THE STRUCTURE OF THE PERMANENT JOB WAGE PREMIUM: EVIDENCE FROM EUROPE

by

Lawrence M. Kahn, Cornell University, CESifo, IZA, and NCER (Queensland)

Email: LMK12@CORNELL.EDU

Postal Address:
Cornell University
258 Ives Hall
Ithaca, New York 14583 USA

Phone: +607-255-0510
Fax: +607-255-4496

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Abstract

Using longitudinal data on individuals from the European Community Household Panel (ECHP) for thirteen countries during 1995-2001, I investigate the wage premium for permanent jobs relative to temporary jobs. The countries are Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Spain, and the United Kingdom. I find that among men the wage premium for a permanent vs. temporary job is lower for older workers and native born workers; for women, the permanent job wage premium is lower for older workers and those with longer job tenure. Moreover, there is some evidence that among immigrant men, the permanent job premium is especially high for those who migrated from outside the European Union. These findings all suggest that the gain to promotion into permanent jobs is indeed higher for those with less experience in the domestic labor market. In contrast to the effects for the young and immigrants, the permanent job pay premium is slightly smaller on average for women than for men, even though on average women have less experience in the labor market than men do. It is possible that women even in permanent jobs are in segregated labor markets. But as noted, among women, the permanent job wage premium is higher for the young and those with less current tenure, suggesting that even in the female labor market, employers pay attention to experience differences.

JEL Classification: J31, J42.
Keywords: wage structure, segmented labor markets, temporary jobs.

I. Introduction

A considerable volume of economic research has been devoted over the last two decades to explaining and suggesting remedies for the stubbornly high unemployment rates in a number of European countries. Among the suggested policy remedies for reducing joblessness is the relaxation of systems of employment protection by allowing firms greater freedom to create temporary jobs. Such dual employment systems produce barriers into the protected, permanent job sector, since firms may be reluctant to create permanent jobs in the presence of high firing costs. Moreover, the bargaining power of insiders in the protected sector is likely to produce a pay gap relative to those in temporary jobs, since the firm must pay firing costs if it decides to discharge workers (Blanchard and Landier 2002; Booth, Francesconi and Frank 2002; Boeri 2011; Kahn 2007 and 2012; Stancanelli 2002). And previous research has found that the young, immigrants and women are disproportionately concentrated in temporary jobs, which are sometimes seen as part of a process leading to labor market dualism, due to the lower pay in temporary jobs and barriers to entering permanent jobs (Kahn 2007).

While previous research on temporary and permanent employment outcomes treats the temporary sector in the aggregate, some workers may still accumulate training and valuable experience in temporary jobs, even if this is less extensive than in permanent jobs. If so, then temporary jobs may themselves represent less of a dead end in the labor market than otherwise imagined. Moreover, we might expect different types of workers to experience different gains upon obtaining a permanent job, depending on their experience while employed in a temporary job. And this heterogeneity in the wage gains for permanent employment implies that the dual employment system can indirectly affect wage inequality even beyond the average pay gap between permanent and temporary jobs.

In this paper, I use European Community Household Data to investigate the premium workers command in permanent jobs relative to temporary jobs across thirteen European countries: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, the

Netherlands, Portugal, Spain, and the United Kingdom. A basic framework to understand this issue comes from Blanchard and Landier's (2002) research in which it is assumed that firms start workers in temporary jobs. Then as the expiration of the job approaches, the firm must decide whether to promote the worker into a permanent job or whether to start over with a new match in a temporary job. In the presence of higher firing costs for permanent jobs relative to temporary jobs, firms will be reluctant to make such promotions unless the economic circumstances of the firm warrant it. Once promoted, workers are able to appropriate some of the firing costs, since these raise the value of continuing the employment match once the worker is promoted. Thus, an important determinant of the wage premium in a permanent job is the value of the match relative to breaking it up and starting over with a temporary employee.

I hypothesize that before being promoted into a permanent job, inexperienced workers must receive training in the temporary job to which they have been hired. In equilibrium, their starting wages in the temporary job will be below the level of starting wages for experienced, trained workers starting a temporary job. After they have become trained, the firm may receive a productivity shock which will determine whether it will promote the workers. By this time, experienced and inexperienced workers will each be trained and thus will be treated similarly by the firm. Because of the wage discount at the beginning of the temporary job for less experienced workers, the wage gain conditional on promotion to a permanent job will be greater for them. We observe a higher incidence of permanent employment among more experienced workers because they have had more opportunities to be in firms that receive a favorable productivity shock, and the exit probability from permanent jobs is relatively low.

