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 the OECD summary measures of Employment Protection Legislation (EPL) strictness on one hand, and Unemployment Insurance Benefits (UIB) generosity on the other. We 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. These correlations also arise for permanent public jobs, yet in a much attenuated way, suggesting that such jobs are perceived to be by and large insulated from labor market fluctuations.

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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, on the other hand, are 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 a public sentiment 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 almighty 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 appealing

[1]See Addison and Texeira (2003), OECD (1999) or the excellent survey in Cahuc and Zylberberg (2004). One should also mention that, as it affects private decisions about job creation and destruction, EPL can obviously be thought of as serving more general purposes 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* (Communist Revolutionary League), the *Parti Communiste Français* (French

to a nontrivial fraction of the 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] recommends 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 labor market institutions choice, 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

Communist Party) and the *Parti des Travailleurs* (Workers' Party)—also contained similar proposals concerning 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

[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 thus 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 identify individual fixed effects in the perceived job security. In addition, we explicitly model worker selection into four different types of labor market status (permanent private job, permanent public job, temporary job, and nonemployment) to capture the endogeneity of worker allocation into jobs.

We have four main findings. First, job security in permanent private and temporary jobs is positively correlated with UIB generosity across countries. Second, 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, the second result suggests that job protection is not the best response to the problem—real or supposed (see OECD, 1997)—of job insecurity. Third, public sector jobs are largely considered to be the most secure, and the correlation of this security with UIB or EPL is much more tenuous than that found for other job types. This suggests that public sector jobs are perceived to be by and large insulated from labor market fluctuations. Fourth, the difference in terms of perceived job security between permanent and temporary jobs, 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 makes a first pass at examining the correlations between job security and labor market policy indicators. This section also contains a brief review of the related empirical literature. 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 the various employment states; and second, we focus specifically on the relationship between labor market institutions and job security. 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 the indication of 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]

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

"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 a variety of labor market institutions related to either Employment Protection Legislation (EPL) or to 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, originally covering fifteen EU countries. 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."*

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 (imperfectly) deal with the former by allowing for unobserved individual heterogeneity in our statistical analysis below. Concerning the latter, there is not much one can do besides keeping it in mind when interpreting the results. Specifically, our general approach will be to interpret the replies to

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 formulation differences across countries resulting from translation.

that question as proxy measures of *the workers' subjective assessment of the expected change in utility associated with a job loss multiplied by the (subjective) probability of losing their job.*[8] This interpretation has the advantage of explicitly acknowledging the "confounded issues" problem in a natural and convenient way. Yet we are fully aware that it is not immune to criticism, and the use of a probabilistic question would be a useful complement to the current analysis. Unfortunately, we know of no multi-country panel including this information. So, with that series of caveats in mind, we shall proceed.

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. As a result, 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. Our final sample consists of 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 distributions of replies to the job security question (among employed wage earners).

< Figure 1 about here. >

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 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.

[8] Formally, consider some employed worker i answering the job security question at some date t. Denote worker i's expected lifetime utility from job continuation at date t by V_{it}^{E}, that 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 that it is a measure of $q_{it} \cdot \left(V_{it}^{E} - V_{it}^{U}\right)$.

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, then we 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. 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?

< Figures 2 and 3 about here. >

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.[9]

First, looking at the vertical scales on both Figures, we obtain a job security ranking of the 12 countries represented in our sample. The basic picture seems to be that workers in "Southern" countries (Portugal, Italy, France, Spain and Greece) feel overall less secure than their counterparts

[9] 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. An important drawback of this indicator is that it fails to take account of the criteria governing eligibility for Unemployment Insurance. Since these criteria vary widely across countries, this is potentially problematic. As an alternative indicator of UIB generosity, we could use average UI expenditures per unemployed which are arguably more complete measures of UIB generosity. The reason why we choose to use average replacement rates is that there is a mechanical negative correlation between mean UI expenditures per unemployed and the unemployment rate, which in turn is likely to be negatively correlated with job security for reasons that we discuss below. This mechanical correlation can thus be suspected of causing 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 using average UI expenditures per unemployed (which are available upon request) go in the same general direction as the results that we present here.

in Northern countries (the Netherlands, Denmark and Ireland, with the most secure country being Austria). Workers in the United Kingdom also occupy a rather low position in this ranking.

