analysis has pointed to a seemingly aberrant positive bivariate relationship between job insecurity and EPL strictness. Böckerman (2004) uses data from 16 European countries in the 1998 "Employment Options for the Future" survey, and reveals a positive correlation between job insecurity and EPL, and a negative correlation with the replacement rate. Postel-Vinay and Saint-Martin (2004) find similar correlations using three different job security questions from 2 different data sources—wave 6 of the ECHP and the 1997 "Work Orientations II" wave of the International Social Survey Programme (ISSP). A recent paper with a somewhat different aim is Deloffre and Rioux (2003), who use data on eleven countries from one wave of the ECHP (1999) to examine the role of (endogenously chosen) contract type, and to assess whether employees' evaluations of their job security are "correct". Finally, Boeri et al.(2001) analyze unique, one-time survey data in which 5,500 citizens from France, Germany, Italy and Spain were asked (*inter alia*) a series of questions about the extent to which they would be willing to pay for more generous unemployment insurance. One of the onclusions reached by these authors is that proposals to increase UIB generosity find more support in countries offering both less generous UIB and more stringent EPL. More generally, Di Tella *et al.* (2003) show that life satisfaction is positively correlated with the replacement rate.

Our contribution in this paper is to add to this literature by distinguishing between job security in different kinds of jobs, and by taking seriously the possibility of endogenous choice of sector by workers.

4 A statistical model of job security

In this section we present the complete statistical framework that we use to analyze job security. Here we take a two-step approach, similar in spirit to (e.g.) Eckstein and Wolpin (1999). In a first step we decompose job security (as measured by the replies to the aforementioned question) into a component capturing the effects of time-varying local and aggregate labor market conditions and another component measuring each individual's "long-run" perception of job security given job characteristics. In a second step, we propose a statistical decomposition of these estimated individual measures of job security into permanent, observed and unobserved individual/job/labor market characteristics, with special interest in the role of country-level EPL and UIB indicators in explaining individual perceptions of job security.

4.1 Step 1: A decomposition of job security

The job security equation. Let s^*_{it} designate perceived job security for individual i at date t. We first decompose s^*_{it} as:

$$s^*_{it} = x'_{it}\beta + \sum_{e\in\mathcal{E}} \varphi^e_i \mathbf{1}_{e_{it}=e} + u_{it}, \tag{1}$$

where the notation is the following. First, x_{it} includes year dummies and date t local labor market conditions.[18] Second, e_{it} is the individual's job type—or job state—at date t. We consistently use

[18] As summarized by the local unemployment rate at date t (see subsection 3.2 above for a definition), taken as a deviation from its mean value over the five-year observation period. We shall return later to the reason why we use deviations from mean values here.

the notation $\mathbf{1}_{e_{it}=e} = 1$ if individual i is in a job type e at date t. We distinguish the three job types described in subsection 3.2, plus a fourth corresponding to non-employment (denoted $e = \text{none}$).

In terms of equation (1)'s notation, $\mathcal{E} = \{\text{ppriv}, \text{ppub}, \text{temp}\}$ is the set of job states in which the idea of job security is meaningful—that is, all states bar non-employment. We thus allow for individual differences in the effect of job statuses on perceived job security: individual random effects φ_i^e capture the influence on perceived job security *in a particular job type* e of all time-invariant, observed and unobserved individual heterogeneity variables. Our step 2 below will conduct a detailed exploration of the determinants of these random effects. For now, we let $\varphi_i = \left(\varphi_i^{\text{ppriv}}, \varphi_i^{\text{ppub}}, \varphi_i^{\text{temp}}\right)'$ denote the vector of labor market status/individual effects. Finally, $u_{it} \sim \mathcal{N}(0,1)$ is an i.i.d. error term, independent of the regressors and the individual effects.

Equation (1) thus decomposes s_{it}^* into a first component $x_{it}'\beta$ which captures the effects of temporary variations in local and aggregate labor market conditions, plus a second component $\varphi_i^{e_{it}}$ that captures the "permanent" (individual) impact on perceived job security of holding a particular job type e_{it}. Implicit in equation (1) is the assumption that the former component is "objective"—i.e. common to all workers—, while the latter is "subjective", so that the φ_i^e's are individual-specific.

The job security equation (1) implies that the conditional distribution of *reported* job security s_{it} given the explanatory variables $(x_{it}, e_{it}, \varphi_i)$ is the standard Ordered Probit. Defining the thresholds $-\infty = \tau_0 < \tau_1 < \ldots < \tau_6 = +\infty$ such that $s_{it} = h \Leftrightarrow \tau_{h-1} \leq s_{it}^* < \tau_h$, we obtain:

$$\Pr\left(s_{it} = h | x_{it}, e_{it}, \varphi_i; \Theta\right) = \left| \begin{array}{ll} N\left(\tau_h - x_{it}'\beta - \sum\limits_{e\in\mathcal{E}} \varphi_i^e \mathbf{1}_{e_{it}=e}\right) - N\left(\tau_{h-1} - x_{it}'\beta - \sum\limits_{e\in\mathcal{E}} \varphi_i^e \mathbf{1}_{e_{it}=e}\right) & \text{if } e_{it} \in \mathcal{E}, \\ 1 & \text{if } e_{it} = \text{none}. \end{array} \right. \tag{2}$$

where Θ denotes the entire set of parameters and $N(\cdot)$ denotes the c.d.f. of the standard normal distribution. The only subtlety here is that s_{it} is only observed when the individual is employed, i.e. if $e_{it} \in \mathcal{E}$.

The selection of workers into employment states. Our interpretation of the individual effect φ_i in equations (1) and (2) is that it partly captures psychological traits reflecting the taste or aversion for specific employment states. For instance, individuals with lower values of the φ_i^{temp} component of φ_i particularly dislike or fear the idea of being employed under temporary contracts and are thus likely to try and select themselves away from temporary jobs. We thus face a potential endogeneity problem in that φ_i is likely correlated with the observed employment state e_{it}.

The strategy we adopt is to model state-to-state transitions by a simple first-order Markov process in which the transition probabilities from an initial employment state ℓ are individual-specific:

$$\Pr\left(e_{it} = j | e_{it-1} = \ell, \mathcal{M}_i; \Theta\right) = \mathcal{M}_i\left(j, \ell\right), \tag{3}$$

where $\mathcal{M}_i\left(j, \ell\right)$ is the (j, ℓ) element of individual i's 4×4 (unobserved) transition matrix $\mathcal{M}_i$. The matrix $\mathcal{M}_i$ itself is treated as another individual random effect, which we shall allow to be correlated with φ_i in order to capture the potential selection of specific worker types into specific job types as evoked above.[19]

Individual likelihood contributions. It may clarify matters at this point to write down individual i's contribution to the sample likelihood. A typical observation is a set:

$$X_i = \left\{\overrightarrow{s}_i, \overrightarrow{x}_i, \overrightarrow{e}_i\right\},$$

where $\overrightarrow{s}_i = (s_{i1}, \ldots, s_{iT})$, $\overrightarrow{x}_i = (x_{i1}, \ldots, x_{iT})$ and $\overrightarrow{e}_i = (e_{i1}, \ldots, e_{iT})$ (individuals are observed for $T = 5$ periods).

Appending the missing unobserved heterogeneity variables to the observed data X_i, we obtain the complete data $\{X_i, \varphi_i, \mathcal{M}_i\}$. The contribution of individual i to the complete likelihood is a function of the parameters Θ and the complete data, $\mathcal{L}_i^c\left(\Theta; X_i, \varphi_i, \mathcal{M}_i\right)$, which can be factored as

[19] Having specified the process governing individual state-to-state transitions, we are left with the usual initial conditions problem, i.e. we have to model the marginal distribution of individual i's initial state, e_{i1}. This distribution depends on the same heterogeneity parameter as the transition process: $\Pr\left(e_{i1} = j | \mathcal{M}_i; \Theta\right) = \pi^1_{\mathcal{M}_i}\left(j\right)$. Another possible approach (advocated by Wooldridge, 2002), is to condition the whole problem on the individual's initial state, e_{i1}.

follows:

$$\mathcal{L}_i^c(\Theta; X_i, \varphi_i, \mathcal{M}_i) = \overbrace{\Pr(\overrightarrow{s}_i | \overrightarrow{x}_i, \overrightarrow{e}_i, \varphi_i; \Theta)}^{1}$$
$$\times \underbrace{[\Pr(\overrightarrow{e}_i | e_{i1}, \mathcal{M}_i; \Theta) \times \Pr(e_{i1} | \mathcal{M}_i; \Theta)]}_{2} \times \underbrace{\Pr(\varphi_i, \mathcal{M}_i | \Theta)}_{3} \times \underbrace{\Pr(\overrightarrow{x}_i)}_{4}. \quad (4)$$

The right hand side of (4) is a product of four terms. The first is the joint conditional probability of the sequence of reported job security values. This can be derived from the job security equation (1), as in equation (2). The second term is the joint conditional probability of the sequence of employment states in which individual i is observed. This joint probability is simply a product of the transition probabilities given by (3), for all dates $t \geq 2$, multiplied by the marginal probability of the initial state e_{i1} (see footnote 19). The fourth and last term is independent of the parameters and can be ignored. With respect to this fourth term, however, we should underline an important assumption made implicitly in (4), which is that $\overrightarrow{x}_i$ is independent of $(\varphi_i, \mathcal{M}_i)$. This assumption is important for step 2 of our estimation procedure, which we will discuss at the end of sub-section 4.2.