I test the prediction that the permanent job wage premium falls as labor market experience rises using longitudinal data from the ECHP. Taking into account individual fixed effects, I find that among men the wage premium for a permanent vs. temporary job is indeed lower for older workers and native born workers; for women, the permanent job wage premium is lower for older workers and those with longer current job tenure. Moreover, there is some evidence that among immigrant men, the permanent job premium is higher for those who migrated from outside the European Union. These findings all suggest that the gain to promotion into permanent jobs is indeed higher for those with less experience in the domestic labor market; moreover, previous findings that immigrants and the young are more likely to be in temporary jobs than the native born and older workers are consistent with the view outlined above as well (OECD 2002; Kahn 2007). In contrast to the effects for the young and immigrants, the permanent job pay premium is slightly smaller on average for women than for men, even though on average women have less experience in the labor market than men do. It is possible that women even in permanent jobs are in segregated labor markets with a different distribution of productivity shocks from those in men's jobs. But as noted, among women, the permanent job wage premium is higher for the young and those with less current tenure, suggesting that even in the female labor market, employers pay attention to experience differences. I then present evidence that among those with temporary jobs, workers with more experience and tenure earn higher wages and are also more likely to have received some formal company-sponsored or subsidized education or job training These patterns are consistent with the model of promotion into permanent jobs, since it predicts smaller permanent job wage gains for more experienced workers, due to their greater likelihood of having received training.

Analysis of the permanent job wage premium can also reveal some sources of wage inequality to the extent that temporary jobs comprise a sizeable portion of employment. To study this issue, I examine the contribution of the overall permanent job pay premium as well as heterogeneity in its level across labor force groups to wage inequality in the country with the largest incidence of temporary employment, Spain. I find that differences in this premium can explain a modest proportion of overall Spanish wage inequality, although for the other countries, the incidence of temporary jobs is too small for the pay gap to play much of a role in accounting for overall wage inequality.

II. Prior Research on the Wage Premium for Permanent Jobs

Recent research on the wage effects of permanent vs. temporary employment provides some guidance for studying its structure. For example, Stancanelli (2002) used ECHP micro data and an extensive set of controls to find hourly wage effects of permanent relative to temporary jobs across ten countries averaging 0.116 for women and 0.121 for men. Boeri (2011) used ECHP and other European microdata and found monthly wage effects for 12 of the 13 countries (i.e., all except Finland) in the current study averaging 19.3%, although his list of controls was far less extensive than Stancanelli's (2002), and his use of monthly rather than hourly earnings may have also helped lead to his larger estimate. Specifically, Boeri (2011) controlled for education and tenure, while Stancanelli (2002) controlled for these as well as age, sector, occupation, and unemployment history.

While these estimates are suggestive, they may be upward biased because workers on permanent jobs are likely to have higher levels of unmeasured productivity than workers on temporary jobs. Supporting this idea, Booth, Francesconi, and Frank (2002) used individual panel data for Britain and found that fixed effects estimates of the permanent job premium were smaller than cross-sectional estimates. For example, the cross-sectional effect for men was 0.171 log points but the fixed effects estimate was only 0.069; for women, the cross-sectional estimate was 0.144, and the fixed effects estimate was 0.109.

In earlier work (Kahn 2012), I used the ECHP to estimate the impact of permanent jobs on hourly wages using both cross-sectional and fixed effects methods across 11 European countries. The controls included age, age squared, dummy variables for low (ISCED levels 0-2) and middle levels (ISCED level 3) of schooling with high levels of schooling the omitted category (ISCED levels 5-7), the regional unemployment rate, and year dummy variables. The cross-sectional estimate was 0.128 log points (which is much closer to Stancanelli's (2002) estimates than Boeri's (2011) results), while the fixed effects estimate was only 0.026 log points, and both effects were statistically significant. Thus, these estimates of the wage effects of permanent jobs averaged across European countries range from a low of 0.03 (my fixed effects estimate) to a high of 0.21 (Boeri's 2011 estimate), with a middle range of 0.12-0.13 (my cross-sectional estimate and Stancanelli's estimates). The smaller fixed effects estimates I found and that Booth, Francesconi and Frank (2002) found suggest that an important portion of the cross-sectional estimate represents unmeasured individual heterogeneity rather than a true effect of permanent jobs.[1] The individual heterogeneity can occur both across workers in the same firm and among workers across firms, since workers differ in their unmeasured skills and firms differ in the product market shocks they are affected by.

