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Wages, implicit contracts, and the business cycle: Evidence from a European panel ☆

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HIGHLIGHTS

► We study the cyclical co-movement of hours and wages in Europe.
► Their behavior is consistent with the presence of implicit insurance contracts.
► The nature of the contracts depends on the country's labor market institutions.
► The elasticity of labor supply is much smaller compared to the U.S. labor market.

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ABSTRACT

We study the joint behavior of hours and wages over the business cycle in a unique panel of 13 European countries, and document significant history dependence in wages. Workers who experience favorable market conditions during their tenure on the job have higher wages, and work fewer labor hours. Unobserved differences in productivity, such as varying job quality, or match-specific productivity are not likely to explain this variation. The results instead point to the importance of contractual arrangements in wage determination. In economies with decentralized bargaining practices, such arrangements resemble self-enforcing insurance contracts with one-sided commitment (by the employer). On the other hand, in countries with strong unions and centralized wage bargaining, wage behavior is better approximated by full-commitment insurance contracts. The co-movement of hours and wages further confirms a contractual framework with variable worker hours. Despite the strong prevalence of contracts in Europe, however, the elasticity of labor supply is considerably smaller compared to the U.S. labor market.

1. Introduction

Understanding the process of wage formation is central to evaluating the empirical performance of equilibrium models of the labor market. Most models in macroeconomics assume that the behavior of wages over the business cycle reflects movements in contemporaneous productivity (Kydland and Prescott, 1982; Long and Plosser, 1983; Mortensen and Pissarides, 1994). Nevertheless, evidence from micro-data shows that wages display significant history dependence, a feature that is considered to be at odds with the spot market model of wages. In particular, using the data from the U.S. labor market, Beaudry and DiNardo (1991) (BD hereon), find that a worker's wage depends crucially on the history of economic conditions during his tenure on the job, and that the contemporaneous conditions are irrelevant when history dependence is accounted for.[1] This pattern is considered to be in line with a contractual market where employers and workers are engaged in long-term contracts to insure workers against temporary drops in their productivity (Baily, 1974; Azariadis, 1975).[2]

The BD methodology has been applied in a relatively limited number of country-specific studies. Nevertheless, the evidence outside the North American markets remains mixed and inconclusive. McDonald and Worswick (1999) and Green and Townsend, (2010) confirm the relevance of implicit contracts for Canada and Macis (2006) for Italy, while studies for Australia (Seltzer and Merrett, 2000), Britain (Devereux and Hart, 2007), Finland (Kilponen and Santavita, 2010) and Germany (Vilhuber, 1999) argue otherwise. This is puzzling, because, in many

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[1] Grant (2003) confirms these findings in an extended sample for the U.S. See also Kudlyak (2010) for a similar result.
[2] See Rosen (1985) for a survey of the implicit contract literature.

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ways, the European style labor markets are more conducive to contractual arrangements compared to the U.S. The tighter labor regulations, along with the strong presence of labor unions in some countries, facilitate the enforceability of wage contracts (Hogan, 2001). In addition, the limited turnover relative to the U.S. increases the expected duration of an employment relationship, raising the potential welfare gains from long-term contracts.

In this paper, by exploiting a unique data set for a panel of European countries, we are able to get a better idea of the relevance of implicit wage contracts, and address a variety of issues recently raised against the existing studies. In particular, we conduct a more robust test of contractual effects based on the behavior of wages over the business cycle, and reinforce our findings by examining whether the joint behavior of hours and wages is consistent with the theory of implicit insurance contracts. The upshot of our paper is that we find a significant role for contracts in Europe.

We begin our analysis by applying the standard BD methodology to Europe. We regress a worker's wage rate on the current unemployment rate in his country of residence, the initial unemployment rate at the start of the job, and the lowest unemployment rate since the start of the job, controlling for individual productivity characteristics. Since, in contractual markets, wages contain information on the economic conditions when the contract was (re)negotiated, the relevance of past unemployment rates has been considered as evidence for implicit contracts. We tackle two issues that cloud this interpretation.

First, such history dependence could also reflect unobserved, match-specific productivity differences among jobs that are created during different phases of the business cycle (Okun, 1973). For instance, Bowlus (1995) shows that jobs that start during recessions have shorter expected duration, indicating a lower match quality. More recently, Hagedorn and Manovskii (2010) argue that the empirical patterns in Beaudry and DiNardo (1991) are also consistent with a spot market model with heterogeneous match quality and endogenous quits. We address this concern in two ways. First, we directly control for measures of match quality. Second, we focus on the wage growth of workers who do not change jobs between two consecutive periods. If the match-specific productivity is time-invariant, this eliminates the potential biases that are due to the changing composition of job quality over the business cycle (Bellou and Kaymak, 2011).

Second, in a single country study, the estimates of contracting effects are identified by the co-variation over time of wages and the aggregate unemployment rate. This can lead to multicollinearity problems in BD type regressions unless the data spans a sufficiently long time period with several cohorts of workers, a feature missing in most studies. Our data spans a relatively short period of time, but by employing a panel of countries, we are able to pursue a more robust identification strategy, which relies on the variation in wages and unemployment within a country over time for several countries. We are also able to control for arbitrary time effects in wages, which is infeasible at a country level.

Next, we seek evidence for the presence of contracts by studying the behavior of hours over the business cycle, and provide an estimate of the elasticity of labor supply for Europe.[3] When hours are part of the negotiation, the welfare maximizing contract aims to provide a constant utility flow to the worker by varying wage payments along with working hours, hence, leisure. Our test relies on the cross-sectional variation in wages conditional on productivity, a distinct characteristic of contractual markets. For instance, a worker who was hired in an expansion, makes more than an identical worker who was hired in the recession, because the former was insured against a possible downfall in productivity prior to the recession. Since both workers have the same productivity, a higher wage rate constitutes a pure income effect, and, therefore must be accompanied by a fall in hours if leisure is a normal good (Beaudry and DiNardo, 1995).

In a spot market model, on the other hand, any variation in the wage rate reflects productivity differences, and, therefore, leads to both income and substitution effects. This mitigates the correlation between hours and wages, and can generate a positive correlation instead. We test for this prediction by projecting the changes in hours on changes in the contractual variation in wages using a two-step procedure.

Our main finding is that implicit contracts overall play a significant role in the determination of wages in Europe. This is robust to unobserved differences in match quality, which appear to be important components of wages but do not affect our main conclusions regarding the importance of past labor market conditions on wages. Despite the statistical strength of our results, we find that the elasticity of wages to past unemployment rates is dampened relative to the reported estimates for the U.S.