We are left with the third term to model, i.e. the joint distribution of the unobserved heterogeneity $(\varphi_i, \mathcal{M}_i)$. This is the subject of the next paragraph.

Unobserved individual heterogeneity. The last part of the model that we have to specify is the joint distribution of individual random effects, $\Pr(\varphi_i, \mathcal{M}_i | \Theta)$. Here we use a finite mixture approach and assume that individual i belongs to one of K classes of individuals, where all members of a given class $k \in \{1, \ldots, K\}$ share the same value $(\varphi_k, \mathcal{M}_k)$ of the various unobserved individual effects. Formally, we model $(\varphi_i, \mathcal{M}_i)$ as

$$\varphi_i = \sum_{k=1}^{K} \varphi_k \times \mathbf{1}_{k_i=k}, \quad \mathcal{M}_i = \sum_{k=1}^{K} \mathcal{M}_k \times \mathbf{1}_{k_i=k}, \quad (5)$$

where k_i is the unobserved class index of individual i. The joint distribution of $(\varphi_i, \mathcal{M}_i)$ is thus entirely characterized by that of k_i, i.e. the distribution of individuals into classes. The latter has K points of support. The class probabilities $\Pr(k_i = k | \Theta) = p_k$ form a set of $K - 1$ parameters to be estimated.

Estimation. With the above set of assumptions, the individual contribution to the complete likelihood (4) simplifies into:[20]

$$\mathcal{L}_i^c\left(\Theta; X_i, k_i\right) = \Pr\left(s_{i1}|x_{i1}, e_{i1}, k_i; \Theta\right) \times \Pr\left(e_{i1}|k_i; \Theta\right) \\ \times \prod_{t=2}^{T}\left[\Pr\left(s_{it}|x_{it}, e_{it}, k_i; \Theta\right) \times \Pr\left(e_{it}|e_{it-1}, k_i; \Theta\right)\right] \times \Pr\left(k_i|\Theta\right) \times \Pr\left(\overrightarrow{x}_i\right). \quad (6)$$

Now, since k_i is unobserved, we have to integrate it out of the likelihood function and maximize the sample log-likelihood:

$$\ln \mathcal{L}\left(\Theta; \mathbf{X}\right) = \sum_{i=1}^{N} \ln\left[\sum_{k=1}^{K} \mathcal{L}_i^c\left(\Theta; X_i, k\right)\right], \quad (7)$$

where $\mathbf{X} = \{X_i\}_{i=1}^N$ denotes the set of N individual observations in the sample. We carry out this maximization using a version of the EM algorithm described in Appendix B. Finally, standard errors are computed using the delta method.

4.2 Step 2: Analysis of job security indicators

Objectives. The individual/job type effects φ_i in our job security equation (1) reflect the impact of all permanent individual characteristics—observed or otherwise—on perceived job security across different job types. For instance, the subjective "cost" of having a temporary relative to a permanent job may vary from one individual to another as a function of unobserved psychological traits such as risk aversion. It may also be the case that the effect of temporary and permanent contracts on perceived job security depends on observed individual characteristics such as how distinct temporary and permanent contracts really are from the individual's viewpoint, which in turn depends on the particular legislation framing the use of temporary contracts in the individual's country of residence. Here we will highlight the correlations between φ_i, which we take as a "filtered" indicator of job security, and a set of covariates.

Construction of summary indicators of job security. We require a predictor of φ_i for each individual i in the sample. As a preliminary step in the construction of this predictor, we first

[20]Given a set of parameter values, our discrete factor model implies that the complete data $\{X_i, \varphi_i, \mathcal{M}_i\}$ is fully characterized by the set $\{X_i, k_i\}$, as the individual effects $(\varphi_i, \mathcal{M}_i)$ are fully characterized by individual class indices and parameter values.

compute the posterior probability that an individual i belongs to class k given the data X_i for this individual and our set of parameter estimates, $\widehat{\Theta}$. Using the notation introduced in step 1, this probability is given by:

$$\Pr\left(k_i = k | X_i; \widehat{\Theta}\right) = \frac{\mathcal{L}_i^c\left(\widehat{\Theta}; X_i, k\right)}{\sum_{\ell=1}^{K} \mathcal{L}_i^c\left(\widehat{\Theta}; X_i, \ell\right)}. \tag{8}$$

It turns out that these probabilities are in fact by-products of the EM algorithm that we use in the estimation procedure of step 1 (see Appendix B).

With these probabilities in hand, we construct a predictor of φ_i by simply taking the expectation of φ_k using the individual-specific class weights defined in (8):

$$\widehat{\varphi}_i = E_{\Pr\left(k|X_i;\widehat{\Theta}\right)}\varphi_k = \sum_{k=1}^{K}\left[\varphi_k \cdot \Pr\left(k_i = k|X_i;\widehat{\Theta}\right)\right]. \tag{9}$$

Note that this predictor $\widehat{\varphi}_i$ is a weighted average of all points in the support of the random effect φ_k. Because this support is discrete, the predictor $\widehat{\varphi}_i$ defined in (9) will not be one of the support points. While this may feel disturbing, one has to bear in mind that using $\widehat{\varphi}_i$ in the forthcoming analysis is equivalent to using a "simulation" approach where one (or any number of) realization(s) $\varphi_{\widehat{k}_i}$ of the random effect is attributed to each individual i by drawing one (or any number of) realization(s) of individual i's class $\widehat{k}_i$ from the distribution (8).[21]

We thus obtain a 3-dimensional vector of state-specific indicators of subjective job security:

$$\widehat{\varphi}_i = \left(\widehat{\varphi}_i^{\text{ppriv}}, \widehat{\varphi}_i^{\text{ppub}}, \widehat{\varphi}_i^{\text{temp}}\right)', \tag{10}$$

where each component of the vector corresponds to a particular job type in $\mathcal{E}$.

Explaining job security. We now turn to our statistical decomposition of the individual components of perceived job security. The basic idea that we pursue is to run OLS regressions of the

[21]Obviously, this method for obtaining a predictor of the individual job security random effect is not the only one imaginable. For instance, an alternative option would consist in predicting individual i's class $\widehat{k}_i$ by minimizing the mean squared prediction error as follows:

$$\widehat{k}_i = \underset{k\in\{1,\ldots,K\}}{\arg\min} \sum_{\ell=1}^{K}\left(\Pr\left(k_i = \ell|X_i;\widehat{\Theta}\right)\|\varphi_\ell - \varphi_k\|^2\right),$$

and then using $\widehat{\varphi}_i = \varphi_{\widehat{k}_i}$ as a predictor. The results with this alternative method are extremely similar to those presented here.

form:

$$\widehat{\varphi}_i^e = z_i'\alpha^e + v_i^e \tag{11}$$

for each separate job state $e \in \mathcal{E}$, where z_i is a vector of permanent characteristics of the individual, the individual's job and her particular labor market. Most importantly, z_i includes country-level policy indicators. We describe the exact specifications that we use below, as we discuss the estimated values of α^e.

Before we consider the estimation results, we should make three important remarks about this last step of our analysis in which we run regressions of the form (11).

First, in terms of the interpretation of the regression results, it may be useful to emphasize that this method merely provides a *descriptive* decomposition of individual perceived job security $\widehat{\varphi}_i$ into an observed heterogeneity component—the $z_i'\alpha^e$'s—and an orthogonal residual component—the v_i^e's. While it may be natural to think of perceived job security φ_i as a function of the z_i's and some unobserved heterogeneity variable, say ε_i, this function is fundamentally unidentified. Our goal here is to provide an intuitive description of job security data rather than to estimate a structural model of job security.

Second, the likelihood functions from step 1—see equations (4) and (6)—were written under the implicit assumption that $X_i = \{\overrightarrow{s}_i, \overrightarrow{x}_i, \overrightarrow{e}_i\}$ is independent of z_i conditional on k_i (or conditional on the pair (φ_i, ψ_i)). Moreover, as we already emphasized in the previous sub-section, those likelihood functions also contain the assumption that $\overrightarrow{x}_i \perp k_i$. The combination of these two assumptions implies independence of $\overrightarrow{x}_i$ and z_i. At this point, recall that $\overrightarrow{x}_i$ contains indicators of regional labor market conditions. Since z_i typically contains country-level policy indicators, the assumption that $\overrightarrow{x}_i \perp z_i$ may sound a bit heroic. To attenuate this criticism, we only include *temporary* variations in local labor market conditions in the vector of explanatory variables $\overrightarrow{x}_i$ so that $\overrightarrow{x}_i$ may be considered orthogonal to the country dummies (specifically, $\overrightarrow{x}_i$ contains year dummies, and the regional unemployment rate *taken in deviations from its 5-year mean value* rather than the level of the regional unemployment rate—see footnote 18). Permanent labor market conditions (as captured by the 5-year mean regional unemployment rate) are then included in the z_i variables.

Last, the dependent variables $\widehat{\varphi}_i^e$ in regressions (11) are affected by prediction errors. These render the computation of the standard errors on α^e difficult. Proper computation of those standard errors would involve running many bootstrap replications of our step 1, which is computationally too costly to consider undertaking on a PC. In the implementation below, we *do not* account for those estimation errors, but we note that the reported standard errors are likely underestimated.[22]

With those three remarks in mind, we now present and discuss the estimation results.

