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Abstract

This paper examines the joint behavior of hours and wages over the business cycle in a unique panel of 11 European countries, and documents 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 confirm the predictions of a contractual framework with variable worker hours. Despite the strong prevalence of contracts in Europe, however, we find that the elasticity of wages to past labor market conditions and the elasticity of labor supply to be considerably smaller compared to the U.S labor market.

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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 reflect 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 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 was 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) 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 (Vilhubert, 1999) argue otherwise. This is puzzling, because, in many 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

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\(^1\) Grant (2003) confirms these findings in an extended sample for U.S. See also Kudlyak (2010) for a similar result.

\(^2\) See Rosen (1985) for a survey of the implicit contracts literature.
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 best 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, 2010).

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 for Europe, 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 intertemporal labor
supply for Europe.\textsuperscript{3} When hours are part of the negotiation, the welfare maximizing con-
tract 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 subsequent recession, because the former was insured against a pos-
sible 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 pro-
ductivity differences, and, therefore, leads to both income and substitution effects. This
mitigates the correlation between hours and wages, and can generate a positive correla-
tion 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 deter-
mination 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 conclu-

\textsuperscript{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
sions 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 intertemporal elasticity of substitution in our preferred specification is -0.20, which is slightly lower than the estimates in Beaudry and DiNardo (1995) for the U.S., consistent with some of the existing estimates for Europe.

Finally, the ECHP 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, our findings are consistent with contracts that are binding on the employer, but not on the worker. Not surprisingly, our findings for workers who have short-term, casual jobs without an explicit contract, indicate that wages are much more responsive to the current economic conditions, and past markets have no effects on their wages.

In the next section, we discuss our empirical specification. Section 2 describes our data, and section 3 presents our results. Section 4 concludes.
1 Empirical Methodology

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:

$$\ln w_{ict} = \beta_1 U_{ct} + \beta_2 U_{ic}^0 + \beta_3 U_{ict}^{\min} + \beta_4 X_{ict} + \theta_t + \mu_c + \delta_i + \epsilon_{ict}$$ (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 unemployment rate in country $c$; $U_{ic}^0$ is the unemployment rate when worker $i$ started his job in country $c$; $U_{ict}^{\min}$ is the minimum unemployment rate between the start year of the job and year $t$. $X$ is a vector of covariates. $\theta_t$ is a vector of indicators for survey years, $\mu_c$ for country and $\delta_i$ for individual fixed effects. $\epsilon_{ict}$ 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 $\beta_1 = \beta_2 = 0$.

Our specification does not exhaust all possible contractual designs. For instance, we
exclude contracts which enslave the worker, but are non-binding on the firm. 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 (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 raises 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.\footnote{Despite the lack of a theoretical underpinning, it has become common practice to include the maximum unemployment rate in (1). While the maximum unemployment may capture some of the variation in wages, the coefficients do not have a meaningful interpretation.} We relax these restrictions when we study the co-movement of hours and wages below.

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. 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.

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, Barsky, and Parker, 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 (1) with a linear time trend. Since we observe
workers with different employment histories in various countries over time, we can con-
trol for time-specific intercepts without jeopardizing the identification of the contempora-
neous unemployment rate in a country. In addition, we include country dummies to cap-
ture cross-country variations in unemployment rates and wages, so that the identification
of the parameters of interest \((\beta_1, \beta_2, \beta_3)\) essentially stems from variation in unemployment
and wages within a country over time for several countries.

We also extend our analysis to test if the behavior of hours worked is consistent with
contracting arrangements that specify variable working hours along with wages. In this
case, the optimal contract aims to provide a constant utility flow to the worker by vary-
ing wage payments along with working hours, and thereby, leisure. The approach that
we employ is based on the cross-sectional covariation of contracted hours and wages
conditional on productivity. The key to our test is the variation in wages conditional on
productivity, which arises naturally in contractual markets. For instance, a worker who
was hired in an expansion, makes more than a worker who was hired in a subsequent re-
cession, 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 con-
stitutes a pure income effect, and, hence 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, wages equal marginal product at all times.
Consequently, any variation in the wage rate necessarily represents a variation in produc-
tivity, 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.

To test for the negative correlation between hours and wages, conditional on produc-
tivity, we regress the change in hours worked on wage growth:

\[
\Delta_k h_{ict} = \alpha_1 \Delta_k \ln w_{ict} + \alpha_2 \Delta X_{ict} + \nu_{ict}
\]  

(2)
When good measures of productivity are not available, estimation of (2) requires a set of instruments for wage growth that are uncorrelated with productivity growth. In contractual markets, wages display strong history dependence, and indicators for past labor market conditions are a natural set of instruments. Following Bellou and Kaymak (2010), 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. We include a full set of lagged year and current year interactions, \( I(t - k \times t) \), in the main regression 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.\(^5\) This specification encompasses different types of contracts, including, for instance, when the contracting arrangements are not fully binding on workers and employers.