Moreover, we show that hours worked respond negatively to differences in wages that arise due to disparities in contractual terms, a result consistent with the presence of risk-sharing agreements. The estimated elasticity of labor supply in our preferred specification is −0.19, which is slightly lower than the estimates in Beaudry and DiNardo (1995) for the U.S., but consistent with the existing estimates for Europe.

Finally, the data provides useful information on whether the employment relationship of a respondent is covered by an explicit contract. By combining this unique feature with variation in the institutional factors across European countries, we are also able to provide valuable insights for the interaction of such factors with insurance contracts, and test whether the theory of implicit contracts is a good description of wages governed by explicit contracts in the data.

While our tests indicate strong contractual effects for workers who are reportedly on a long-term employment contract, we find that the nature of the contract depends crucially on the strength of labor unions and the centralization of the bargaining process in the country. In countries with strong unions and centralized bargaining practices, wages depend significantly on the unemployment rate at the start of the job in support of a contractual model with full commitment by both workers and employers. In countries with decentralized bargaining, the findings are consistent with contracts that are binding on the employer, but not on the worker. Not surprisingly, however, for workers who have short-term, casual jobs without an explicit contract, wages are much more responsive to the current economic conditions, and past market conditions have no effects on their wages.

The empirical specifications are introduced in the next section. Section 3 describes the data. Section 4 presents the results on wages and Section 5 examines the cyclical behavior of hours. Section 6 concludes.

2. Empirical tests of implicit contracts

Following Beaudry and DiNardo (1991), we estimate a wage regression that nests three different models of wage determination. The general economic conditions are approximated by the unemployment rate in each country. The real wage at each period is projected on the contemporaneous unemployment rate (spot market model), the unemployment rate at the start of the employment relationship (full-commitment risk-sharing model), and the minimum unemployment rate since the start of the job (one-sided lack of commitment with worker mobility). Our specification is:

$$\ln w_{ict} = \beta_1 U_{ct} + \beta_2 U_{ic}^0 + \beta_3 U_{ict}^{min} + \beta_4 X_{ict} + \theta_t + \mu_c + \delta_i + \varepsilon_{ict} \tag{1}$$

where i, c, t index individual, country and time respectively; $\ln w_{ict}$ is the current log-real wage rate for individual i; U_{ct} is the contemporaneous

[3] Although the implications of implicit contracts for wage movements have been extensively studied, much less attention has been spent on the implications for hours with the notable exceptions of Abowd and Card (1987) and Beaudry and DiNardo (1995).

unemployment rate in country c; U^{0}_{ic} is the unemployment rate when worker i started his job in country c; U^{min}_{ict} is the minimum unemployment rate between the start year of the job and year t. X is a vector of covariates. θ_t is a vector of indicators for survey years, μ_c for country and δ_i for individual fixed effects. ε_{ict} is a random error term.

The spot market model is consistent with $\beta_1<0$ and $\beta_2=\beta_3=0$, i.e. wages depend only on contemporaneous conditions. In a full-commitment risk-sharing model, the optimal wage contract features a fixed wage, which equals the expected productivity of the worker at the time the contract was signed. Since both parties fully commit, wage is not renegotiated. This is consistent with $\beta_2<0$ and $\beta_1=\beta_3=0$. When the contract is binding on the firm, but the worker can change employers, the optimal contract features a downward rigid wage with raises whenever the economic conditions improve (Harris and Holmstrom, 1982). Then, wages reflect the best economic conditions that the worker ever experienced. This can be tested by $\beta_3<0$ and $\beta_1=\beta_2=0$.

The additional control variables are cubic polynomials in age and tenure, indicators for the industry classification of the job, indicators for the region within the country. We use the cubic polynomial in age as a proxy for the labor market experience.[4] Tenure captures the accumulation of firm-specific human capital. Because the minimum unemployment rate is correlated with tenure by construction, including a cubic polynomial in tenure is essential in order to account for possible nonlinearities in the effect of tenure on wages that could be inaccurately absorbed by the minimum unemployment rate.[5]

The set of industry dummies corrects for varying industry composition at different phases of the business cycle (Okun, 1973). Individual fixed effects control for time-invariant productivity characteristics. They also capture the shifts in the composition of unobserved worker characteristics over the business cycle (Bils, 1985; Solon et al., 1994). If, for instance, low productivity workers are hired primarily during expansions, then wages are negatively correlated with the economic conditions at the start of the job, which would tend to attenuate the coefficient on the initial unemployment rate.

The existing studies typically estimate Eq. (1) with a linear or a quadratic time trend. Since we observe workers with different employment histories in various countries over time, we can control for time-specific intercepts without jeopardizing the identification of the contemporaneous unemployment rate in a country. In addition, we include country dummies to capture cross-country variations in unemployment rates and wages, so that the identification of the parameters of interest (β_1,β_2,β_3) essentially stems from variation in unemployment and wages *within a country over time* for several countries.

2.1. Selection and job quality

The underlying assumption in the estimation of Eq. (1) is that the error term is orthogonal to different unemployment measures. Nonetheless it is possible that ε contains unobserved match-specific components that are correlated with past economic conditions. For instance, Bowlus (1995) finds that total job duration, a proxy often used for match quality, is negatively related to the unemployment rate when the job started. We address this concern in two ways. First we include proxy variables for job quality as controls in Eq. (1). Second, we estimate Eq. (1) in differences for workers who do not switch jobs. Taking differences for job stayers eliminates any persistent components, including permanent match-specific effects. Letting $\varepsilon_{ict}=m_{ij}+v_{ict}$, and taking differences yields the following specification.

$$\Delta_k \ln w_{ict} = \beta_1 \Delta_k U_{ct} + \beta_3 \Delta_k U^{min}_{ict} + \beta_4 \Delta_k X_{ict} + \Delta_k \theta_t + \Delta_k v_{ict}, \tag{2}$$

where $\Delta_k x_t = x_t - x_{t-k}$, and $t-k$ denotes the time of the last survey.

The specifications above do not exhaust all possible contractual designs. For instance, contracts which enslave the worker, but are non-binding on the firm are excluded from the analysis. These contracts predict wages that are rigid upwards, with occasional wage cuts whenever the employer finds a better worker. Wages, in this case, can be summarized by the maximum unemployment rate since the start of the job. Also not captured in Eq. (1) are arrangements when neither the employer nor the worker can credibly commit to the employment relationship. In this case, the optimal contract foresees sluggish wage adjustments including wage cuts as well as increases whenever the cyclical shocks are large (Thomas and Worrall, 1988). Such contracts cannot be captured by any extremum moments, such as the minimum or the maximum unemployment rate.[6] Similarly, Eq. (2) only permits a test of the spot market model against a contractual model with one-sided commitment since the initial unemployment rate drops from Eq. (1) when taking differences. We relax these restrictions when we study the co-movement of hours and wages below.