5 Estimation results

In practice we use $K = 8$ unobserved classes of individuals, each class corresponding to a unique value of φ and $\mathcal{M}$. Eight is the optimal number of classes according to the Normalized Entropy Criterion (NEC, see Celeux and Soromenho, 1996, and Appendix B).[23]

5.1 Step 1

Individual effects. The estimated class probabilities p_k appear in Table 2. The estimated values of the "job security" effect φ, which we denote as $\varphi_1, \ldots, \varphi_8$ are reported in Table 3 for each separate job state in $\mathcal{E}$. Finally, rather than displaying the 8 transition matrices $\mathcal{M}_1, \ldots, \mathcal{M}_8$ (which would take up a lot of space), we present the associated invariant probability distributions π_1^∞ to π_8^∞ in Table 4. (These distributions are defined over the four employment states by $\pi_k^{\infty\prime} \cdot \mathcal{M}_k = \pi_k^{\infty\prime}$ and measure the long-run probability for a member of class k of being in each particular employment state.)

[22]This last problem can be overcome by implementing a slightly different, single-step estimation method. The pros and cons of various approaches to estimating our model are discussed in Appendix C, where we also motivate our methodological choice.

[23]Other commonly used penalized likelihood criteria (AIC, BIC) suggest allowing for even more classes: the Schwarz-Bayesian Information Criterion (BIC) suggests 10 classes, while the Akaike Information Criterion (AIC) is still decreasing after 11 classes, which is as much as our computer could handle using the whole sample.(However, AIC is known to asymptotically overstate the number of classes.) We choose to follow the NEC for 3 reasons: first, as opposed to more general model selection criteria, it is specifically designed to select the number of classes in a finite mixture model; second, the computational cost of maximizing the likelihood increases quickly with the number of classes; and third, beyond 6 classes, increasing the number of classes did not make much qualitative difference to the results.

The class probabilities in Table 2 do not require much comment, besides the fact that they are all well above zero, so that none of the unobserved classes that our estimation procedure detects is of (probabilistically) negligible size.

Table 3 shows evidence of large-scale individual heterogeneity in job security perceptions. Yet it is clear that all classes feel less secure about temporary than permanent (public or private) jobs, bar the a priori puzzling case of class 1 who are somewhat averse to public jobs—feeling just as secure in temporary jobs and permanent public jobs! This becomes less of a paradox when we note (as we will see below) that members of class 1 are actually practically never employed in public sector jobs. Note also that 3 out of 8 classes (numbers 3, 6 and 7) view permanent public and permanent private jobs as equally secure.

Finally, the last row in Table 3 confirms that, on average, people feel more secure in public than in private sector jobs, and less secure in temporary than in permanent jobs. This latter series of results contrast with Table 1, which seemed to imply that *ceteris paribus* workers' average job security perceptions were roughly equal across job types, a result which we described as dubious given the obvious cross-job type differences in "objective" measures of job security (e.g. the year-to-year probability of a transition into nonemployment—see Table A2 in Appendix A). Indeed, Table 3 now suggests that this seeming equality was mostly driven by the selection of workers into different job types. While those effects differ across classes/individuals—and we shall dwell on these differences in the next section—, it seems generally true that "social insecurity" concerns temporary job holders, with public sector employees being relatively insulated. This is not entirely unexpected. Yet we still take it as a general indication that (1) workers' perceptions of their own job security partly reflect their objective condition—i.e. workers make "correct" assessments about their job security, and (2) workers reporting a low level of satisfaction with their job security really mean that they wish their job were *more* (as opposed to less) stable or protected.

Our next task is to consider the allocation of workers into employment states (Table 4). Again we see clear evidence of heterogeneity across worker classes. For instance, workers in classes 1 and 6 clearly tend to end up massively in permanent private sector jobs, while those from classes 5 and

7 are attracted to the public sector. Also, some workers (e.g., classes 4 and 8) seem to have trouble avoiding "undesirable" employment states—namely temporary jobs and nonemployment. This suggests that the "job security" individual effect φ_i may be determined in part by a psychological trait which also impacts on individual productivity (either at work or in job search) which in turn determines the type of jobs to which individuals have access.

Job insecurity and long-run employment states. Finally, we wish to assess the nature and extent of the potential selection biases that we mentioned in subsection 4.1. This amounts to analyzing the relationship between the job security random effects and the patterns of allocation into job states of the various classes.

One way to carry out this analysis is to examine jointly Tables 3 and 4. While this may reveal some "intuitively consistent" elements of the selection process (such as members of class 1 disliking public jobs and consistently selecting themselves away from public jobs), it probably does not provide the most useful summary picture of worker selection. More conveniently, our results allow us to compute the selection biases defined as follows:

$$\mathcal{B}(e_1) = E\left(\varphi_i^{e_1} | e_i = e_1\right) - E\left(\varphi_i^{e_1} | e_i \neq e_1\right). \tag{12}$$

This is the gap between average reported job security for a job of type $e_1 \in \mathcal{E}$ as perceived by workers in employment state e_1 and the average job security for that same job type e_1 as perceived by workers who are in an employment state *other than* e_1. Equation (12) thus takes up the familiar definition of selection biases from the "treatment effects" literature.[24]

The vector $\mathcal{B}$ is shown in Table 5 for the 3 fixed effect values φ^{ppriv}, φ^{ppub} and φ_i^{temp}, and the 3 relevant conditioning employment states "ppriv", "ppub" and "temp". We first note that

[24]The long-run version of (12) can be expressed as a function of the numbers reported in Tables 2 to 4:

$$\mathcal{B}(e_1) = \sum_{k=1}^{K} \left(\varphi_k^{e_1} \frac{p_k \pi_k^{\infty}(e_1)}{\sum_{\ell=1}^{K} p_\ell \pi_\ell^{\infty}(e_1)} \right) - \sum_{k=1}^{K} \left(\varphi_k^{e_1} \frac{p_k \left[1 - \pi_k^{\infty}(e_1)\right]}{1 - \sum_{\ell=1}^{K} p_\ell \pi_\ell^{\infty}(e_1)} \right),$$

where $p_k \pi_k^{\infty}(e_1) / \sum_{\ell=1}^{K} p_\ell \pi_\ell^{\infty}(e_1)$ (resp. $p_k \left[1 - \pi_k^{\infty}(e_1)\right] / \left[1 - \sum_{\ell=1}^{K} p_\ell \pi_\ell^{\infty}(e_1)\right]$) is the probability of belonging to class k conditional on being in employment state e_1 (resp. in an employment state other than e_1).

selection biases are fairly large: their magnitude is comparable to the differences across job states of the levels of the effects themselves (see Table 3). Next, looking at the signs of the three terms of Table 5, there is positive selection into permanent jobs. For instance, workers in permanent, private jobs feel more secure about permanent private jobs than workers in other employment states: $\mathcal{B}\,(\text{ppriv}) > 0$. Likewise, $\mathcal{B}\,(\text{ppub}) > 0$. While these conform with simple intuition, the negative sign of the third term is more puzzling. $\mathcal{B}\,(\text{temp}) < 0$, meaning that temporary job holders tend to be *more* temporary job-averse (in terms of security) than the average worker in other employment states. This again suggests that the allocation process of workers into job states is not entirely governed by workers' free choices based on their taste for particular job types: choices are constrained to some extent, even in the long-run.

We now turn back to our central equation of interest, equation (1). We begin by analyzing the impact of labor market conditions on job security.

Labor market conditions. The estimated coefficients on the observed time-varying covariates x_{it} (the β's in equation (1)) appear in Table 6.[25] Recall our proposed interpretation of latent job security s_{it}^* as a compound of the perceived *utility cost* of job loss and the subjective *probability* of that loss. The covariates entering the right hand side of equation (1) potentially impact both components of perceived job security.

The estimated coefficients on the year dummies suggest that job security tends to follow the cycle—with 1998 to 2000 being somewhat more "secure" years. There is no sign of a time trend; this is not unexpected, given that interviewees are asked to report their job security measured on a fixed 1-6 scale. A higher-than-normal local unemployment rate unsurprisingly tends to make workers more worried. Temporarily high local unemployment rates reflect bad local labor market conditions and thus indicate how easy or difficult it would be to find a new job in the case of dismissal.

[25] We do not report the cutoff points τ_h. Those are available upon request, together with their standard errors.

5.2 Step 2

Job security and individual characteristics. In this final section we present the results of a series of OLS regressions of the type shown in (11) repeated below for convenience:

$$\widehat{\varphi}_i^e = z_i'\alpha^e + v_i^e. \tag{11}$$

We use the same vector z_i of explanatory variables as in subsection 3.2.[26] We run this regression for the three job types: permanent private (ppriv), permanent public (ppub), and temporary (temp). The results of this series of regressions are shown in Table 7. They may be usefully compared to the results of the simple regression ignoring selection issues (see Table 1, subsection 3.2). In this paragraph we shall highlight the salient differences between the two series of results: as we shall see, our treatment of unobserved heterogeneity improves on the results from Table 1 by making them generally more clear-cut and intuitive.

As in Table 1, the cohort effects suggest that job security is U-shaped in age. The point estimates and standard errors are now similar across job types. The effects of education are also similar, though somewhat starker here than in Table 1: low-educated workers now feel less secure about *all* job types (the effect is now strongest for temporary jobs) than workers with high or intermediate levels of education, the latter two feeling roughly equally secure.

In contrast to the inconclusive results of Table 1, there is now evidence that foreign workers feel more insecure than natives in all types of jobs. However, compared to the significantly and substantially negative coefficients found for private and temporary jobs, the estimated coefficient in the public sector job regression is smaller in magnitude by about a half and of weak statistical significance.

[26] This includes age and age-squared, education (3 dummies), marital status, the presence of children under 15 in the household, an indicator of foreign citizenship, an indicator of the existence of a long-term unemployment spell in the recent past, the 5-year average local unemployment rate, and the OECD indicators of EPL strictness and UIB generosity.