Second, Figure 2 strongly suggests a negative correlation between job security and job protection: at first blush, countries with stricter EPL have workers who feel *less* secure in their jobs. Conversely, Figure 3 suggests—somewhat less strongly[10]—that countries with more generous UIB also have workers who feel *more* secure in their jobs.

Those 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 undoubtedly feel insecure (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 Accounting for observed individual heterogeneity

Personal characteristics, labor market conditions and job type. Our first step beyond a 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 personal characteristics: birth cohort and cohort squared, education (3 dummies[11]), cohabitational status, the presence of children under 15 in the household, an indicator of foreign citizenship, and an

[10] 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.

[11] 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. Yet this is the only general education variable available in the ECHP user database.

indicator of the existence of a long-term unemployment spell in the recent past.[12] We also include the 5-year average local unemployment rate as an indicator of local labor market conditions.[13]

Finally, we want to allow for the possibility that holders of different job types have fundamentally different perceptions of their job security. In this perspective we distinguish 3 different job types:[14]

- $e =$ 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 reported 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 holders of each one of the three job types defined in the previous paragraph. The sample is the initial wave (1997) of our panel.[15]

< Table 1 about here. >

Table 1 first reports a constant, which is normalized at zero for permanent employees of the private sector.[16] The constant is non statistically significant for public employees, suggesting that 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

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

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

[14] 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 very costly computationally.

[15] Results from subsequent waves are very similar.

[16] Any ceteris paribus difference in job security is subsumed in this constant. That is, we have imposed equality of the ancillary cutoff parameters of the Ordered Probit across job types. While admittedly restrictive, this protocol facilitates the comparison of job security between job types.

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 effects are non statistically significant among employees of the public sector.

Next, low-educated workers seem to feel somewhat less secure than their high-educated counterparts in all permanent job types, whereas education doesn't make any difference in terms of job security among temporary workers. This may be taken as reflecting the generally less favorable conditions of low-skilled labor markets. Yet these particular results should be taken with caution given the arguably poor quality of the education variable (see footnote 11).

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

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. But the corresponding effect on perceived job security in permanent jobs (public or private) is *positive* (albeit non significant 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 of Table 1 report the estimated effects of our policy indices. Those pertain to the main objective of this paper, which is to explore the link between job security and labor market policy. Here one observes that 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 first series of results corroborates the visual impression given by Figure 2. However, turning to the effect of EPL strictness on perceived job security, one sees that it is negative for permanent private contract holders (as Figure 1 would have suggested), essentially zero for temporary job holders, and positive for permanent employees of the public sector. This tends to mitigate the conclusion that one was tempted to draw from the negative bivariate correlation shown in Figure 1, which only survives the control for observed individual heterogeneity among holders of permanent, private sector jobs.

Discussion. While some of the regression results gathered in Table 1 are either standard or intuitive, others are puzzling. First, the fact that temporary jobs is the only job type for which education is uncorrelated with job security is difficult to explain. Given that temporary workers are objectively more exposed to the risk of losing their jobs, and given that the chances of finding a replacement job are 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 hard to explain. Again, given that temporary jobs are more exposed to the risk of job loss, one would have expected temporary workers to be most responsive (one way or the other) to differences in job protection regulations. Ultimately, there seems to be very little ceteris paribus job security differential between job types: the constant term in Table 1 is only negative and weakly significant for temporary job holders.

While there are probably many potential explanations for those results, what we shall explore in this paper is the role of unobserved individual heterogeneity. The presumption that we have is that the selection process of workers into job types may not be independent of the workers' (unobserved) general attitude toward job security. For example, workers that tend to be "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— turns up as a handicap for job search. As a result, workers that have this trait may end up in "undesirable" job states—typically, temporary jobs or nonemployment—more often than their highly self-confident counterparts. Whichever way it goes, the results gathered in Table 1 may partly be governed by selection effects. Our task in the rest of this paper will be to exploit the longitudinal dimension of the data to try to account for such selection effects.

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 (2003) 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 (2003) 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 conclusions reached by these authors is that proposals to increase UIB generosity find more support in countries offering less generous UIB and more stringent EPL altogether.