2 The European Community Household Panel

The data comes 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 our analysis, we restrict the sample to men of ages 21 to 60 who work full-time (30+ hours a week), in the private sector. Multiple job holders are excluded. The appendix

\(^5\)This specification is slightly different than Beaudry and DiNardo (1995), who use start year indicators and changes in start year indicators. While the results are similar, our specification is more general, and captures more of the variation in wages.
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 individuals who contribute 92,800 person-year observations. The distribution of observations are shown in Figure 2. The sample contains a total of 37,018 employer-employee matches. Figure 3 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% workers switch jobs at least once between 1994 and 2001. Figure 4 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.

Figure 1 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 expansionary periods and the Netherlands having a consistently falling unemployment rate throughout the sample period.

3 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 of our specifications control for individual fixed effects, tenure, age, indicators for industry, survey years, countries and

\textsuperscript{6}Adequately controlling for job tenure is crucial since the minimum unemployment rate is correlated with tenure by construction, which could generate a negative coefficient for the minimum unemployment
regions within a country. The first column projects wages only on the contemporaneous unemployment rate. Real wages display a clear procyclical pattern. A one percent increase in the national unemployment rate leads to a 1.7% fall in the real wage rate. This is very close to existing estimates for the U.S labor market in magnitude. 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 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 percent 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 U.S. For instance, Beaudry and DiNardo (1991) and Grant (2003) provide estimates in the neighborhood of -2.5% for the minimum unemployment rate.

in a spot market if the tenure profiles are concave (Gertler and Trigari, 2009). We include cubic polynomials in tenure and age, but the results below are robust to including higher order polynomials.
3.1 Explicit versus implicit contracts

In this subsection, we investigate whether the behavior of wages for workers with explicit job contracts is consistent with 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.

Table 3 presents our results. The first column shows the benchmark estimation results for 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 workers are in this category.

Turning to workers under explicit contracts, the results closely mirror our 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.

We suspect that the results based on the sample of covered workers might 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 was argued that unions may constitute an enforcing mechanism of 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 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\(^7\) can be substantial in Europe, whereas it is negligible in the US. 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 first divide our sample of countries in two categories: those with a centralized collective bargaining system and those with a more decentralized wage setting mechanism.\(^8\) 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.\(^9\) 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 and Italy. Next, for each category, we re-estimate the specification in (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 1 percent elevated, 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

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\(^7\)Workers whose wages are governed by a collective agreement.

\(^8\)Unfortunately, the ECHP does not report the union status of a worker.

\(^9\)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 data from various sources. See references therein.
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).

The subsequent wage raises 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 raises are likely negotiated across the board, independently of when a worker was hired. This completely separates the start year effects from wage growth within in 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.

3.2 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 on average 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 the one 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 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. We return to this below.\textsuperscript{10}

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 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). The average unemployment rate, on the other hand, does not do a good job of capturing job quality.\textsuperscript{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

\textsuperscript{10}We estimated our model in a smaller sample of countries for which vacancy statistics were available. The coefficient estimates were similar, while the standard errors were elevated. Since we think that our alternative method below is superior to use of proxy variables, we do not report these results.

\textsuperscript{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.
jobs. This method automatically eliminates any time-invariant match characteristic. It is, therefore, immune to inference problems that may arise due to measurement errors in vacancies and a potential multicollinearity between the proxies for match quality, and the minimum unemployment rate.\footnote{Bellou and Kaymak (2010) use a similar strategy to distinguish between the spot market model and the contractual model in the US.}

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 percent 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.

In a way, this result is not surprising, since the coefficient of the change in the minimum unemployment rate is mostly identified during recessions, when job-to-job transitions are not common. The arguments raised in Hagedorn and Manovskii (2010), on the other hand, rely heavily on workers who switch into high quality jobs, which is more likely to take place during booms.

### 3.3 Are hours consistent with implicit contracts?

In this section, we turn our focus to hours worked instead of the wage rate. Although the empirical studies of contracting models using wage data are widely available, less attention has been paid to the cyclical behavior of hours within in the context of contractual
models.

The baseline model of wage contracts can be easily extended to allow for variable hours. In this case, the optimal contract will not only specify a history dependent profile for wages but also a contingent profile for hours worked (Beaudry and DiNardo, 1995). The welfare maximizing insurance contract with variable hours aims to provide a constant flow of utility subject to participation constraints of employers and workers. This involves transfers either in terms of wages or in terms of leisure. The optimal contract achieves 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.