2.2. Implicit contracts and hours worked

When contracting arrangements specify variable working hours along with wages, the data on hours can be used to test for implicit contracts as well. Such contracts achieve allocative efficiency by setting the marginal rate of substitution between consumption and leisure to the marginal product. The wage rate, however, deviates temporarily from the marginal product to insure the worker against fluctuations in productivity. Furthermore, workers that are hired at different times can be paid at different rates conditional on productivity. For instance, a worker who was hired in an expansion, makes more than a worker who was hired in a recession, because the former was insured against a possible downfall in productivity prior to the recession. Since both workers have the same productivity, a higher wage rate constitutes a pure income effect, and, hence must be accompanied by a fall in hours if leisure is a normal good (Beaudry and DiNardo, 1995). This is the empirical test we wish to conduct.

In a spot market model, on the other hand, wages equal marginal product at all times. Consequently, any variation in the wage rate necessarily represents a variation in productivity, and, therefore, generates both an income and a substitution effect. This mitigates the correlation between hours and wages, and can generate a positive correlation instead.

The presence of a negative correlation between hours and wages can be tested by regressing the change in hours worked on the wage growth:

$$\Delta_k h_{ict} = \alpha_1 \Delta_k \ln w_{ict} + \alpha_2 \Delta X_{ict} + \nu_{ict}. \tag{3}$$

In contractual models, fluctuations in the wage conditional on productivity generate pure income effects. Therefore α_1 denotes the income elasticity of labor supply. The difficulty in estimating Eq. (3) is determining whether the fluctuations in wage growth are actually

[4] The ECHP does not report experience nor years of education, which prevents the use of experience in our regressions. We were able to construct a measure of potential experience for a small subset of the sample. The regressions were robust to replacing potential experience with age. Both sets of estimates had somewhat lower precision due to smaller sample size.

[5] Gertler and Trigari (2009) argue that inadequately controlling for job tenure could generate a negative coefficient for the minimum unemployment in a spot market when the tenure profiles are concave. We include cubic polynomials in tenure and age, but the results below are robust to including higher order polynomials.

[6] Despite the lack of a theoretical underpinning, it has become common practice to include the maximum unemployment rate in Eq. (1). While the maximum unemployment may capture some of the variation in wages, the coefficients do not have a meaningful interpretation.

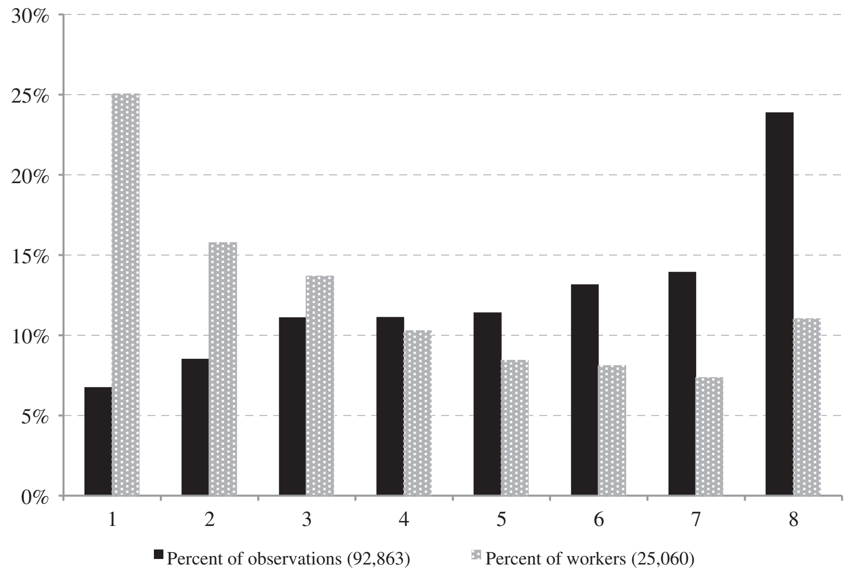

Fig. 1. Number of observations per worker.

independent of productivity, as predicted by the theory of implicit contracts. If good measures of worker productivity were available, one could estimate Eq. (3) by including these measures as control variables. When such information is not available, an alternative is to use a set of instruments for wage growth that are uncorrelated with productivity growth. Since wages display strong history dependence in contractual markets, indicators of past labor market conditions provide a natural set of instruments. For instance, initial unemployment rate in a market with full commitment, or the minimum unemployment rate in a market with one-sided commitment would be valid instruments. Since hours are otherwise allocated efficiently, i.e. they depend on contemporaneous productivity, indicators of past economic performance do not have a direct impact on hours growth.

Since the commitment structure of the market is not known ex ante, a complication remains in deciding which indicators of history are pertinent for wage growth. We avoid this by taking a more general approach here. Following Bellou and Kaymak (2011), we use a full interaction of job start year and lagged year indicators, $I(t_0 \times t-k)$, to capture the contractual variation in wages. This is parsimonious since each indicator captures the entire history of conditions from period t_0 to $t-k$, when the wage rate was last observed. This specification, therefore, encompasses all possible types of contracts, including, for instance, arrangements that are not fully binding on workers or employers. The theory of implicit contracts implies that those with lower wages in year $t-k$, due to unfavorable economic conditions between t_0 and $t-k$, have, on average, a larger wage growth between $t-k$ and t. This prediction of the model for wage growth is used to test for implicit contracts in Bellou and Kaymak (2011).[7]

To control for changes in contemporaneous conditions and to ensure that the contractual variation in wage growth is identified only by prior labor market conditions, we include a full set of lagged year and current year interactions, $I(t-k \times t)$ in the main regression. This prevents measurement issues that might arise from using the unemployment rate as a proxy for economic conditions. Other control variables are differences in cubic polynomials of age and tenure, indicators for industry, country, region within the country, and a country specific indicator for job switchers.

3. The European Community Household Panel

The data come from the 1994–2001 waves of the European Community Household Survey (ECHP). ECHP is a harmonized longitudinal survey of a set of European countries. Households and individuals are interviewed on an annual basis and provide a wide range of information regarding their labor market conditions. Although it is a relatively short panel (33 quarters), respondents provide retrospective information on their jobs. For the first wave of individuals, who are subsequently followed in later years, the earliest reported job-starting date is 1981. Therefore, we effectively exploit variation within and across countries in unemployment between 1981 and 2001 to explain the observed wages in 1994–2001.