4 A statistical model of job security

In this section we present the statistical framework that we use to analyze the determinants of 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:

$$s_{it}^* = x_{it}'\beta + \sum_{e \in \mathcal{E}} \varphi_i^e \mathbf{1}_{e_{it}=e} + u_{it}, \tag{1}$$

where the notation is the following. First, x_{it} includes year dummies and date t local labor market conditions.[17] Second, e_{it} is the individual's job type—or job state—at date t. We consistently use the notation $\mathbf{1}_{e_{it}=e} = 1$ if individual i is in a job type e at date t. Again here we distinguish the 3 job types described in subsection 3.2, plus a fourth corresponding to nonemployment (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 nonemployment. We thus allow for the possibility that the effect on perceived job security of being in a particular job type may be individual-specific. Thus the various 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

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

heterogeneity variables. Our step 2 below will conduct a detailed exploration of the determinants of those random effects. For now, we let $\boldsymbol{\varphi}_i = \left(\varphi_i^{\text{ppriv}}, \varphi_i^{\text{ppub}}, \varphi_i^{\text{temp}}\right)'$ denote the vector of labor market status/person 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) 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" 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 $(x_{it}, e_{it}, \boldsymbol{\varphi}_i)$ 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:

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

where Θ denotes the entire set of parameters and $N(\cdot)$ denotes the cdf 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 $\boldsymbol{\varphi}_i$ is that it partly captures psychological traits reflecting the taste or aversion for specific employment states. For instance, individuals with very low values of the φ_i^{temp} component of $\boldsymbol{\varphi}_i$ 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 $\boldsymbol{\varphi}_i$ is likely to be correlated with the observed employment states 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 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 $\boldsymbol{\varphi}_i$ in order to capture the potential selection of specific worker types into specific job types as evoked above.[18]

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 $\{X_i, \boldsymbol{\varphi}_i, \mathcal{M}_i\}$. The contribution of individual i to the complete likelihood is a function of the parameters Θ and the complete data, $\mathcal{L}_i^c\left(\Theta; X_i, \boldsymbol{\varphi}_i, \mathcal{M}_i\right)$, which can be factored as follows:

$$\begin{aligned}\mathcal{L}_i^c\left(\Theta; X_i, \boldsymbol{\varphi}_i, \mathcal{M}_i\right) = & \overbrace{\Pr\left(\overrightarrow{s}_i | \overrightarrow{x}_i, \overrightarrow{e}_i, \boldsymbol{\varphi}_i; \Theta\right)}^{1} \\ & \times \underbrace{\left[\Pr\left(\overrightarrow{e}_i | e_{i1}, \mathcal{M}_i; \Theta\right) \times \Pr\left(e_{i1} | \mathcal{M}_i; \Theta\right)\right]}_{2} \times \underbrace{\Pr\left(\boldsymbol{\varphi}_i, \mathcal{M}_i | \Theta\right)}_{3} \times \underbrace{\Pr\left(\overrightarrow{x}_i\right)}_{4}.\end{aligned} \tag{4}$$

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. It 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

[18] 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: $\Pr\left(e_{i1} = j | \mathcal{M}_i; \Theta\right) = \pi^1_{\mathcal{M}_i}(j)$. We should also mention the existence of another possible approach (advocated by Wooldridge, 2002), which would be to condition the whole problem on the individual's initial state, e_{i1}.

the transition probabilities given by (3), for all dates $t \geq 2$, multiplied by the marginal probability of the initial state e_{i1} given by (**??**). The fourth and last term is independent of the parameters and can be ignored. Concerning this fourth term, however, one should emphasize an important assumption implicitly made in (4), which is that $\overrightarrow{x}_i$ is independent of $(\boldsymbol{\varphi}_i, \mathcal{M}_i)$. This assumption has implications for step 2 of our estimation procedure, which we will discuss at the end of subsection 4.2.