Moreover, we should recall that the regressions in this section *do not* account for the estimation errors that affect the dependent variables $\widehat{\varphi}_i^e$. So again, the standard errors that we report for step 2 regressions are probably understated and all the ensuing considerations about statistical inference are only indicative.

There is also weak evidence that living in a couple affects job security positively, and that having children in the household makes men feel more insecure about private and temporary jobs. Neither of these effects was found in Table 1. These can be interpreted as an indication of family insurance: having children makes job loss more costly, while the presence of a spouse who can fulfill the role of second breadwinner alleviates insecurity. Moreover, interestingly, the effect of having children in the household vanishes in public sector jobs.

Past experience of long-term unemployment reduces perceived job security in all types of jobs. From the point estimates, this effect is around twice as large for temporary jobs than for any of the permanent job types. We also find a negative correlation between the average local unemployment rate and perceived job security in permanent private and temporary jobs (with a somewhat stronger effect for temporary jobs). But its effect on perceived job security in public jobs is strongly *positive* (and not very precisely estimated). These particular results are qualitatively similar—but again more clear cut—to those found in Table 1. Our tentative interpretation that workers facing adverse labor market conditions would wish to have more protected jobs continues to hold.

Table 7 also reports a constant (first row), the values of which show the job security ranking of job types: temporary job holders feel less secure than permanent, private job holders, while those in permanent public jobs feel more secure. This statement is based on point estimates and the relatively large standard errors affecting the corresponding coefficients make it likely that the *ceteris paribus* differences in job security across job types are of weak statistical significance. However, this implicit security ranking of job types is more intuitive than that obtained from Table 1, a difference which may be taken to confirm the importance of selection effects.

At this point, the picture that Table 7 sketches is that, while the perception of job security for either permanent private or temporary jobs varies significantly with local labor market conditions, recent unemployment experience, citizenship, and to some extent with family status, these controls only appear statistically insignificantly or with much reduced power in the regression (11) for φ^{ppub} (i.e., in the public sector). Compared to private or temporary jobs, public jobs are thus perceived to be safe, and insulated from labor market shocks.

Job security and policy. Finally, the last two rows of Table 7 report the estimated effects of our indicators of EPL strictness and UIB generosity. The results are striking. We first see that correlation between job security and EPL is negative and significant , while that with UIB is positive and significant. This holds in both temporary and permanent, private jobs, thus confirming the visual impression given by Figures 2 and 3. For public jobs, however, perceived job security is essentially *uncorrelated* with the stringency of EPL, and is only very weakly positively correlated with the generosity of UIB (the estimated coefficient on the latter indicator is smaller in magnitude by a factor of 5 to 8 compared to that in the private and temporary job regressions, and its statistical significance is borderline). Hence, again, public sector jobs are perceived as being largely insulated from the risk of job loss.

A closer look at the numbers further reveals that, while perceived job security in temporary and permanent private jobs are roughly equally sensitive to EPL, unemployment compensation is relatively more important for temporary (relative to permanent) jobs. The estimated coefficient on the UIB indicator is about 60 percent larger in magnitude in the temporary than in the permanent private job regression.

At this point it thus seems safe to conclude that male workers holding either a temporary or a permanent, private job feel more secure in countries with generous UIB but relatively *low* EPL (at least as measured by the OECD indicators). Neither composition effects—due to demographic differences between countries or to particular selection patterns of workers into specific job types, based on the former's observed and unobserved individual characteristics—nor the trade-off between EPL strictness and UIB generosity can explain why workers in countries with stricter EPL and less generous UIB are more worried about their job security.

6 Interpretation of the results

How should we interpret these correlations between job security, EPL and UIB? In this section we review some theoretical arguments from two strands of literature: the "macro-labor" literature, which focuses on the causal effect of institutions on labor market outcomes (job security in particular), and the "political economy of institutions" literature, which considers causality running in the opposite direction.

6.1 The impact of EPL and UIB on job security

Stricter EPL leads to longer unemployment durations, both theoretically and empirically. For employees, EPL is therefore a double-edged sword: while protecting them by reducing the risk of job loss, it also increases the associated cost of job loss by reducing the outflow rate from unemployment. Standard calibrations of the Mortensen and Pissarides (1994) matching model—the modern theoretical workhorse of labor market policy scholars—predict that the second phenomenon is most likely to dominate, and most strongly so for the tier of the labor market that is not subject to strong firing restrictions—i.e. for temporary jobs—(see e.g. Cahuc and Postel-Vinay, 2002, for an application with both temporary contracts and firing restrictions). One interpretation of the negative correlations appearing in Table 7 is that these predictions are correct, at least for those job types that are most obviously exposed to some "objective" risk of job loss—namely temporary jobs and permanent private jobs. The generosity of UIB, on the other hand, has no evident cross-country correlation with objective aggregate measures of labor market risk such as mean job or unemployment spell hazards. As a first approximation, it can thus probably be considered preferable to EPL as an insurance tool against labor market risk.[27] This, at least, is the stated view of the European council (see the quotation in the Introduction), and Table 7 suggests that workers, in many cases, share this view.

6.2 Job security as a determinant of institutions

A separate "political economy of institutions" literature inverses the causality running from institutions to labor market outcomes, and considers the arrow running in the opposite direction.[28] A recurring idea in this latter strand of literature is that it is in the interest of "insiders" to support strict EPL, while "outsiders" are more likely to favor generous UIB. As Boeri et al. (2001)

[27]One component of EPL, severance payments, which is a pure transfer from firm to (former) worker, potentially plays a true insurance role. Yet all other components of EPL (procedural costs, waiting periods, judicial costs) are deadweight costs for the firm-worker match.

[28]Saint-Paul (2000, 2002), Boeri et al. (2001) and Boeri et al. (2003) are recent examples addressing the specific issues of UIB and/or EPL.

note, "*EPL concentrates the unemployment risk among 'outsiders'*". While our results do not convey a direct test of this statement, they certainly are consistent with it. To see this, we can construct a measures of the "individual gain (in terms of job security) to being an insider" as either $\widehat{\varphi}_i^{\text{ppub}} - \widehat{\varphi}_i^{\text{temp}}$ or $\widehat{\varphi}_i^{\text{ppub}} - \widehat{\varphi}_i^{\text{ppriv}}$ and regress these measures on the covariates z_i and our country-level measures of UIB and EPL. The coefficient estimates on these latter two variables are shown in the first two columns of Table 8, where we see that the gain to being an insider significantly increases with EPL strictness and significantly decreases with UIB generosity.[29]

We can further show, however, that this result is sensitive to the fact that we are assimilating "being an insider" to "holding a permanent, public job". The differential in perceived job security between permanent private and temporary contracts, $\widehat{\varphi}_i^{\text{ppriv}} - \widehat{\varphi}_i^{\text{temp}}$, appears uncorrelated with EPL or UIB (see the third column of Table 8).

Up to this latter qualification, our results thus seem broadly consistent with the theoretical political economy message about who profits and who suffers from a given mix of EPL and UIB. Yet turning back to the first two columns of Table 7, one sees that those in permanent private jobs—many of whom may arguably be considered to be "insiders" and who surely constitute a political majority—feel less secure when facing stricter EPL and less generous UIB.[30] This is somewhat intriguing, particularly if one seeks to understand the emergence of a low UIB-high EPL regime as a political equilibrium (Boeri et al., 2003).

One, difficult to test, interpretation is that workers suffer from a certain kind of myopia, and do not take into account the negative effects of EPL on unemployment duration, while instead concentrating on the (immediate) positive effect on firing. An explanation along similar lines is over-optimism: while individuals feel bad about others who are potentially harmed when they

[29] We omit the estimated coefficients on the remaining covariates in z_i. They are available upon request. An interesting point to note regarding these coefficients is that the job security gain to being an insider always unambiguously increases in the face of adverse labor market conditions, as measured by a high local unemployment rate.

[30] We use a shortcut here. All the type-(11) regressions summarized in Table 7 and 8 were run on the *entire* population. Running separate regressions for holders of the various specific job types, or weighting the data by the individual long-run probabilities of holding specific job types (the π_k^∞'s) leads to qualitatively unchanged results.

lose their jobs, they do not imagine that they themselves will be in that situation. As such, they overvalue the benefits of EPL relative to the costs.

Another candidate interpretation is that the making of labor market policy escapes the majority rule. It may indeed generically be the case that the political process is sufficiently biased toward public sector employees to make it look like they were the political group deciding upon the strictness of employment protection or the generosity of unemployment compensation. How reasonable such a statement sounds to a layperson probably varies a lot between European societies (as well as within each society, for that matter). One of the main points that we make here, however, is that the data do not seem to contain much evidence against it.

A third explanation would dispute that EPL is, on average, bad for workers. The gist of this argument is that EPL affects security negatively, but some other, unmeasured, aspect(s) of the job positively. The overall positive impact of EPL on workers' jobs leads to them being in favor of it.

Alternatively, *de jure* dismissal restrictions as measured by the EPL index may not translate to *de facto* practice. Yet, for job security to rise with EPL would require that *de facto* and *de jure* restrictions be negatively correlated, which seems unlikely. Further, Pierre and Scarpetta (2004) use international survey data on 17000 firms to show that employers' concerns about regulations are closely correlated with the Laws' *de jure* stringency.

Last, it may be the case that the degree of job insecurity is a driving force behind the desire for protection, and that this latter only partially alleviates the problem. A parallel would then be stronger police presence in high crime areas: it is the latter (ex ante) which determines the former. The small amount of temporal movement in EPL, and the difficulty of accurately measuring changes, preclude any analysis of co-movements in security and protection. We suspect, however, that the endogeneity of institutions may turn out to play a key role in the relationship between EPL and job security.