In contractual models workers that are hired at different times can be paid at different rates conditional on productivity. This arises due to differences in contractual terms. Since the productivity is the same, a higher wage rate must be associated with decreased hours to maintain the equality between the marginal rate of substitution and the marginal product. Therefore, hours are negatively related to wages in contractual markets. This is the empirical test we wish to conduct.

Our strategy is to regress changes in hours worked on the change in the wage rate. It is essential, however, to detach the variation in wages that is coming from differences in contractual terms, and not from differences in productivity. We achieve this by using a full set of interactions of start year and current year indicators as instruments for wages. The interaction variables capture the entire history of economic conditions while on the job and, hence, all possible contractual arrangements, including those with two-sided lack of commitment.13 We control for differences in cubic polynomials of age and tenure, and indicators for industry, country, region within the country, and a country specific indicator for job switchers. To control for the contemporaneous changes in the economic environ-

13Beaudry and DiNardo (1995) employ start year dummies and differences in start year dummies as instruments. While the results are similar with their specification, using full set of interactions provides a more robust identification strategy and, hence, efficiency gains.
ment, we also include a full set of interactions of the current year and the lagged year indicators. This ensures that the variation in wage growth captured by our instruments represents truly the past labor market conditions. In a contractual market, variations in wages due to past labor market conditions constitute a pure income effect, leading to a negative correlation between hours and wages. The estimated coefficient is simply the intertemporal elasticity of labor supply.

There is significant history dependence in wage growth in Europe. The Wald test of our instruments in the first step yields a statistic of 49.1. This rules out any concerns for weak identification. We estimate the intertemporal elasticity of labor supply for workers with explicit contracts to be -0.29 (with a standard error of 0.07). This is larger than most of the existing estimates for European males. The comparable elasticity estimate for US, as provided in Beaudry and DiNardo (1995), is -0.34.

Our benchmark estimate includes a country-specific switcher dummy to control for possible changes in job quality. Since we had a sufficiently large dataset, we also estimated the model using job stayers only. By focusing on changes within job spells, we control for changes in the cyclical composition of match-specific productivity characteristics that are time-invariant. This yields an elasticity of -0.20, with a standard error of 0.07, which indicates that hours are more responsive to wage adjustments for switchers.

Overall, our estimates show a significant negative correlation between contractual hours and wages. This finding is robust across different samples and methods we consider. Our preferred estimate is -0.20. This is 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 Netherlands (van Soest, Woittiez, and Kapteyn, 1990).

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14 See Table 1 in Blundell and MaCurdy (1999) for a survey of elasticity estimates.
15 Since we have data from multiple countries, we also estimated our 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 variation in wage growth that is dependent on the country-specific labor market history. The estimates are very similar with slightly lower standard errors.
16 See Table 2 in Blundell and MaCurdy (1999) for a list of estimates in the literature.
Nonetheless, the magnitude of the elasticity is small, especially compared to 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.

4 Conclusion

Using longitudinal data on 11 European countries, we present evidence that is consistent with the presence of strong contractual effects in wages. We show 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. Since 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.

Our 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 raises do not reflect changes in workers’ outside options. For economies where the wage bargaining is decentralized, we find that the wage behav-
ior is consistent with contractual models that are binding on the employer, but not on the worker.

Our study 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 functioning of the European labor market.

References


## A Data

Data comes from the 1994-2001 waves of the European Community Household Survey (ECHP). The countries included in the analysis are Austria, Belgium, Denmark, Finland, France, Greece, Ireland, Italy, the Netherlands, Portugal, Spain and the UK. Sweden was excluded since ECHP only contains cross-sectional data on this country. Luxembourg was 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 men of ages 21 to 60 who work with an employer in the private sector for at least 30 hours 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 was 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 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 they started working. A job switch was identified as a change in the beginning month/year of a worker’s job between two consecutive interviews. All observations with apparent inconsistencies, such as when the beginning year goes back in time, were dropped.

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

Unemployment rates were taken from 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 was used whenever information on the beginning month of the job was available. The annual rate was substituted otherwise.
B Tables and Figures

Table 1: Descriptive Statistics

<table>
<thead>
<tr>
<th></th>
<th>Standard</th>
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<tr>
<td></td>
<td>Mean</td>
</tr>
<tr>
<td>Age</td>
<td>36.0</td>
</tr>
<tr>
<td>Tenure</td>
<td>4.9</td>
</tr>
<tr>
<td>Income (log)</td>
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<tr>
<td>Weekly hours</td>
<td>43.6</td>
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<tr>
<td>Number of workers</td>
<td>25,060</td>
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<tr>
<td>Number of jobs</td>
<td>37,018</td>
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<td>Number of observations</td>
<td>92,863</td>
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</table>