What remains to be modeled here is the third term, i.e. the joint distribution of the unobserved heterogeneity $(\boldsymbol{\varphi}_i, \mathcal{M}_i)$. 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, $\Pr(\boldsymbol{\varphi}_i, \mathcal{M}_i|\Theta)$. Here we use a finite mixture approach and assume that any 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 $(\boldsymbol{\varphi}_k, \mathcal{M}_k)$ of the various unobserved individual effects. Formally, we model $(\boldsymbol{\varphi}_i, \mathcal{M}_i)$ as

$$\boldsymbol{\varphi}_i = \sum_{k=1}^{K} \boldsymbol{\varphi}_k \times \mathbf{1}_{k_i=k}, \quad \mathcal{M}_i = \sum_{k=1}^{K} \mathcal{M}_k \times \mathbf{1}_{k_i=k}, \tag{5}$$

where k_i is the unobserved class index of individual i. The joint distribution of $(\boldsymbol{\varphi}_i, \mathcal{M}_i)$ 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:[19]

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

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

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

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

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 effect $\boldsymbol{\varphi}_i$ in our job security equation (1) picks up the impact of all permanent individual characteristics—observed or otherwise—on perceived job security in all job types. For instance, it may be the case that the subjective "cost" of holding a temporary relative to a permanent contract varies from one individual to another according to 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 $\boldsymbol{\varphi}_i$, 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 $\boldsymbol{\varphi}_i$ for each individual i in the sample. This is equivalent to constructing a predictor $\widehat{k}_i$ of the particular class k_i to which individual i belongs. First we 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:[20]

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

[20] In fact, these probabilities are by-products of the EM algorithm that we use in the estimation procedure of step 1. See Appendix A for details.

With these probabilities in hand, we define our predictor $\widehat{k}_i$ as follows:

$$\widehat{k}_i = \underset{k\in\{1,\ldots,K\}}{\arg\min} \sum_{\ell=1}^{K} \left(\Pr\left(k_i = \ell | X_i; \widehat{\Theta}\right) \sum_{e\in\mathcal{E}} \pi_\ell^\infty(e) \left(\varphi_\ell^e - \varphi_k^e\right)^2 \right), \tag{9}$$

where π_ℓ^∞ is the invariant probability distribution associated with the transition matrix $\mathcal{M}_\ell$. This latter distribution is defined over all four employment states by $\pi_\ell^{\infty\prime} \cdot \mathcal{M}_\ell = \pi_\ell^{\infty\prime}$ and measures the long-run probability of finding a member of class ℓ in each particular employment state.

The construction (9) of $\widehat{k}_i$ can be explained as follows. Suppose that we assign to class k an individual who really belongs to class ℓ. Then, each time this individual is observed in some employment state $e \in \mathcal{E}$, the squared prediction error that we are making on perceived job security is $\left(\varphi_\ell^e - \varphi_k^e\right)^2$. Since this individual really belongs to class ℓ, the (long-run) probability with which he is observed in job state e is $\pi_\ell^\infty(e)$. Therefore, the mean squared prediction error that we are making is $\sum_{e\in\mathcal{E}} \pi_\ell^\infty(e) \left(\varphi_\ell^e - \varphi_k^e\right)^2$. Equation (9) minimizes the expectation of that mean squared prediction error, given the data X_i, for each individual i in the sample.[21]

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

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

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

Explaining job security. We now turn to our statistical decomposition of perceived job security. The basic idea that we pursue is to run OLS regressions of the 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 the particular labor market in which the individual trades. Most importantly, z_i includes country-level policy indicators. We describe the exact specifications that we use below, as we comment on the estimated values of α^e.

[21] Obviously, this is not the only imaginable minimization criterion. For instance, an alternative (simpler) option would consist in minimizing $\sum_\ell \left(\Pr\left(k_i = \ell | X_i; \widehat{\Theta}\right) \|\boldsymbol{\varphi}_\ell - \boldsymbol{\varphi}_k\|^2\right)$, without taking account of the long-run probability of each employment state. The results under this alternative criterion are extremely similar to those we present in this paper.

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 how one should interpret the regression results, it maybe useful to emphasize that this method merely provides a *descriptive* decomposition of individual perceived job security $\widehat{\boldsymbol{\varphi}}_i$ 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 $\boldsymbol{\varphi}_i$ as a function of the z_i's and some unobserved heterogeneity variable, say ε_i, this function is fundamentally unidentified. We hence reiterate that our goal in this paper is to provide an intuitive description of job security data rather than to estimate a structural model of job security.

Second, the likelihood function from step 1—see equations (4) and (6)—were written using 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 $(\boldsymbol{\varphi}_i, \psi_i)$). 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 one should recall that $\overrightarrow{x}_i$ contains indicators of local (i.e. 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 the force of this criticism, we only incorporate *temporary* variations in local labor market conditions in the vector of explanatory variables $\overrightarrow{x}_i$ in such a way that $\overrightarrow{x}_i$ be orthogonal to the country dummies. (Specifically, $\overrightarrow{x}_i$ contains year dummies, and the regional unemployment rate *taken in deviation from its 5-year mean* rather than the regional unemployment rate in level—see footnote 17.) The permanent labor market conditions (as captured by the 5-year mean regional unemployment rate) are then incorporated in the z_i regressors.