7 Concluding remarks

This paper contributes to the economic policy debate by examining the link between labor market institutions and job security. We use data from the European Community Household Panel to construct indicators of perceived job security for 3 different types of job contracts—permanent private, permanent public, and temporary—in 12 different EU countries. We then examine the relationship between job security and labor market institutions, specifically employment protection and unemployment benefit generosity.

The overall conclusions are that perceived job security in non-public sector jobs is lower in countries with stricter employment protection legislation but higher in countries with more generous unemployment benefits. These effects are not found for public jobs, which seem to be more "universally" perceived as safe jobs (i.e. insulated from labor market shocks). These conclusions hold controlling for composition effects and controlling for sorting by workers into job types.

Our interpretation of these results remains speculative, as we cannot carry out direct tests of many hypotheses. One key point to bear in mind is that the effect of EPL on job security broadly defined (including future employment prospects) is theoretically ambiguous. It is also possible that we have uncovered some kind of a political equilibrium, whereby those who profit from higher EPL (secure insiders, say) push for more protection, and politicians are responsive to this pressure. It remains to be explained, however, why this would hold when it appears that the majority of workers feel less secure in higher EPL/less generous UIB environments.

Nevertheless, it seems clear that employment protection, as measured by the OECD indicator, does not by itself afford good protection against the feeling of job insecurity, whereas unemployment benefits do play something of an insurance role. Interestingly, the European Union's own observatory on the quality of work makes extensive reference to protection (European Foundation for the Improvement of Living and Working Conditions, 2002). However, this is always in the context of social protection, rather than pure job protection. In this sense, the Danish model of "flexicurity" may be what workers really want, although they do not necessarily realize it.

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APPENDIX

A Sample description

Construction. The European Panel (ECHP) is a common-questionnaire panel of household and individual data gathered by EUROSTAT, covering twelve EU countries over a maximum of eight waves (1994 to 2001).[31] The survey was extended to fifteen countries with the later addition of Austria, Finland and Sweden. Due to missing data, we only use twelve of these fifteen countries, and five of the eight waves: satisfaction with job security was not asked in Germany, Luxembourg or Sweden. Also, there are abnormally high proportions of non-responses to that question among temporary job holders in France in wave 3.[32]

It should also be noted that the UK left the ECHP in 1997, and that subsequent data is ex-post harmonized from the British Household Panel Study (BHPS). It turns out that BHPS data have higher non-response rates than do ECHP data to the job security question.

Our final sample consists of male workers aged between 20 and 55 in 1997, who are observed to be either wage earners or nonemployed at every yearly interview between 1997 and 2001. Individuals who were self-employed in at least one year during our observation window were left out of the sample, as were individuals consistently reporting nonparticipation. We thus end up with 12,091 individuals $\times$ 5 waves, the country distribution of which is described in Table A1.

< Table A1 about here. >

Job security. The job security variable is known as item PE032 in the ECHP user data base. The exact wording of that question and the per-country distribution of replies are presented in the main text. We note that this question comes second in a series of 7 job satisfaction questions that appear in the "employment" section of the questionnaire. It turns out that the responses to many of these satisfaction questions are highly correlated.

Job mobility. Individual i's employment state at date t is denoted by e_{it}. We distinguish four employment states: employed in the private sector with a permanent contract (e_{it} = ppriv); employed in the public sector with a permanent contract (e_{it} = ppub); employed with a temporary contract (e_{it} = temp);[33] and nonemployed (e_{it} = none). Note that all the information used to construct the state indicator e_{it} is reported by the individual. In particular, the definition of "temporary contract" is somewhat arbitrary and can vary from one individual to the other.

[31] Details on the ECHP are available at the European Panel Analysis Group (EPAG) website (http://www.iser.essex.ac.uk/epag/user-network.php).

[32] Similar problems appear to a lesser extent in the U.K. data in waves 7 and 8. Moreover, the contract type (short vs. long term) is largely missing in Portugal in wave 7. These latter problems can be fixed to a great extent by bringing missing information over from adjacent waves.

[33] In principle we could have split further between temporary, private and temporary, public. However, this raises considerable computational difficulties due to the scarcity of transitions from, e.g. permanent, public jobs to temporary, private jobs.

< **Figure A1 and Table A2 about here.** >

The distribution of individuals across states, and the matrix of observed transitions are displayed in Figure A1 and Table A2, respectively.

B The EM algorithm

In this Appendix we briefly describe our application of the EM algorithm for finite mixtures. For a general presentation see Dempster et al. (1977), or Bilmes (1998) for an excellent applied tutorial. The algorithm goes through the following two steps:

1. **Expectation (E)-step.** Given starting values of the parameters Θ_0, we first update the mixing proportions which are equal to the posterior joint density of k conditional on the observables X_i:

$$\Pr\left(k_i = k | X_i; \Theta_0\right) = \frac{\mathcal{L}_i^c\left(\Theta_0; X_i, k\right)}{\sum_{\ell=1}^K \mathcal{L}_i^c\left(\Theta_0; X_i, \ell\right)}. \tag{13}$$

We then use those mixing proportions to compute the *expected value* of individual i's contribution to the sample *log*-likelihood, given our initial parameter value Θ_0:

$$E\left[\ln \mathcal{L}_i^c\left(\Theta; X_i^c\right) | X_i; \Theta_0\right] = \sum_{k=1}^K \Pr\left(k_i = k | X_i; \Theta_0\right) \times \ln \mathcal{L}_i^c\left(\Theta; X_i, k\right). \tag{14}$$

2. **Maximization (M)-step.** The M-step simply consists in maximizing the *expected* sample log likelihood, given the starting parameter values Θ_0:

$$\begin{aligned} \widehat{\Theta}|_{\Theta_0} &= \arg\max_{\Theta} \sum_{i=1}^N E\left[\ln \mathcal{L}_i^c\left(\Theta; X_i^c\right) | X_i; \Theta_0\right] & (15) \\ &= \arg\max_{\Theta} \sum_{i=1}^N \sum_{k=1}^K \Pr\left(k_i = k | X_i; \Theta_0\right) \times \ln \mathcal{L}_i^c\left(\Theta; X_i, k\right). & (16) \end{aligned}$$

This delivers an updated set of parameter estimates, $\widehat{\Theta}|_{\Theta_0}$, which we then use as new starting values in the E-step.

In our application, using equations (2) to (6), individual i's contribution to the complete log-likelihood can be written as the following function of the parameters:

$$\begin{aligned} \ln \mathcal{L}_i^c\left(\Theta; X_i, k_i\right) = \sum_{t=1}^T \ln\left(N\left[\tau_h - x_{it}'\beta - \varphi_{k_i}^{e_{it}}\right] - N\left[\tau_{h-1} - x_{it}'\beta - \varphi_{k_i}^{e_{it}}\right]\right) \\ + \ln \pi_{k_i}^1\left(e_{i1}\right) + \sum_{t=2}^T \ln \mathcal{M}_{k_i}\left(e_{it}, e_{it-1}\right) + \ln p_{k_i}. \end{aligned} \tag{17}$$

Given our set of initial parameters Θ_0, we first compute the mixing proportions using (13) and (17). We then maximize

$$\sum_{i=1}^{N}\sum_{k=1}^{K}\Pr\left(k_i=k|X_i;\Theta_0\right)\times\left(\sum_{t=1}^{T}\ln\left(N\left[\tau_h-x_{it}'\beta-\varphi_{k_i}^{e_{it}}\right]-N\left[\tau_{h-1}-x_{it}'\beta-\varphi_{k_i}^{e_{it}}\right]\right)\right.$$
$$\left.+\ln\pi_k^1\left(e_{i1}\right)+\sum_{t=2}^{T}\ln\mathcal{M}_k\left(e_{it},e_{it-1}\right)+\ln p_k\right) \quad (18)$$

with respect to $\Theta=\left(\beta,\varphi,\tau,\pi^1,\mathbf{M},\mathbf{p}\right)$, where boldface letters designate vectors of parameters (e.g. $\mathbf{M}=\left(\mathcal{M}_k\right)_{1\leq k\leq K}$). This maximization problem is separable to some extent: a first subset of parameters—those involved in the first line of (18), namely (β,φ,τ)—are estimated by running a weighted ordered probit regression of reported job security $\overrightarrow{s}_i$ against $(\overrightarrow{x}_i,\overrightarrow{e}_i)$ and a class index k_i according to the job security equation (1) using the mixing proportions (13) as weights. The complementary subset of parameters—those involved in the second line of (18), namely $\left(\pi^1,\mathbf{M},\mathbf{p}\right)$—can be obtained in closed form from the relevant first-order conditions:

$$\widehat{\pi}_k^1\left(e\right)|_{\Theta_0}=\frac{\sum_{i=1}^{N}\left[\mathbf{1}_{e_{i1}=e}\Pr\left(k_i=k|X_i;\Theta_0\right)\right]}{\sum_{i=1}^{N}\Pr\left(k_i=k|X_i;\Theta_0\right)}; \quad (19)$$

$$\widehat{\mathcal{M}}_k\left(j,\ell\right)|_{\Theta_0}=\frac{\sum_{i=1}^{N}\sum_{t=2}^{T}\left[\mathbf{1}_{e_{it-1}=\ell}\mathbf{1}_{e_{it}=j}\Pr\left(k_i=k|X_i;\Theta_0\right)\right]}{\sum_{i=1}^{N}\sum_{t=2}^{T}\left[\mathbf{1}_{e_{it-1}=\ell}\Pr\left(k_i=k|X_i;\Theta_0\right)\right]}; \quad (20)$$

$$\widehat{p}_k|_{\Theta_0}=\frac{1}{N}\sum_{i=1}^{N}\Pr\left(k_i=k|X_i;\Theta_0\right). \quad (21)$$