For the baseline analysis, we restrict the sample to men of ages 21 to 64 who work full-time (30+ hours a week), in the private sector. In an auxiliary analysis we separately examine women who satisfy the same selection criteria. The appendix provides a more detailed description of the data and lists further sample restrictions. The selection criteria are chosen so that the sample is comparable to those used in similar studies. In total, we observe approximately 25,000 males who contribute 92,800 person-year observations. The distribution of observations is shown in Fig. 1. The sample contains a total of 37,018 employer–employee matches. Fig. 2 shows the total number of jobs held per worker. 56% of the workers in the sample hold 1 job throughout the sample period. The remaining 44% of the workers switch jobs at least once between 1994 and 2001. Fig. 3 shows the empirical distribution of job duration. Over 60% of the jobs last less than 4 years, however note that the observed duration is truncated since many jobs are ongoing when the sample ends in 2001. Table 1 reports the summary statistics for the core variables used in the study.

Fig. 4 depicts the trajectory of unemployment rates since 1981 for each of the countries comprising our sample. This figure essentially describes the variation we exploit for the identification of the baseline parameters. As is evident, there is considerable variability in unemployment both within and across countries through time. In 1981, the first observation year in the sample, the unemployment rate varied from as low as 2.5% in Austria to as high as 14% in Spain. Moreover, the cyclicality of unemployment varied significantly from one country to another with Portugal, for instance, experiencing multiple distinct recessionary and

[7] Our specification is slightly different than Beaudry and DiNardo (1995), who use start year indicators and changes in start year indicators. While the results here are similar to those using their specification, a full set of interactions provides a more robust identification strategy and is more efficient.

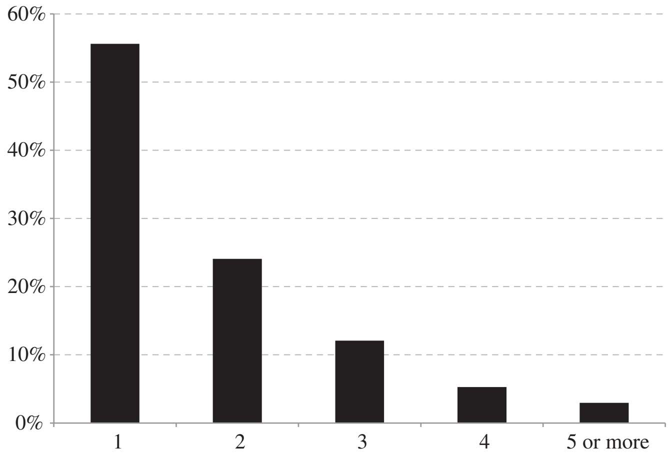

Fig. 2. Total number of jobs per worker. Note — The figure shows the fraction of workers by the total number of jobs held during the sample period. Data is taken from the European Community Household Panel (1994–2001). Sample includes men of ages 21 to 64 who work for full time in the private sector.

expansionary periods and the Netherlands having a consistently falling unemployment rate throughout the sample period.

4. Contracts and wages in Europe

Table 2 presents the main results of the paper. The coefficients are multiplied by 100 to facilitate the interpretation of the baseline effects. All specifications control for individual fixed effects, cubic polynomials in age and tenure, indicators for industry, survey years, countries and regions within a country. The first column projects wages only on the contemporaneous unemployment rate. Real wages display a clear procyclical pattern. A one percentage point increase in the national unemployment rate leads to a 1.7% fall in the real wage rate. This is very close in magnitude to the existing estimates for the U.S. labor market. Bils (1985), for instance, reports an estimate of 1.6%.

The importance of current economic conditions is considerably reduced when indicators of past labor market conditions are introduced in the following columns. The effect is most dramatic when we introduce the minimum unemployment rate since the beginning of the job, where the coefficient of the contemporaneous unemployment rate declines by more than half to 0.70%. On the other hand, the initial

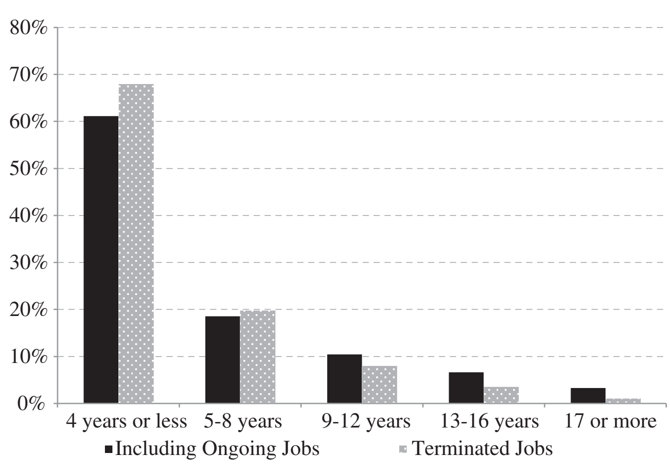

Fig. 3. Distribution of job duration. Note — The figure shows the percent of jobs with the indicated duration. There are a total of 37,018 jobs, of which 14,340 end during our sample period. Data is taken from the European Community Household Panel (1994–2001). Sample includes men of ages 21 to 64 who work for full time in the private sector.

Table 1
Descriptive statistics.

	Men	Women
Number of workers	25,060	16,403
Age	36.0 (9.9)	35.0 (9.8)
Tenure	4.9 (4.6)	4.6 (4.4)
Income (log)	2.3 (0.6)	2.0 (0.6)
Weekly hours	43.6 (7.9)	40.0 (6.0)
Number of jobs	37,018	16,374
Number of observations	92,863	50,990

Note — Standard deviations are reported in parentheses. Data come from the European Community Household Panel (1994–2001). Sample includes full-time, private sector workers of ages 21 to 64.

unemployment rate and the minimum unemployment rate have significant predictive power for current wages. The coefficients are −1.0% and −1.3%, when included separately.

In the last column, we let the three theories compete by simply including all indicators of economic conditions in the same specification. All three measures are statistically significant. A one percentage point higher unemployment rate when the job started is associated with 0.6% lower wages. This effect is additionally amplified if the worker has experienced more favorable conditions during her career on the job. A percentage point increase in the minimum unemployment rate leads to a 0.8% reduction in the current wage. Conditional on past labor market conditions, a one-percentage point decline in the current unemployment rate is associated with an approximately 0.8% increase in wages.