Third, 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 many bootstrap replications of our step 1, which takes some time to converge. 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.

< Tables 2 to 4 about here. >

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,

[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 didn't seem to make much qualitative difference for the results.

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 jobs. One can also note that 3 out of 8 classes (number 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 jobs, and less secure in temporary than in permanent jobs. 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" chiefly concerns temporary job holders while public employees are relatively insulated. This is not entirely unexpected, but we still take it as a general indication that 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 look at 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 jobs, while those from classes 5 and 7 go to the public sector. Also, some workers (e.g., classes 4 and 8) seem to have trouble avoiding the "undesirable" employment states—namely temporary jobs and nonemployment. This suggests that the "job security" individual effect $\boldsymbol{\varphi}_i$ 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 may want 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 won't provide the most synthetic picture of worker selection. More conveniently, our results allow us to compute the selection biases defined as follows:

$$\mathcal{B}(e_1, e_2) = E\left(\varphi_i^{e_1} | e_i = e_2\right) - E\left(\varphi_i^{e_1} | e_i \neq e_2\right). \tag{12}$$

This is the gap between average job security for a job of type $e_1 \in \mathcal{E}$ as perceived by workers in employment state e_2 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_2. Equation (12) thus takes up the familiar definition of selection biases from the "treatment effects" literature.[24]

< Table 5 about here. >

The matrix $\mathcal{B}$ is shown in Table 5 for the 3 fixed effect values φ^{ppriv}, φ^{ppub} and φ_i^{temp}, and the 3 conditioning employment states "ppriv", "ppub" and "temp". The first thing to note is that the 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 first two diagonal 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: $\mathcal{B}(\text{ppriv}, \text{ppriv}) > 0$. Likewise, $\mathcal{B}(\text{ppub}, \text{ppub}) > 0$. While these conform with simple intuition, the negative sign of the third diagonal term is more puzzling. $\mathcal{B}(\text{temp}, \text{temp}) < 0$, meaning that temporary job holders tend to be *more* temporary job-averse 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 follow this discussion of selectivity biases by turning back to our central equation of interest, equation (1). We begin by analyzing the impact of labor market conditions on job security.

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

$$\mathcal{B}(e_1, e_2) = \sum_{k=1}^{K} \left(\varphi_k^{e_1} \frac{p_k \pi_k^{\infty}(e_2)}{\sum_{\ell=1}^{K} p_\ell \pi_\ell^{\infty}(e_2)} \right) - \sum_{k=1}^{K} \left(\varphi_k^{e_1} \frac{p_k \left[1 - \pi_k^{\infty}(e_2)\right]}{1 - \sum_{\ell=1}^{K} p_\ell \pi_\ell^{\infty}(e_2)} \right),$$

where $p_k \pi_k^{\infty}(e_2) / \sum_{\ell=1}^{K} p_\ell \pi_\ell^{\infty}(e_2)$ (resp. $p_k \left[1 - \pi_k^{\infty}(e_2)\right] / \left[1 - \sum_{\ell=1}^{K} p_\ell \pi_\ell^{\infty}(e_2)\right]$) is the probability of belonging to class k conditional on being in employment state e_2 (resp. in an employment state other than e_2).

Labor market conditions. The estimated coefficients on the observed time-varying covariates x_{it} (the β's from 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. Most of the covariates entering the right hand side of equation (1) potentially impact both components of perceived job security.

< Table 6 about here. >

The year dummies suggest that job security tends to follow the cycle—with 1998 to 2000 appearing to be slightly more "secure" years—with no sign of a time trend. This last point was not unexpected, given that interviewees are asked to report their feeling about job security measured on a fixed 1-6 scale. A higher-than-normal local unemployment rate tends to make workers more worried. This is unsurprising, as 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.

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 here for convenience:

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

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

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

[26] A reminder: it includes the individual's birth cohort and cohort squared, education (3 dummies), cohabitational 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.

see, our treatment of unobserved heterogeneity improves on the results from Table 1 by making them generally more clear-cut and more intuitive.