At this point, we have an update $\widehat{\Theta}|_{\Theta_0}$ for all the parameters, which we compare with our initial guess, Θ_0. If they are close enough, the algorithm stops, else we start again at the E-step using $\widehat{\Theta}|_{\Theta_0}$ as a new initial guess.[34]

Finally, the parsimony criterion used to select the number of classes K is the Normalized Entropy criterion (NEC) proposed by Celeux and Soromenho (1996), which is given by:

$$NEC\left(K\right)=\frac{-\sum_{k=1}^{K}\sum_{i=1}^{N}\Pr\left(k_i=k|X_i;\widehat{\Theta}_K^*\right)\ln\left[\Pr\left(k_i=k|X_i;\widehat{\Theta}_K^*\right)\right]}{\ln\mathcal{L}\left(\widehat{\Theta}_K^*;\mathbf{X}\right)-\ln\mathcal{L}\left(\widehat{\Theta}_1^*;\mathbf{X}\right)}, \quad (22)$$

where $\widehat{\Theta}_K^*$ is the vector of parameter estimates for a model with K classes. The denominator in the latter formula is thus the log of the likelihood ratio between the K-class model and the single-class model. In the case of this paper, $NEC\left(K\right)$ is minimized at $K=8$.

C Methodological issues

In this Appendix we briefly discuss the pros and cons of our two-step method, vis-à-vis a more direct, one-step approach.

[34] In practice our convergence criterion is that the maximum relative increase among the components of Θ be less than 1/100th of a percentage point. When this criterion is met, the marginal percent increase in the sample likelihood following an additional iteration is in the order of 10^{-2} percent.

The basic problem that we are trying to solve is the following. We have a model positing that subjective (reported) job security, $\overrightarrow{s}_i$, depends on a number of time-varying observed characteristics $(\overrightarrow{x}_i, \overrightarrow{e}_i)$, on some time-invariant individual characteristics z_i, and on an unobserved time-invariant characteristic k_i. Taking up the notation (from the main text) $X_i = \{\overrightarrow{s}_i, \overrightarrow{x}_i, \overrightarrow{e}_i\}$, we can write the joint probability of a typical observation i (given parameter values Θ) as:

$$\begin{aligned} \Pr(X_i, k_i, z_i|\Theta) &= \Pr(X_i|k_i, z_i; \Theta) \times \Pr(k_i|z_i; \Theta) \times \Pr(z_i) \\ &= \Pr(X_i|k_i; \Theta) \times \Pr(k_i|z_i; \Theta) \times \Pr(z_i), \end{aligned} \tag{23}$$

where the second equality comes from the implicit assumption that $X_i \perp z_i|k_i$.

Our problem is to estimate the parameter Θ. Our approach to this problem takes two steps. We first maximize the *marginal* sample likelihood of $\mathbf{X}$, $\mathcal{L}(\Theta; \mathbf{X}) = \int \Pi_{i=1}^N \Pr(X_i, k_i|\Theta)\, d\mathbf{k}$—that is, we integrate z_i out of (23) and maximize the resulting marginal likelihood. Second, we predict a value $\widehat{k}_i$ of k_i for each i following the protocol presented insubsection 4.2, and look at moments of the conditional distribution *of this predictor* $\widehat{k}_i$ given z_i—essentially, regressions of the form (11) compute $E\left(\widehat{k}_i|z_i\right)$. An obvious drawback of this two-step approach is that $\widehat{k}_i$ is only an imperfect predictor of the true k_i. As discussed in the main text, this somewhat weakens the results obtained in our second step.

An alternative, more direct (one-step) approach to this problem would be to maximize the full sample log-likelihood $\mathcal{L}(\Theta; \mathbf{X}, \mathbf{z}) = \int \Pi_{i=1}^N \Pr(X_i, k_i, z_i|\Theta)\, d\mathbf{k}$. This can be done using for instance an EM algorithm similar to the one described in Appendix B. Note however that this approach requires that one specifies (parametrically) the conditional probability $\Pr(k_i|z_i; \Theta)$. Given a parametric specification, this single-step approach has the advantage (over the two-step approach) of directly delivering an estimate of the conditional probability $\Pr(k_i|z_i; \Theta)$, which is essentially what we are interested in in our step 2.

The problem with the single step approach is that the conditional distribution $\Pr(k_i|z_i; \Theta)$ is not non-parametrically identified. Yet the estimates obtained in the single-step method are *a priori* sensitive to the particular parametric assumption made about the form of $\Pr(k_i|z_i; \Theta)$. This problem is circumvented by the two-step method, where we only impose an arbitrary structure on $\Pr(k_i|z_i; \Theta)$ in the second and last step. Step 1 of the two-step method thus delivers estimates of the subset of parameters that enter $\Pr(X_i|k_i; \Theta)$ and of the marginal class probabilities $\Pr(k_i)$ which are not polluted by potential specification errors affecting $\Pr(k_i|z_i; \Theta)$.

An additional advantage of the two-step method is that it is considerably less burdensome in terms of computation. In particular, once step 1 is completed and the $\widehat{k}_i$'s are constructed, we can try any specification we want in the second-step regressions (11) at practically zero computational cost (since those are simple linear regressions). By contrast, changing the specification of $\Pr(k_i|z_i; \Theta)$ in the single-step method implies re-running the whole likelihood maximization, which takes hours of computing time. This last practical argument convinced us to opt for the two-step approach.

Table A1: Number of individuals per country

AUT	BEL	DNK	ESP	FIN	FRA	
853	809	777	772	742	1,981	
GBR	**GRC**	**IRL**	**ITA**	**NLD**	**PRT**	**Total**
1,471	936	339	931	1,521	959	**12,091**

Table A2: Observed transitions between employment states

		Current state $e_{it} = \ldots$			
		Permanent Private	**Permanent Public**	**Temporary**	**Nonemployment**
Past state	**Permanent Private**	92.41	1.75	2.89	2.94
	Permanent Public	4.44	92.36	1.30	1.90
$e_{it-1} = \ldots$	**Temporary**	32.21	6.57	47.67	13.55
	Nonemployment	22.79	4.23	20.52	52.46

Table 1: Ordered Probit regression of job security[1]

Dependent variable: perceived job security (s_{it}), 1997.			
Explanatory variables	**Effect among workers whose job type is $e = \ldots$**		
	Permanent Private	Permanent Public	Temporary
Constant	0 (ref.)	−0.433 (.548)	−0.948 (.583)
Age (/10)	−0.479 (.131)	−0.322 (.251)	−0.731 (.295)
Age-squared (/100)	0.053 (.017)	0.042 (.031)	0.093 (.040)
High education	0 (ref.)	0 (ref.)	0 (ref.)
Intermediate education	−0.041 (.034)	−0.145 (.047)	0.119 (.093)
Low education	−0.061 (.035)	−0.255 (.052)	−0.034 (.091)
Foreign	−0.305 (.092)	0.048 (.238)	0.102 (.245)
Couple	0.015 (.038)	0.081 (.064)	0.013 (.083)
Has children	0.005 (.030)	0.006 (.048)	0.031 (.076)
Past unemployment experience	−0.115 (.061)	−0.367 (.123)	−0.345 (.096)
Mean local unempl. rate[2]	0.241 (.356)	2.382 (.414)	−1.685 (.634)
EPL	−0.138 (.017)	0.059 (.029)	0.038 (.054)
UIB	0.038 (.003)	0.036 (.005)	0.047 (.009)

Notes: [1]Standard errors in parentheses.
[2]Mean over the observation period, 1997-2001.

Table 2: Class probabilities

p_1	p_2	p_3	p_4	p_5	p_6	p_7	p_8
0.061	0.249	0.044	0.158	0.134	0.208	0.055	0.092
(.006)	(.007)	(.006)	(.009)	(.006)	(.010)	(.006)	(.006)

Table 3: Job security fixed effects

Job state:	Perm. priv. (φ_k^{ppriv})	Perm. pub. (φ_k^{ppub})	Temporary (φ_k^{temp})
φ_1	0	-3.76	-3.73
	(ref.)	(.124)	(.099)
φ_2	-0.98	-0.68	-1.16
	(.067)	(.093)	(.130)
φ_3	-2.10	-2.11	-3.02
	(.087)	(.111)	(.084)
φ_4	-2.11	0.03	-2.72
	(.076)	(.080)	(.112)
φ_5	-3.29	-1.10	-4.06
	(.076)	(.110)	(.093)
φ_6	-2.11	-2.12	-3.71
	(.074)	(.103)	(.092)
φ_7	-0.47	-0.48	-0.97
	(.116)	(.105)	(.083)
φ_8	0.84	1.29	0.43
	(.090)	(.197)	(.109)
Mean[1]	**-1.80**	**-0.98**	**-2.41**

Note: [1]The mean effect for each job state e equals $\sum_k p_k \varphi_k^e$, where the p_k values are those in Table 2.