Overall, the results in Table 2 suggest that contracts, following the BD interpretation, are important and quantitatively at least as relevant as the spot market in the determination of wages. Although contracts play an important role in wage determination in Europe, this effect is not nearly as large as in the U.S. For instance, Beaudry and DiNardo (1991) and Grant (2003) provide estimates in the neighborhood of −2.5% for the minimum unemployment rate.

4.1. Explicit contractual arrangements

In this subsection, we investigate whether the behavior of wages for workers with explicit job contracts is consistent with the theory of implicit insurance contracts. To this end, we look at two sets of workers: those who are reportedly covered by a permanent or a fixed term contract with a duration of one year or more, and those who do not have a wage contract. The latter group contains mainly workers who undertake seasonal or casual jobs without an explicit contractual arrangement, and constitutes less than 5% of all workers.

Table 3 presents the results. The first column shows the benchmark estimation results for male workers without an explicit contract. The current unemployment rate has a sizable coefficient of −3.6%, whereas past labor market conditions are not statistically relevant. This is consistent with the spot market being the main determinant of wages of workers under no explicit contracts, though it should be noted that only a small group of them belongs to this category.

Turning to workers under explicit contracts, the results closely mirror earlier findings. Past labor market conditions play a significant role in wage determination. Wages also adjust to current economic conditions, however, suggesting that contracts are more flexible than the simple insurance contracts with one-sided commitment.

The results based on the sample of covered workers may be affected by the underlying heterogeneity among countries regarding their intensity of unionism and the degree of centralization of the wage determination process. Theoretically, the impact of unions on the presence of implicit contracts could go either way. On one hand, it has been argued that unions may constitute an enforcing mechanism of

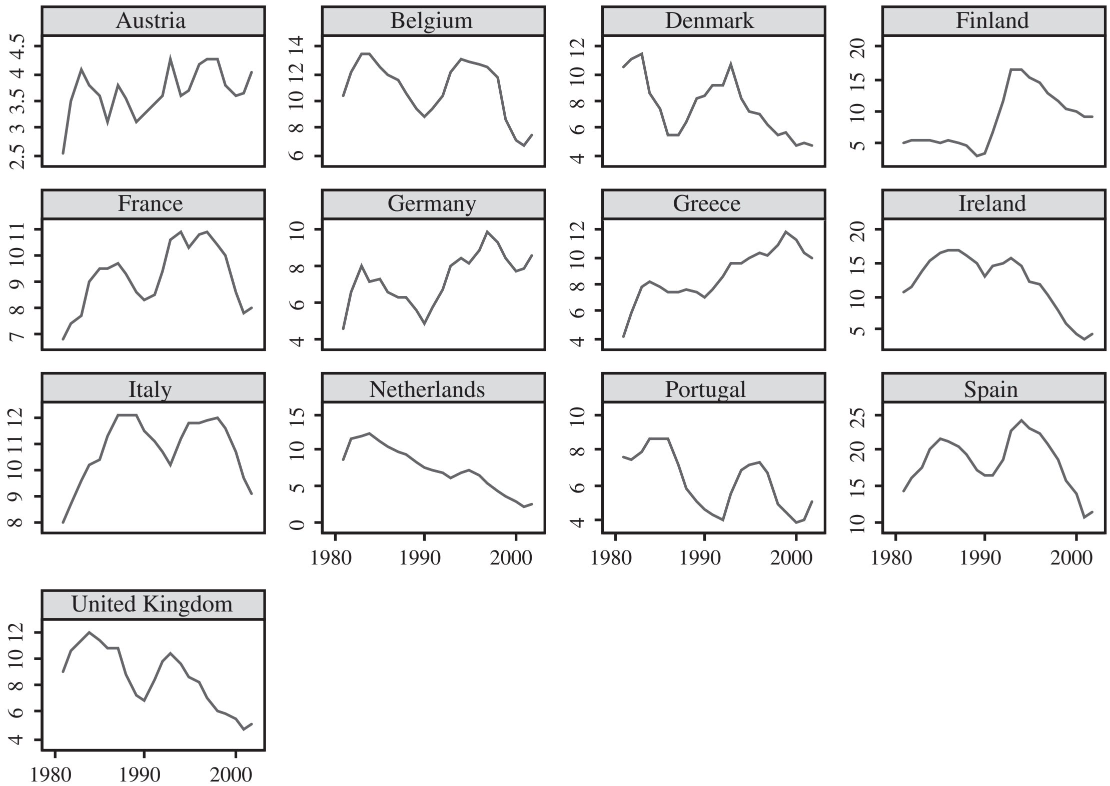

Fig. 4. The civilian unemployment rate in Europe: 1981–2002.
Source: OECD database.

existing contractual arrangements (Hogan, 2001). In that sense, one might anticipate that in countries where unions are more prevalent, implicit contracting should be stronger. Grant (2003) using the U.S. data finds that implicit contracts are indeed more prevalent in the unionized sectors.

Nevertheless, the structure of unions in the U.S. is considerably different compared to Europe. In particular, the difference between union membership and union coverage can be substantial in Europe, whereas it is negligible in the U.S.[8] Furthermore, the bargaining process in the U.S. is mostly decentralized, perhaps with the exception of the auto industry. On the contrary, in Europe, it is common to see industry-wide collective agreements. Both of these factors can influence how wage contracts are written.

To see this, we divide the sample of countries in two categories: those with a centralized collective bargaining system and those with a more decentralized wage setting mechanism.[9] For a country to belong in the first category, the union coverage rate should be at least 50% and the country's score on the level of wage setting should be at least 0.5 over 1.[10] In other words, a country should simultaneously display a relatively high degree of unionization and centralization in the wage setting process. In our sample, these countries are Belgium, Denmark, Finland, Italy and the Netherlands. Next, for each category, we re-estimate the specification in Eq. (1).

The last two columns in Table 3 show the estimation results. In the set of countries with a highly centralized collective bargaining system, the current unemployment rate does not have an effect on wages. The initial unemployment rate, on the other hand, has a significant effect on current wages. Workers that start their jobs when the unemployment rate is elevated by 1 percentage point, accept 1% lower wage rates. The minimum unemployment rate, by contrast, does not play an important role in wage determination.