< Table 7 about here. >

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 in *all* job types (including temporary jobs) than workers with a high or intermediate level of education, while the latter two feel equally secure. Again, the orders of magnitude are similar across job types.

In contrast to the unconclusive results of Table 1, there is now evidence of foreign workers feeling 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 for public jobs is smaller in magnitude by about a third and of borderline 10% significance.

Also, there is 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 effect was present in Table 1. Interestingly, these effects vanish in public sector jobs. These latter results can be interpreted as a (remote) sign that insurance within the family plays a role in some countries: the presence of children makes job loss more costly, while the presence of a spouse who can fulfill the role of second breadwinner alleviates it.

Past experience of long-term unemployment reduces perceived job security in all types of jobs. Again, looking at point estimates, this effect is about twice as strong in permanent private jobs than in permanent public jobs, and three times as strong in temporary jobs than in permanent public jobs. Table 7 also reports a negative correlation between the average local unemployemnt rate and perceived job security in permanent private and temporary jobs (with a somewhat stronger estimated effect in the case of temporary jobs). But its effect on perceived job security in public jobs is strongly *positive*. These particular results are qualitatively similar—yet again more clear cut—to what was found in Table 1. Our tentative interpretation that workers facing adverse labor market conditions aspire to more insulated jobs thus remains.

Table 7 also reports a constant (first row), the values of which once more confirm the ranking of our three job types in terms of job security: temporary job holders feel less secure than permanent, private job holders, while holding a permanent public job makes people feel more secure. This implicit ranking of contract types now looks clear and is more intuitive than the one obtained in Table 1, a difference which may be taken as a confirmation of 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 come out either non statistically significant or with much weaker absolute values in the regression (11) for φ^{ppub} (i.e., for public jobs). Compared to private or temporary jobs, public jobs thus seem to be perceived as safe jobs, that are 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. Here we first see that EPL has a negative and significant correlation with job security, and UIB has a positive and significant correlation with job security in all types of jobs. In the end, the visual impression given by Figures 2 and 3 is thus confirmed.

Second, both of the effects are estimated as slightly larger in temporary than in permanent private jobs—although the difference is likely not significant. By contrast, the correlation between job security and EPL is smaller by a factor of three in permanent public jobs than in private and temporary jobs. Likewise, the correlation between job security and UIB is smaller by a factor of five in permanent public jobs than in other job types. Again, public sector jobs seem to be perceived as being largely insulated from the risk of job loss.

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 the highlighted 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: it does indeed protect by reducing the risk of job loss, but it also increases the cost of job loss by reducing the outflow rate from unemployment. One interpretation of the negative correlations appearing in Table 7 is that the second phenomenon dominates. 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 probably be considered preferable to EPL as an insurance tool against labor market risk.[27] What Table 7 seems to suggest is that workers, in many cases, view it this way.

6.2 Job security as a determinant of institutions

The correlations found above do not inform about the causality of the relationship. The existing "macro-labor" literature dealing with labor market policy is primarily interested in causality running from institutions to labor market outcomes, whereas the "political economy of institutions" literature considers the arrow running in the opposite direction.[28] A recurring idea in this latter

[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

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) put it, "*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 measure of the "individual gain (in terms of job security) to being an insider" as $\widehat{\varphi}_i^{\text{ppriv}} - \widehat{\varphi}_i^{\text{temp}}$ and regress it on the covariates z_i and our country-level measures of UIB and EPL. The coefficients on these latter two variables are shown in the first column 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 that this result does not depend on the fact that we are assimilating "being an insider" to "holding a permanent, private job" and "being an outsider" to "holding a temporary job". Similar and even stronger results obtain when one considers $\widehat{\varphi}_i^{\text{ppub}} - \widehat{\varphi}_i^{\text{temp}}$ or $\widehat{\varphi}_i^{\text{ppub}} - \widehat{\varphi}_i^{\text{ppriv}}$, as in the second and third columns of Table 8.

< Table 8 about here. >

6.3 Taking stock

Thus our results seem consistent with messages from both strands of literature.

Yet turning back to the first two columns of Table 7, one sees that even holders of permanent jobs—who can to a first approximation be considered to be "insiders" and who arguably 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 possible interpretation is that insiders 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.

[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.