Table 4: Invariant job state distributions

Job state	**Invariant distributions ($\pi_k^{\infty\prime} \cdot \mathcal{M}_k = \pi_k^{\infty\prime}$):**[1,2]								
	π_1^∞	π_2^∞	π_3^∞	π_4^∞	π_5^∞	π_6^∞	π_7^∞	π_8^∞	π_{mean}^∞[3]
Perm. priv.	.822	.905	.469	.299	.266	.855	.084	.468	**.604**
Perm. publ.	.023	.013	.353	.498	.628	.090	.738	.333	**.273**
Temporary	.124	.031	.068	.137	.035	.019	.072	.010	**.053**
Nonemp.	.026	.051	.109	.066	.070	.036	.106	.190	**.070**

Notes: [1]Subscripts designate classes.
[2]Standard errors are not reported (available on request).
[3]The mean distribution equals $\sum_k p_k \pi_k^\infty$, where the p_k values are those in Table 1.

Table 5: Selection

Vector $\mathcal{B}$		
$E\left(\varphi^{\text{ppriv}} \vert e = \text{ppriv}\right) - E\left(\varphi^{\text{ppriv}} \vert e \neq \text{ppriv}\right)$	$=$	0.345
$E\left(\varphi^{\text{ppub}} \vert e = \text{ppub}\right) - E\left(\varphi^{\text{ppub}} \vert e \neq \text{ppub}\right)$	$=$	0.587
$E\left(\varphi^{\text{temp}} \vert e = \text{temp}\right) - E\left(\varphi^{\text{temp}} \vert e \neq \text{temp}\right)$	$=$	-0.275

Table 6: Estimated coefficients from the job security equation (1)[1]

1997	0 (ref.)	2000	0.067 (.020)
1998	0.046 (.017)	2001	0.028 (.021)
1999	0.036 (.018)	local unempl. rate[2]	-1.226 (.455)

Notes: [1]Standard errors in parentheses.
[2]In deviation from its 1997-2001 mean.

Table 7: Second-step regressions—equation (11)[1]

Explanatory variables (z_i)	Dependent variable: $\widehat{\varphi}_i^{\text{ppriv}}$ (perm. priv.)	$\widehat{\varphi}_i^{\text{ppub}}$ (perm. pub.)	$\widehat{\varphi}_i^{\text{temp}}$ (temporary)
Constant	−0.593 (.179)	−0.440 (.189)	−1.400 (.224)
Age (/10)	−0.403 (.099)	−0.367 (.104)	−0.488 (.124)
Age-squared (/100)	0.050 (.013)	0.054 (.014)	0.064 (.016)
High education	0 (ref.)	0 (ref.)	0 (ref.)
Intermediate education	−0.009 (.026)	−0.074 (.028)	−0.032 (.033)
Low education	−0.163 (.028)	−0.199 (.029)	−0.246 (.035)
Foreign	−0.254 (.081)	−0.126 (.085)	−0.274 (.101)
Couple	0.052 (.030)	0.031 (.032)	0.077 (.038)
Has children	−0.024 (.026)	−0.009 (.026)	−0.037 (.031)
Past unemployment experience	−0.172 (.043)	−0.119 (.045)	−0.258 (.054)
Mean local unempl. rate[2]	−0.424 (.241)	0.380 (.254)	−0.618 (.301)
EPL	−0.148 (.015)	0.007 (.016)	−0.134 (.019)
UIB	0.883 (.079)	0.179 (.083)	1.400 (.099)

Notes: [1]Standard errors in parentheses.
[2]Mean over the observation period, 1997-2001.

Table 8: Job security and policy

Dependent variable:	$\widehat{\varphi}^{\text{ppub}} - \widehat{\varphi}^{\text{temp}}$	$\widehat{\varphi}^{\text{ppub}} - \widehat{\varphi}^{\text{ppriv}}$	$\widehat{\varphi}^{\text{ppriv}} - \widehat{\varphi}^{\text{temp}}$
EPL	0.141 (.013)	0.155 (.018)	−0.013 (.011)
UIB	−0.746 (.069)	−0.705 (.097)	−0.041 (.060)

Note: Standard errors in parentheses.

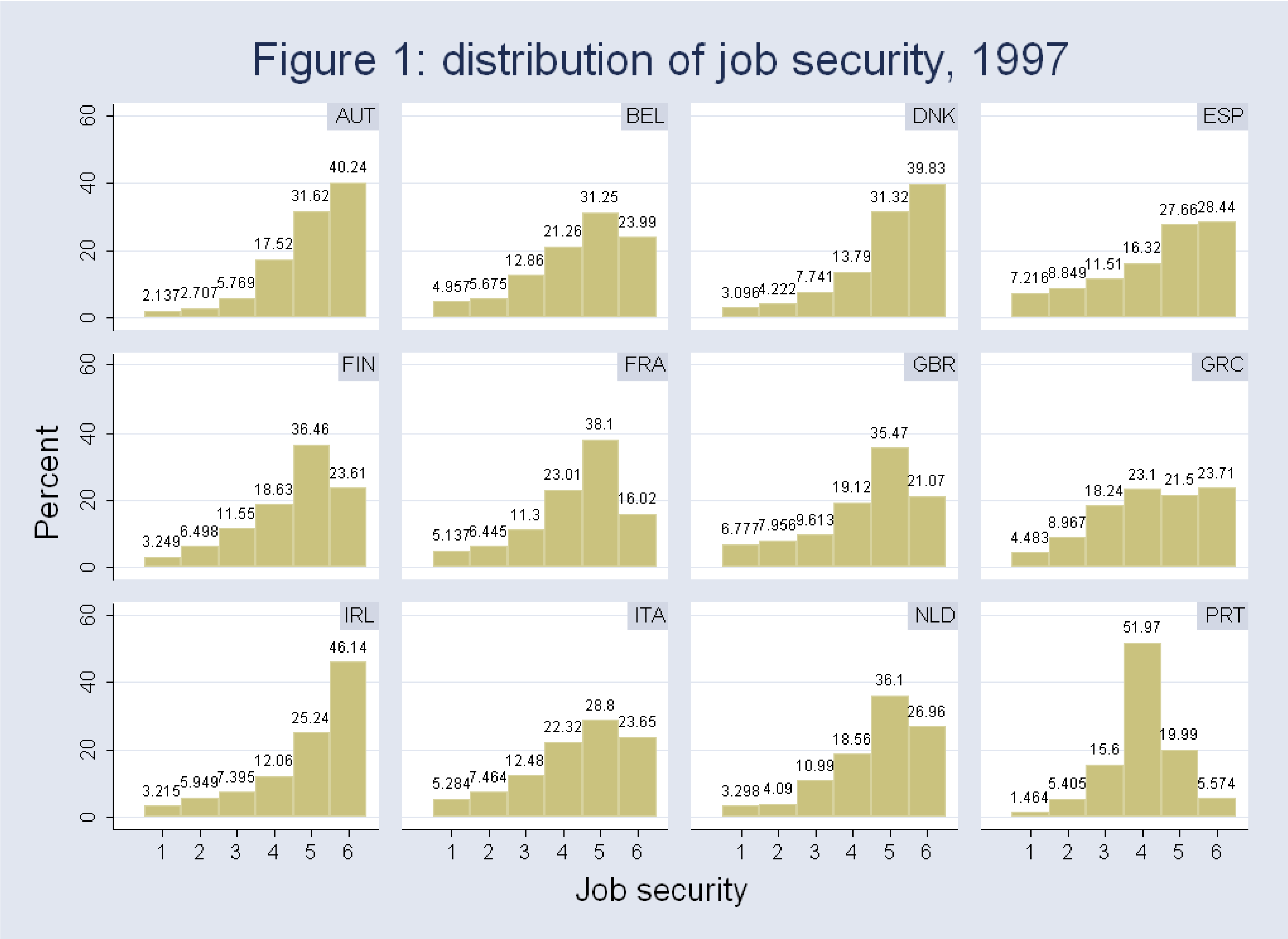


Figure 1: distribution of job security, 1997
AUT
BEL
DNK
ESP
FIN
FRA
GBR
GRC
IRL
ITA
NLD
PRT
Percent
Job security
2.137 2.707 5.769 17.52 31.62 40.24
4.957 5.675 12.86 21.26 31.25 23.99
3.096 4.222 7.741 13.79 31.32 39.83
7.216 8.849 11.51 16.32 27.66 28.44
3.249 6.498 11.55 18.63 36.46 23.61
5.137 6.445 11.3 23.01 38.1 16.02
6.777 7.956 9.613 19.12 35.47 21.07
4.483 8.967 18.24 23.1 21.5 23.71
3.215 5.949 7.395 12.06 25.24 46.14
5.284 7.464 12.48 22.32 28.8 23.65
3.298 4.09 10.99 18.56 36.1 26.96
1.464 5.405 15.6 51.97 19.99 5.574