That union wages do not display spot market behavior is not surprising considering that union contracts are usually not state-dependent, and they cover several years at a time. Our findings indicate that the wage rate for new workers is not rigid, but reflects the conditions at the time of the hire. The observed wage pattern is consistent with full commitment contracts, which may not be too far from reality, since most workers in this category have reportedly permanent job contracts. It is, however, at odds with Tayloristic contracting models that impose rigidity on the wages of new hires, a feature that is crucial in amplifying the volatility of employment over the business cycle (Gertler and Trigari, 2009). Our results rather support the models of the labor market that allow for flexibility of the hiring

Table 2
Real wages and unemployment.

Dependent variable: $\ln w_{ijt}$				
U_t	-1.69^{**} (0.17)	-1.30^{**} (0.17)	-0.70^{**} (0.20)	-0.84^{**} (0.20)
U_{t_0}		-1.00^{**} (0.21)		-0.60^{**} (0.25)
U_t^{min}			-1.31^{**} (0.22)	-0.81^{**} (0.26)
Obs.	92,863	92,863	92,863	92,863

Note — U_t is contemporanous unemployment rate, U_{t_0} is the unemployment rate at the start of the job and U_{min} is the minimum unemployment rates since the start of the job. All specifications also control for individual fixed effects, cubic polynomials in tenure and age, indicators for industry, survey years, countries and regions within a country. Data comes from the European Community Household Panel (1994–2001). Sample includes full-time, private sector workers of ages 21 to 64. Standard errors are clustered by individual.
** Statistically significant at 1%.

[8] Union coverage is the fraction of workers whose wages are governed by a collective agreement.
[9] Unfortunately, the ECHP does not report the union status of a worker.
[10] Both the union coverage statistics and the index of the centralization of wage bargaining are provided by Wallerstein and Western (2008) (Table 1, pg. 215, and Table 3, pg. 224). Authors use the data from various sources. See references therein.

Table 3
Real wages and unemployment for workers with explicit contracts.

Dependent variable: lnw_{ijt}				
		Permanent/fixed term contracts		
	Casual/seasonal jobs	All	Collective bargaining	Decentralized bargaining
U_t	-3.60^{*}	-0.86^{**}	-0.22	-0.74^{**}
	(1.42)	(0.23)	(0.34)	(0.27)
U_{t_0}	-1.67	-0.65^{**}	-1.03^{*}	-0.40
	(1.22)	(0.29)	(0.47)	(0.34)
U_t^{min}	2.05	-0.64^{**}	0.84	-1.12^{**}
	(1.93)	(0.29)	(0.44)	(0.34)
Obs.	4541	71,260	22,964	48,296

Note — U_t is contemporanous unemployment rate, U_{t_0} is the unemployment rate at the start of the job and U_{min} is the minimum unemployment rate since the start of the job. All specifications also control for individual fixed effects, cubic polynomials in tenure and age, indicators for industry, survey years, countries and regions within a country. Data is taken from the European Community Household Panel (1994–2001). Sample includes men of ages 21 to 64 who work full-time in the private sector. Standard errors are clustered by individual. Countries with centralized bargaining systems are Belgium, Denmark, Finland, Italy and the Netherlands.
* Statistically significant at 5%.
** Statistically significant at 1%.

Table 4
Real wages and cyclical match quality.

Dependent variable: lnw_{ijt}				
U_t	-0.84^{**}	-0.91^{**}	-0.99^{**}	-1.00^{**}
	(0.20)	(0.20)	(0.19)	(0.19)
U_{t_0}	-0.60^{*}	-0.58^{*}	-0.54	-0.67^{*}
	(0.25)	(0.25)	(0.29)	(0.28)
U_t^{min}	-0.81^{**}	-0.66^{*}	-0.70^{**}	-0.61^{*}
	(0.26)	(0.26)	(0.26)	(0.26)
Job duration		0.79^{**}		0.98^{**}
		(0.22)		(0.32)
$\sum_{t_0}^{t_f} U_t/100$			-0.06	0.11
			(0.06)	(0.07)
Obs.	92,863	92,863	92,863	92,863

Note — U_t is contemporaneous unemployment rate, U_{t_0} is the unemployment rate at the start of the job and U_{min} is the minimum unemployment rates since the start of the job. Match quality is measured by the total duration of the job and the sum of unemployment rates during the entire duration of the job. All specifications also control for individual fixed effects, cubic polynomials in tenure and age, indicators for industry, survey years, countries and regions within a country. Data comes from the European Community Household Panel (1994–2001). Sample includes men of ages 21–64 who work full-time in the private sector. Standard errors are clustered by individual.
* Statistically significant at 5%.
** Statistically significant at 1%.

wage followed by limited cyclicality for wage growth on-the-job (Pissarides, 2009).

The subsequent wage increases once a worker is hired are independent of the general economic conditions. This implies that wages are not renegotiated in response to changes in the outside options of individual workers. This is expected since wage increases are likely negotiated across the board, independently of when a worker was hired. This completely separates the start year effects from wage growth within a job spell and from current year effects, if any.

The last column reports our findings for the set of countries with decentralized bargaining. Wages in these countries respond most to the minimum unemployment rate, then to the current unemployment rate, but not the initial unemployment rate. We consider these results to be indicative of the presence of contractual arrangements, that resemble one-sided insurance contracts. These results are in line with the findings of Green and Townsend (2010) using Canadian data. In particular, the authors show that while a contracting model with renegotiations is a good description of wages in the non-union sector, a full commitment insurance model is a better description of wage dynamics in the union sector.

4.2. Cyclical composition of job quality

The main conclusion from the analysis performed to this point is that both spot and implicit contract models characterize the movement of real wages over the business cycle in Europe. However, the baseline estimates could be biased because they do not account for the potentially confounding effect of match-specific productivity. As suggested recently by Hagedorn and Manovskii (2010), the observed importance of past labor market conditions could be due to cyclical selection of unobserved match quality. In particular, if wage offers are procyclical, then when times are good and jobs are plenty, only the best job matches survive as workers in weaker matches leave their jobs in pursuit of better employment opportunities. Since differences in match quality are not directly observable to the researcher, they are captured by indicators of past economic conditions, hence the history of dependence in wages.