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 also found, although in a much attenuated way, 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 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 of the sample. The European Panel (ECHP) is a common-questionnaire panel of household and individual data gathered by EUROSTAT, originally covering fifteen EU countries over eight waves (1994 to 2001).[31] However, due to missing data, we are only able to use twelve of the fifteen countries, and five of the eight waves: the satisfaction with job security question 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] Finally, Austria only joined the ECHP in 1995 (wave 2) and Finland in 1996 (wave 3).

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.

As a result, 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 observed in self-employment in at least one year during our observation window were left out of the sample, and so were individuals consistently reporting that they were nonparticipants. 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. Here we just add that this question comes second in a series of 7 job satisfaction questions that are asked in a row to the interviewee in the "employment" section of the questionnaire. It turns out that the responses to many of those satisfaction questions are highly correlated.

Job mobility. Individual i's employment state at date t is denoted by e_{it}. As described in the main text, we distinguish four possible employment states for any individual: employed in the private sector under a permanent contract (e_{it} = ppriv); employed in the public sector under a permanent contract (e_{it} = ppub); employed under a temporary contract (e_{it} = temp);[33] and nonemployed (e_{it} = none). Note that all the

[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. Those latter problems could be fixed to a large extent by bringing the 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 because of the scarcity of transitions from, e.g. permanent, public jobs to temporary, private jobs.

information used to construct the state indicator e_{it} is reported by the individual. In particular, the definition of what a "temporary contract" is can be somewhat arbitrary and vary from one individual to the other.

< 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 can use as starting values to start over at the beginning of the E-step.

In our application, using equations (2) to (6), individual i's contribution to the complete log-likelihood writes out 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 $\Theta=\left(\beta,\boldsymbol{\varphi},\boldsymbol{\tau},\boldsymbol{\pi}^1,\mathbf{M},\mathbf{p}\right)$, 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 $(\beta,\boldsymbol{\varphi},\boldsymbol{\tau})$—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 $\left(\boldsymbol{\pi}^1,\mathbf{M},\mathbf{p}\right)$—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 is stopped.[34] Else we start over at the E-step using $\widehat{\Theta}|_{\Theta_0}$ as a new initial guess.

Finally, the parsimony criterion that we use 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 is in two steps. In a first step, we 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. Then, in a second step, we predict a value $\widehat{k}_i$ of k_i for each i following the protocol presented in subsection 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, however, lies in that (as we already argue in the main text) the conditional distribution $\Pr(k_i|z_i; \Theta)$ is not nonparametrically identified. Yet the estimates obtained in the single-step method are a priori sensitive to the particular parametric assumption that one makes about the form of $\Pr(k_i|z_i; \Theta)$. This problem is circumvented by the two-step method, where we are only imposing 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 once 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 that one re-runs 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 $e_{it} = \dots$			
		ppriv	**ppub**	**temp**	**none**
Past	**ppriv**	92.41	1.75	2.89	2.94
state	**ppub**	4.44	92.36	1.30	1.90
$e_{it-1} = \dots$	**temp**	32.21	6.57	47.67	13.55
	none	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 $e = \dots$** ... perm. private	... perm. 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	−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})
$\boldsymbol{\varphi}_1$	0	−3.76	−3.73
	(ref.)	(.124)	(.099)
$\boldsymbol{\varphi}_2$	−0.98	−0.68	−1.16
	(.067)	(.093)	(.130)
$\boldsymbol{\varphi}_3$	−2.10	−2.11	−3.02
	(.087)	(.111)	(.084)
$\boldsymbol{\varphi}_4$	−2.11	0.03	−2.72
	(.076)	(.080)	(.112)
$\boldsymbol{\varphi}_5$	−3.29	−1.10	−4.06
	(.076)	(.110)	(.093)
$\boldsymbol{\varphi}_6$	−2.11	−2.12	−3.71
	(.074)	(.103)	(.092)
$\boldsymbol{\varphi}_7$	−0.47	−0.48	−0.97
	(.116)	(.105)	(.083)
$\boldsymbol{\varphi}_8$	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 1.