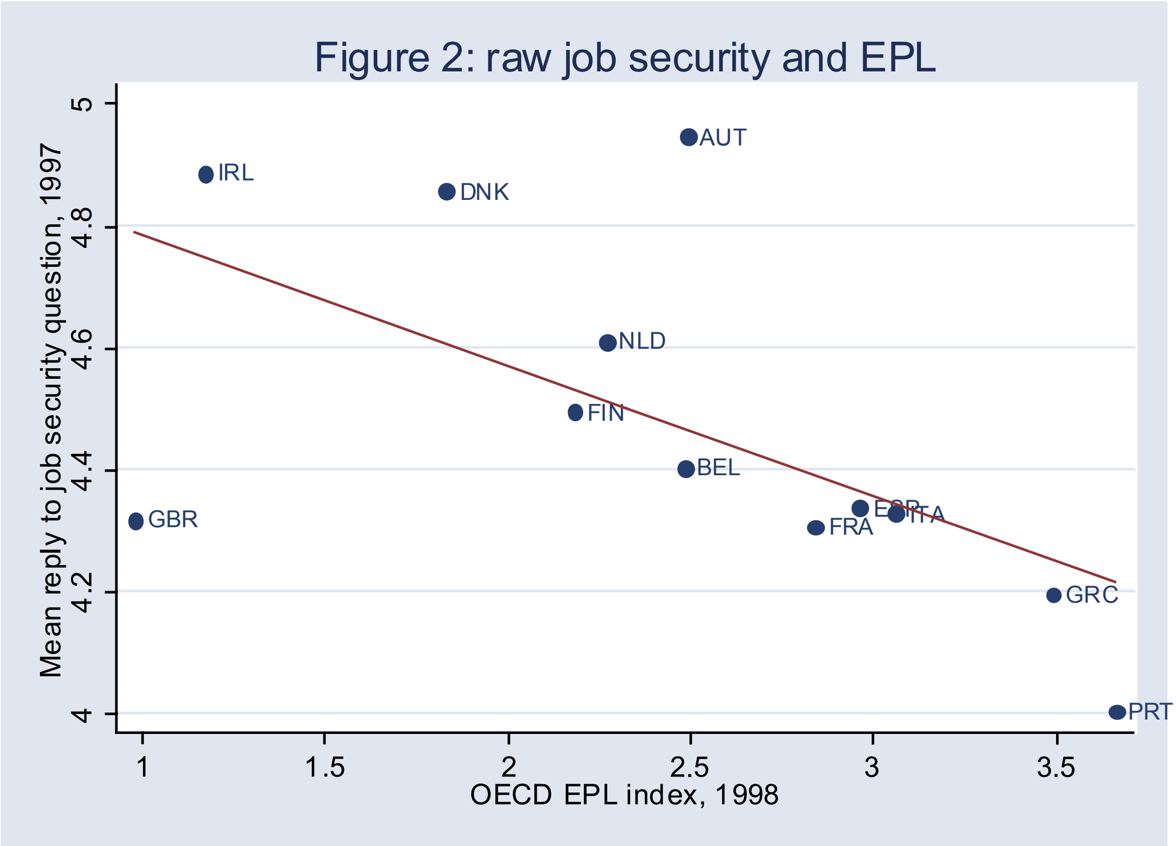


Figure 2: raw job security and EPL

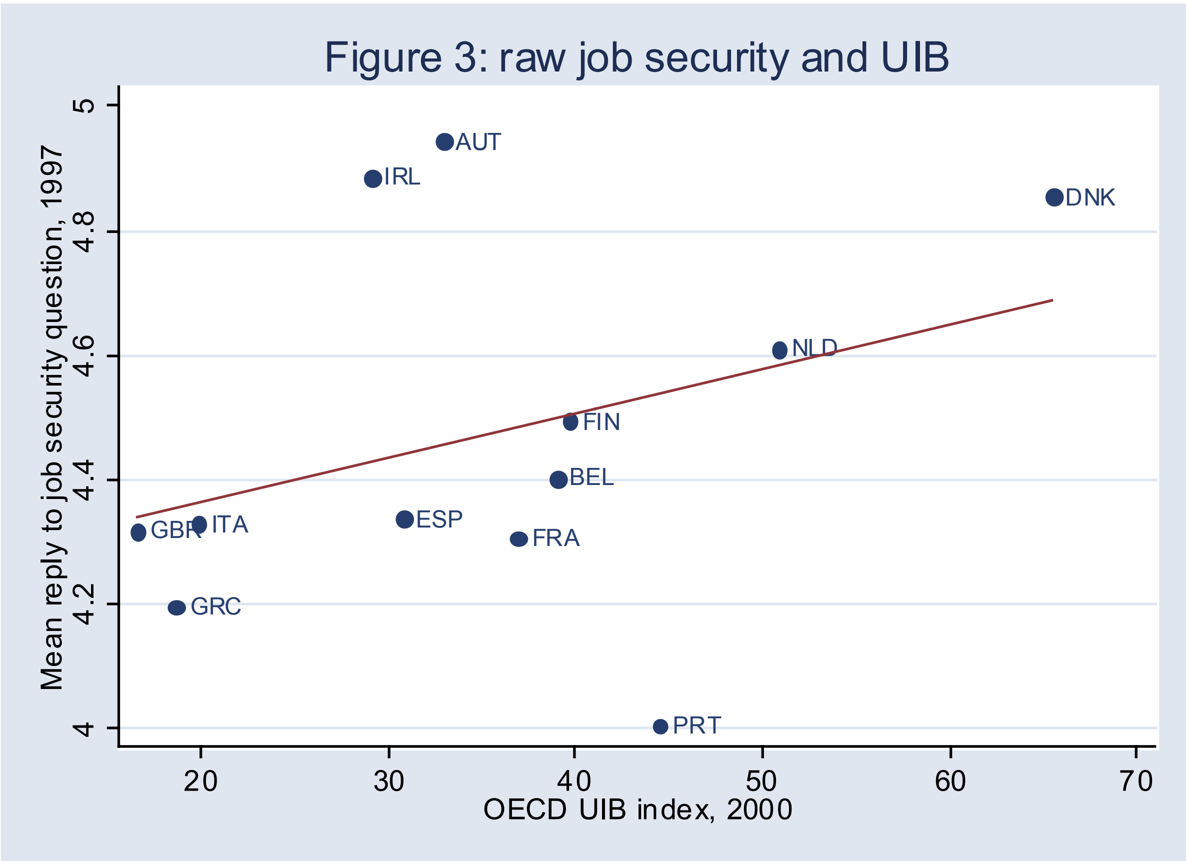
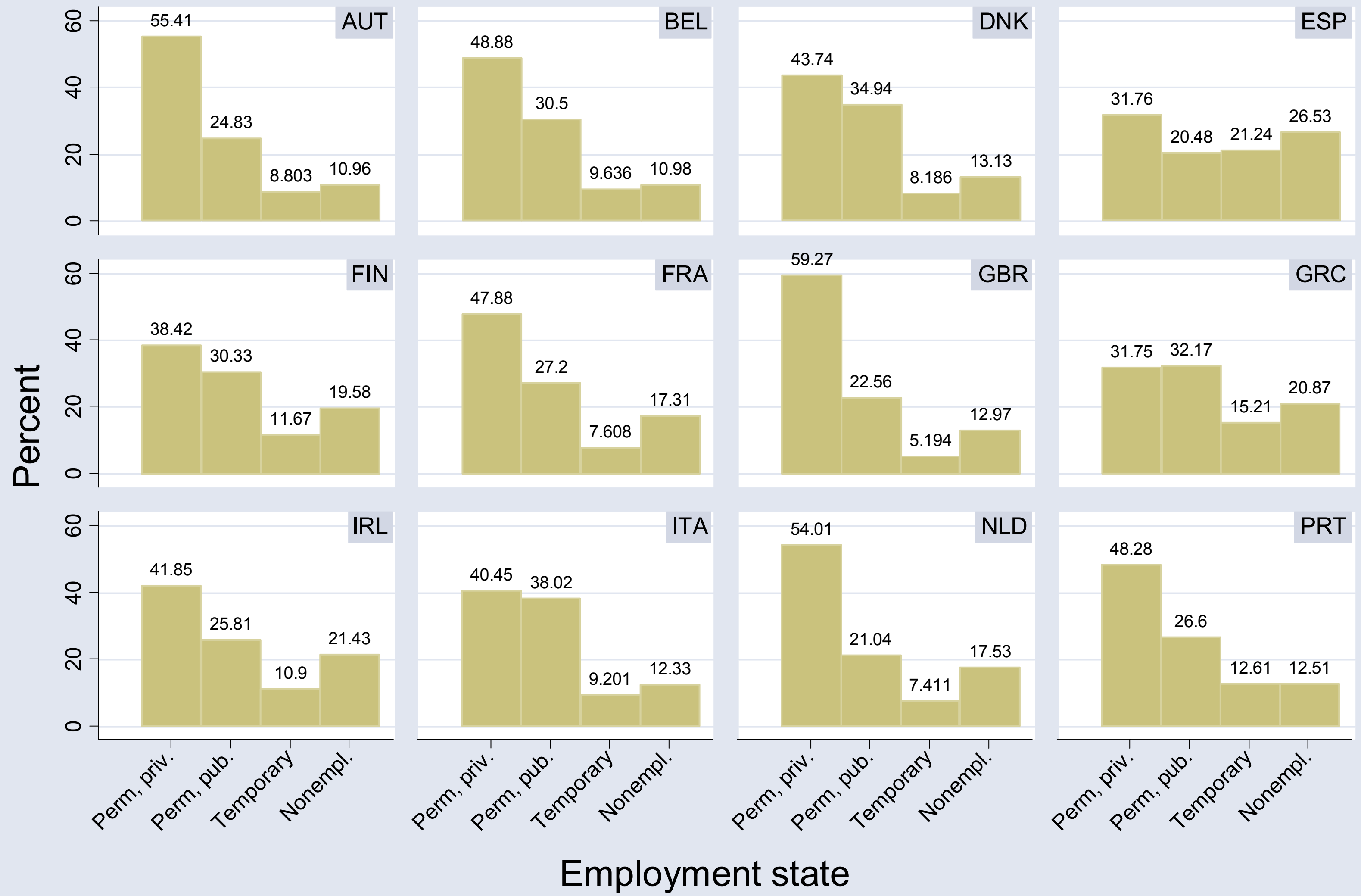


Figure 3: raw job security and UIB

Figure A1: distribution of workers across job states, 1997



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