To control for cyclical changes in the composition of jobs, we supplement our baseline specification with two measures of match quality: the total job duration of a job (or completed tenure), and the average unemployment rate over the duration of the match. The first measure captures the idea that jobs that last longer reflect high productivity matches. The second reflects the idea that such selection is more stringent during expansions, when vacancies are plenty. A job that survived an expansion is, on average, more productive than a job that only survived a recession. The two measures reflect the difference between the quantity and the quality of the time survived by the match. Our measures differ from those used by Hagedorn and Manovskii (2010) who use the log of the sum of aggregate labor market tightness instead of the unemployment rate. Our choice was driven by the lack of reliable measures of vacancies for all of the countries and years in our sample. Note, however, that despite this limitation, this exercise has some important implications. The match quality proxy used by Hagedorn and Manovskii (2010) can be decomposed into summation of two terms: the log of the total job duration and the log of average market tightness. These two terms are to an extent separately captured by the inclusion of our two match quality measures. Therefore our specification allows gauging whether history dependence is sensitive to allowing these measures to have a separate effect on wage determination.

Table 4 presents the results. The duration of a job is a strong predictor of the wage. A job that eventually lasts an additional year is associated with an average of 0.8–1.0% higher wage rate during the entire duration of the job. The average unemployment rate, on the other hand, does not do a good job of capturing job quality. The inclusion of job duration in the regression decreases the coefficient on the minimum unemployment rate from 0.8% to 0.6%. This is consistent with the finding that jobs that start during a recession are, in general, shorter (Bowlus, 1995). More importantly, however, the fact that the minimum unemployment rate remains strongly significant despite the inclusion of the match quality proxies, casts doubt on the conjecture that history dependence is entirely driven by cyclical selection in unobserved match quality.[11]

The additional variables we use are imperfect measures of match quality. Our alternative strategy to control for changes in average job quality over the business cycle is to run the benchmark regression in differences using only the workers who do not change jobs (Table 5). This method automatically eliminates any time-invariant match characteristic. It is, therefore, immune to inference problems that

[11] Total job duration is truncated from above for most of the jobs in our sample since they are not completed. We have also estimated our results with the sample of jobs that were completed. The findings are similar, if not more favorable for contracting effects. The minimum unemployment rate is persistently significant albeit all estimates have higher standard errors due to the considerable reduction in sample size.

Table 5
Real wage growth and unemployment.

Dependent variable: $\Delta_k w_t$		
$\Delta_k U_t$	−0.69*	−0.37
	(0.24)	(0.25)
$\Delta_k U_{min}$		−0.83*
		(0.33)
N	54,899	54,899

Note — $\Delta_k U_t$ and $\Delta_k U_{min}$ are the changes in the contemporaneous unemployment rate and the minimum unemployment rate since the start of the job. All specifications also control for differences in cubic polynomials in tenure and age, indicators for industry, countries and regions within a country. Data is taken from the European Community Household Panel (1994–2001). Sample includes men of ages 21 to 64 who work full-time in the private sector. Standard errors are clustered by individual.
* Statistically significant at 5%.

may arise due to measurement errors in vacancies and a potential multicollinearity between the proxies for match quality, and the minimum unemployment rate.

When we regress real wage growth on the change in the contemporaneous unemployment rate alone, we find that wages of job stayers are slightly procyclical. A one percentage point increase in the unemployment rate leads to a 0.7% reduction in wages. This is lower than the coefficient of −1.7% in the first column of Table 2, consistent with the hypothesis that the cyclicality of wages stems mostly from wages of new hires. Nevertheless, when we include the change in the minimum unemployment rate, the elasticity of wages with respect to the contemporaneous unemployment rate decreases to −0.4%, and becomes statistically insignificant. Meanwhile, the coefficient on the change in the minimum unemployment rate is −0.8%, virtually equal to the estimate in the last column of Table 2. This implies that the history dependence that we detected in the benchmark specification is not an artifact of cyclical selection of match quality.

4.3. Gender differences in contractual arrangements

The main analysis thus far focused on a sample of males and showed that history dependence is a significant determinant of the wages of workers covered by explicit contractual arrangements even after accounting for unobserved match-specific productivity. In this section, we explore whether the main conclusions hold for female workers as well. To this end, we focus on women who report that they are in an explicit contractual relationship with their employer and we replicate the baseline specification (Eq. (1)) with and without the inclusion of match quality proxies.[12]

The results are presented in Table 6. Column 1 shows the estimates from our baseline specification. The estimates suggest that history dependence is also a prominent feature of wage dynamics among women. A one percentage point higher unemployment rate at the start of the job reduces the wage rate by 1.5% on average. On the other hand, the minimum unemployment rate has no predictive power on current wages. Conditional on past labor market conditions, a one-percentage point decline in the contemporaneous unemployment rate is associated with an approximately 0.88% increase in wages, an effect which is of almost identical magnitude to that for males. These results suggest that, while the cyclical behavior of wages for men is better described by a one-sided lack of commitment model non-binding on the worker side, female wages are consistent with full commitment insurance contracts.

In column 2 we include the total job duration and the average unemployment rate over the job spell as match quality proxies. The results reinforce our contractual interpretation. While total job duration is an important component of wages, its inclusion does not minimize

[12] Wages of women with casual work arrangements or short-term contracts do not show any systematic behavior over the cycle.

Table 6
Real wages and unemployment: women.

Dependent variable: $\ln w_{ijt}$		
U_t	−0.88**	−1.14**
	(0.30)	(0.24)
U_{t_0}	−1.50**	−0.90*
	(0.46)	(0.19)
U_t^{min}	−0.35	−0.19
	(0.53)	(0.48)
Job duration		1.34*
		(0.52)
$\sum_{t_0}^{t_f} U_t/100$		0.07
		(0.11)
Obs.	38,116	38,116

Note — U_t is contemporaneous unemployment rate, U_{t_0} is the unemployment rate at the start of the job and U_{min} is the minimum unemployment rates since the start of the job. Match quality is measured by the total duration of the job and the sum of unemployment rates during the entire duration of the job. All specifications also control for individual fixed effects, cubic polynomials in tenure and age, indicators for industry, survey years, countries and regions within a country. Data come from the European Community Household Panel (1994–2001). Sample includes women of ages 21–64 who work full-time in the private sector with an explicit job contract that lasts more than a year. Standard errors are clustered by individual.
* Statistically significant at 5%.
** Statistically significant at 1%.

the significance of economic conditions at the start of the job. Whereas the impact of initial unemployment rate on current wages slightly declines in absolute value, it still remains a significant predictor of wages. When we estimate the main specification in differences, as in Eq. (2), we find that the change in the current unemployment rate is associated with a coefficient of −0.84 (standard error of 0.36) while the coefficient for the change in the minimum unemployment rate is statistically insignificant. Hence, these results corroborate the finding that the initial market conditions are more important for women than men, and that cyclical updating of wages through renegotiations is not as common for women.