Table 4: Invariant job state distributions

Job state	Invariant distributions ($\pi_k^{\infty\prime} \cdot \mathcal{M}_k = \pi_k^{\infty\prime}$):[1,2]		
$\begin{pmatrix} \text{Perm., priv.} \\ \text{Perm., publ.} \\ \text{Temporary} \\ \text{Nonemp.} \end{pmatrix}$	$\pi_1^{\infty} = \begin{pmatrix} .822 \\ .023 \\ .124 \\ .026 \end{pmatrix}$	$\pi_2^{\infty} = \begin{pmatrix} .905 \\ .013 \\ .031 \\ .051 \end{pmatrix}$	$\pi_3^{\infty} = \begin{pmatrix} .469 \\ .353 \\ .068 \\ .109 \end{pmatrix}$
$\begin{pmatrix} \text{Perm., priv.} \\ \text{Perm., publ.} \\ \text{Temporary} \\ \text{Nonemp.} \end{pmatrix}$	$\pi_4^{\infty} = \begin{pmatrix} .299 \\ .498 \\ .137 \\ .066 \end{pmatrix}$	$\pi_5^{\infty} = \begin{pmatrix} .266 \\ .628 \\ .035 \\ .070 \end{pmatrix}$	$\pi_6^{\infty} = \begin{pmatrix} .855 \\ .090 \\ .019 \\ .036 \end{pmatrix}$
$\begin{pmatrix} \text{Perm., priv.} \\ \text{Perm., publ.} \\ \text{Temporary} \\ \text{Nonemp.} \end{pmatrix}$	$\pi_7^{\infty} = \begin{pmatrix} .084 \\ .738 \\ .072 \\ .106 \end{pmatrix}$	$\pi_8^{\infty} = \begin{pmatrix} .468 \\ .333 \\ .010 \\ .190 \end{pmatrix}$	$\boldsymbol{\pi}_{\text{mean}}^{\infty} = \begin{pmatrix} \mathbf{.604} \\ \mathbf{.273} \\ \mathbf{.053} \\ \mathbf{.070} \end{pmatrix}^{3}$

Notes: [1]Subscripts designate classes.
[2]Standard errors are not reported (available upon 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

Matrix $\mathcal{B}$	Conditioning state (e_2)		
	ppriv	ppub	temp
$E\left(\varphi^{\text{ppriv}} \vert e = e_2\right) - E\left(\varphi^{\text{ppriv}} \vert e \neq e_2\right)$	0.345	−0.518	−0.091
$E\left(\varphi^{\text{ppub}} \vert e = e_2\right) - E\left(\varphi^{\text{ppub}} \vert e \neq e_2\right)$	−0.638	0.587	−0.027
$E\left(\varphi^{\text{temp}} \vert e = e_2\right) - E\left(\varphi^{\text{temp}} \vert e \neq e_2\right)$	0.105	−0.258	−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	−0.373 (.168)	0.539 (.161)	−0.743 (.203)
Age (/10)	−0.518 (.093)	−0.786 (.090)	−0.675 (.112)
Age-squared (/100)	0.063 (.012)	0.104 (.012)	0.084 (.015)
High education	0 (ref.)	0 (ref.)	0 (ref.)
Intermediate education	0.002 (.025)	−0.033 (.024)	0.025 (.030)
Low education	−0.147 (.026)	−0.163 (.025)	−0.173 (.031)
Foreign	−0.247 (.076)	−0.132 (.073)	−0.217 (.091)
Couple	0.059 (.028)	0.021 (.027)	0.101 (.034)
Has children	−0.037 (.023)	0.009 (.022)	−0.051 (.028)
Past unemployment	−0.174 (.040)	−0.083 (.039)	−0.257 (.049)
Mean local unempl. rate[2]	−0.599 (.226)	0.711 (.216)	−0.844 (.272)
EPL	−0.158 (.014)	−0.057 (.013)	−0.185 (.017)
UIB	0.914 (.074)	0.407 (.071)	1.103 (.089)

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

Table 8: Job security and policy

Dependent variable:	$\widehat{\varphi}^{\text{ppriv}} - \widehat{\varphi}^{\text{temp}}$	$\widehat{\varphi}^{\text{ppub}} - \widehat{\varphi}^{\text{temp}}$	$\widehat{\varphi}^{\text{ppub}} - \widehat{\varphi}^{\text{ppriv}}$
EPL	0.027 (.007)	0.128 (.013)	0.101 (.014)
UIB	−0.189 (.068)	−0.696 (.068)	−0.507 (.075)

Note: Standard errors in parentheses.