Note.— Data comes from the European Community Household Panel (1994 - 2001). Sample includes full-time, private sector workers of ages 21 to 64.
Table 2: Real Wages and Unemployment in Europe

<table>
<thead>
<tr>
<th>$U_t$</th>
<th>-1.69**</th>
<th>-1.30**</th>
<th>-0.70**</th>
<th>-0.84**</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(0.17)</td>
<td>(0.17)</td>
<td>(0.20)</td>
<td>(0.20)</td>
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<tr>
<td>$U_{t0}$</td>
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<td>-0.60**</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.21)</td>
<td></td>
<td>(0.25)</td>
<td></td>
</tr>
<tr>
<td>$U_{tmin}$</td>
<td>-1.31**</td>
<td></td>
<td>-0.81**</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.22)</td>
<td></td>
<td>(0.26)</td>
<td></td>
</tr>
</tbody>
</table>

Note.— $U_t$ is contemporaneous unemployment rate, $U_{t0}$ is the unemployment rate at the start of the job and $U_{tmin}$ 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 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 5% and 1%.

Table 3: Real Wages and Unemployment for Workers with Explicit Contracts

|                      | Casual/Seasonal Jobs | Permanent/Fixed Term Contracts | | |
|----------------------|----------------------|--------------------------------|----------------|
|                      |                      | All                            | Collective Bargaining | Decentralized Bargaining |
| $U_t$                | -3.60*               | -0.86**                        | -0.22                 | -0.74**                  |
|                      | (1.42)               | (0.23)                         | (0.34)                | (0.27)                   |
| $U_{t0}$             | -1.67                | -0.65**                        | -1.03 *               | -0.40                    |
|                      | (1.22)               | (0.29)                         | (0.47)                | (0.34)                   |
| $U_{tmin}$           | 2.05                 | -0.64**                        | 0.84                  | -1.12**                  |
|                      | (1.93)               | (0.29)                         | (0.44)                | (0.34)                   |

Note.— $U_t$ is contemporaneous unemployment rate, $U_{t0}$ is the unemployment rate at the start of the job and $U_{tmin}$ 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 and Italy.

*,** statistically significant at 5% and 1%.
Table 4: Real Wages and Unemployment: Controlling for Match Quality

<table>
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<tr>
<th>Dependent Variable: $\ln w_{ijt}$</th>
<th>$U_t$</th>
<th>$U_{t0}$</th>
<th>$U_{tmin}$</th>
<th>Job Duration</th>
<th>$\sum_{t0}^{t} U_t / 100$</th>
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<tbody>
<tr>
<td>$U_t$</td>
<td>-0.84**</td>
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<td>-0.99**</td>
<td>-1.00**</td>
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<td></td>
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<td>(0.20)</td>
<td>(0.19)</td>
<td>(0.19)</td>
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<tr>
<td>$U_{t0}$</td>
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<td>-0.58*</td>
<td>-0.54</td>
<td>-0.67*</td>
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<tr>
<td></td>
<td>(0.25)</td>
<td>(0.25)</td>
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<td>(0.28)</td>
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<tr>
<td>$U_{tmin}$</td>
<td>-0.81**</td>
<td>-0.66*</td>
<td>-0.70**</td>
<td>-0.61*</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.26)</td>
<td>(0.26)</td>
<td>(0.26)</td>
<td>(0.26)</td>
<td></td>
</tr>
<tr>
<td>Job Duration</td>
<td>0.79**</td>
<td></td>
<td>0.98**</td>
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</tr>
<tr>
<td></td>
<td>(0.22)</td>
<td></td>
<td>(0.32)</td>
<td></td>
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<tr>
<td>$\sum_{t0}^{t} U_t / 100$</td>
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<td>0.11</td>
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<td></td>
<td>(0.06)</td>
<td>(0.07)</td>
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Note.— $U_t$ is contemporaneous unemployment rate, $U_{t0}$ is the unemployment rate at the start of the job and $U_{tmin}$ 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% and 1%.

Table 5: Real Wage Growth and the Unemployment Rate

<table>
<thead>
<tr>
<th>Dependent Variable: $\Delta_k w_t$</th>
<th>$\Delta_k U_t$</th>
<th>$\Delta_k U_{min}$</th>
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<tr>
<td>$\Delta_k U_t$</td>
<td>-0.69*</td>
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<tr>
<td>$\Delta_k U_{min}$</td>
<td>-0.83*</td>
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<td></td>
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<tr>
<td>N</td>
<td>54,899</td>
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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.
Figure 1: The rate of unemployment in Europe: 1981 - 2002
Figure 2: Total number of valid observations per worker

Figure 3: Total number of jobs per worker
Figure 4: Empirical distribution of total job duration.

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.