5. Are hours consistent with implicit contracts?

In this section, we turn our focus to hours worked instead of the wage rate. The results from the first step of the TSLS estimation of Eq. (3) show significant history dependence in wage growth. The Wald test of the instruments yields a statistic of 49.1, which rules out concerns for weak identification. The estimate of the elasticity of labor supply for male workers with explicit contracts is −0.27 with a standard error of 0.06 (Table 7). This is higher than most of the existing estimates for European males, but comparable to some of the earlier estimates of the labor supply elasticities that exploit the peculiarities of the tax structure in Britain (Blundell and Walker, 1986; Blundell et al., 1988) and the Netherlands (van Soest et al., 1990).[13] The comparable elasticity estimate for the U.S., as provided in Beaudry and DiNardo (1995), is −0.34.

The benchmark estimate includes a country-specific switcher dummy to control for possible changes in job quality. Since the dataset is sufficiently large, we also estimated the model using job stayers only. Focusing on changes within job spells controls for changes in the cyclical composition of match-specific productivity characteristics that are time-invariant. This specification yields an elasticity of −0.19, with a standard error of 0.05, which indicates that hours are more responsive to wage adjustments for switchers.[14]

[13] See Table 2 in Blundell and MaCurdy (1999) for a list of estimates in the literature.
[14] Since the data contains multiple countries, we also estimated the model using interactions of the start year and lagged year with the country of residence as instruments for wage growth. The elasticity in this case is identified by the variation in wage growth that is dependent on the country-specific labor market history. The estimates are very similar with slightly lower standard errors.

Table 7
Income elasticity of labor supply.

Sample	Elasticity (Std. error)	Observations
Males		
All workers	−0.27 (0.06)**	51,980
Job stayers	−0.19 (0.05)**	51,669
Females		
All workers	−0.11 (0.03)**	31,135
Job stayers	−0.13 (0.03)**	27,274

Note — Table shows the estimated (income) elasticity of weekly hours with respect to the wage rate conditional on productivity. Data is taken from the European Community Household Panel (1994–2001). Sample includes full-time workers of ages 21 to 64 in the private sector. Standard errors are clustered by individual.
** Statistically significant at 1%.

Overall, the estimates show a significant negative correlation between contractual hours and wages. Nonetheless, the magnitude of the elasticity is small, especially compared to the US. To see the economic significance of this estimate, consider a two-percentage point fall in the unemployment as the economy exits a recessionary period. This drop, according to the estimates of Table 2 (column 4), is related to a 1.62% increase in the wage of a worker hired during the recession relative to a new hire. Suppose now that the wage increase is not associated with a productivity boost. Then, based on the preferred elasticity estimate, the 1.62% increase in the wage rate will induce an approximate 0.32% reduction in the weekly hours worked. Given the average of 43.6 hours per week in our sample, this amounts to a reduction of work time by 7 hours annually for an employee working 50 weeks per year. This is less than a regular workday of extra vacation time. The equivalent effect of a two-percentage point fall in unemployment for the U.S. economy would be an extra 3 to 5 days of vacation, which is more sizable.

The second panel in Table 7 reports the estimates of the income elasticity for women. Since the findings in Section 4.3 suggest that wages of female workers depend more strongly on the initial unemployment rate, we use the indicators for the start year of the job interacted by the worker's country of residence as instruments for subsequent wage growth. The estimated elasticity of labor supply is −0.11 for the entire sample, and −0.13 for the sample of job stayers. These estimates are consistent with other estimates of the income elasticity of labor supply for women in the literature.[15]

6. Conclusion

Using longitudinal data on 13 European countries, we find that past labor market conditions are significant predictors of current wages, but the effect is smaller compared to the range of estimates available for the U.S. labor market. The nature of the history dependence is consistent with the theory of implicit contracts. Since, however, we also find that the contemporaneous unemployment rate has an independent and significant effect on wages, we cannot rule out the spot market models completely.

The analysis of workers with explicit contracts provides crucial insights to the wage determination process in Europe. In countries with strong labor unions and centralized bargaining practices, we find significant time of entry effects, consistent with full commitment insurance contracts. By contrast, wages do not respond to contemporaneous economic conditions, and wage increases do not reflect changes in workers' outside options. For economies where the wage bargaining is decentralized, we find that the wage behavior is consistent with contractual models that are binding on the employer, but not on the worker.

[15] See Tables 1–4 in Bargain et al. (2011) and Table 1.7 in Alesina et al. (2006) for a long list of studies on the income elasticity of labor supply by gender.

The cyclical behavior of hours worked indicates a negative correlation between the contractual variation in wages and weekly hours worked, which delivers additional evidence in support of the contractual essence of the European labor market.

Appendix A. Data

The data come from the 1994–2001 waves of the European Community Household Survey (ECHP). The countries included in the analysis are Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Spain and the UK. Sweden is excluded since ECHP only contains cross-sectional data on this country. Luxembourg is also excluded from the analysis since a large fraction of its workers reside outside the country, rendering the national unemployment rate an irrelevant measure of workers' outside options. The sample is restricted to workers of ages 21 to 64 who work with an employer in the private sector for at least 30 h a week. Workers with multiple jobs are excluded.

Hourly wage rate is calculated as the gross monthly wage and salary income divided by four times the usual weekly hours reported by the respondent. All wage data are converted to 2000 values using the national CPI of each country, and then converted to British pounds using PPP adjusted conversion rates provided in the ECHP country files. This ensures cross-country comparability of wages without incorporating exchange rate fluctuations over time between countries.

The respondents in the ECHP report the beginning year of their jobs back to 1981. Jobs that began before 1981 were coded into one category by the ECHP, and therefore were dropped from our sample. Job tenure is measured as the difference between the interview year and the beginning year of a worker's job. For all jobs that began after 1993, the ECHP also reported the month of the year when the job started. A job switch is identified as a change in the beginning month/year of a worker's job between two consecutive interviews. All observations with apparent inconsistencies were dropped.

Information on the region of residence is aggregated to single digits in the UK. Missing observations are assigned separate region indicators for Belgium, France, Italy and Spain except in Finland and Portugal where the missing observations appear to be from the first regions in the dataset (based on our analysis of the longitudinal nature of the missing data). Denmark, Greece, Ireland and the Netherlands are considered as a single region each.

Unemployment rates are taken from the OECD database. Current unemployment rate is the quarterly unemployment rate at the time of the interview. For calculation of the minimum unemployment rate, quarterly unemployment rate is used whenever information on the beginning month of the job was available. The annual rate is substituted otherwise.

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