In this paper, I use fixed effects methods to investigate the structure of the permanent job wage premium. Previous work suggests that the aggregate estimate of roughly 3% is modest, certainly compared to other factors that affect wages even in countries with highly centralized wage setting mechanisms.[2] Yet the small average effect may mask large differences across groups in the premium to getting a permanent job. An analysis of the structure of this premium can reveal differences in labor market outcomes within groups such as the young or immigrants.

III. Conceptual Framework

The basis for the empirical work to be described below comes from Blanchard and Landier's (2002) theoretical model of a labor market with both temporary and permanent jobs.

[1] While not directly comparable to these worker-level estimates, Bentolila and Dolado (1994) found for five European countries (Denmark, France, Spain, West Germany, and the United Kingdom), that manufacturing wages were negatively affected by the fraction of workers on temporary contracts, a result consistent with a pay premium for permanent employees.

[2] For example, in such countries, the standard deviation of industry wage effects tends to be much larger than this figure, as do the effects of a one standard deviation difference in educational attainment or cognitive ability (Kahn 1998; Blau and Kahn 2005).

Before discussing this framework, it is worth mentioning the case of an unregulated labor market with complete knowledge of workers' and firms' productivities. In such a world, firms offering only temporary jobs would have to pay a premium relative to jobs that offer employment security. Blanchard and Landier (2002) in effect build a more realistic setting in which the government imposes penalities on firms that fire workers and in which there are productivity shocks that cannot be perfectly anticipated. Specifically, policy typically distinguishes jobs according to the level of firing costs, with the permanent jobs of course having higher costs of termination. In this setup, all jobs start out as temporary, with low firing costs. The firm may receive a productivity shock, measured such that a higher value indicates a more favorable level of productivity, and then decides whether to promote the worker to permanent status.[3] The authors show that there is a level of the shock—the reservation level—above which the firm will promote and below which the firm will terminate the employment relationship. The model assumes a wage determination mechanism in which wages are set in a Nash bargaining framework, although as I discuss below, this assumption is not necessary. On both permanent and temporary jobs, firms and workers share the gains to continuing the match. These gains include the avoidance of firing costs. Since these are higher for permanent jobs, the model predicts a pay premium for those promoted into permanent jobs, which the firm will take into account before making the decision to promote the worker.

In what follows below, I generalize this framework to include the possibility that some workers (the "inexperienced") require training to enable them to perform permanent jobs and that this is acquired during employment in a temporary job. Suppose, realistically, that is more costly to fire someone from a permanent job than from a temporary job. Then an inexperienced

[3] An example of a productivity shock is the collapse of the housing market in 2007-8, which led to an especially sharp reduction in demand facing firms supplying that industry.

worker's wage in a temporary job will be lowered due to the costs of getting training, and the worker will still accept the job rather than be unemployed, due to the expected benefits of training. In contrast, an experienced worker starting a temporary job is already trained, so there is no need for a wage discount during the temporary job. Upon promotion, the experienced and inexperienced workers are in a similar situation, since they are now both trained. The Nash bargain after promotion will thus have the same result for both inexperienced and experienced workers. Therefore, the wage gain to promotion for an inexperienced worker will be greater than for a experienced worker. In this model, the promotion probability for the two workers in the same firm will be the same because it assumes that the inexperienced worker on a temporary job receives training before any productivity shock. However, the experienced workers will have had more chances to be promoted; we will therefore observe a higher incidence of permanent employment among experienced than inexperienced workers, since the exit probability from permanent jobs is relatively low.[4]

I illustrate these ideas using Blanchard and Landier's (2002) set up. Assume that a match begins between a firm and a worker who has already been trained. Suppose that all employment relationships begin with temporary jobs with firing costs c_0, productivity levels for trained workers y_0, and with wages (to be determined by bargaining) w_0. Then assume that productivity shocks occur with probability λ and that these shocks have cumulative distribution function F(-). Assume that at the point of the shock, the firm must decide whether to promote the worker to a permanent job with firing cost c or terminate the employment relationship.[5] Let y be the realized productivity level of the shock, and w(y) be the wage on the permanent job. Let k be the cost of

[4] As discussed further below, in the case where the inexperienced worker may not have received training before promotion, we still expect to see a larger wage gain for inexperienced than experienced workers upon promotion.

[5] Further following Blanchard and Landier (2002), assume that the firing costs represent administrative expenses rather than severance pay, which according to Lazear's (1990) analysis represent a transfer between the company and the worker and thus need not affect resource allocation.

creating a new vacancy, s be the exogenous probability of retirement, and r be the discount rate. Then Blanchard and Landier (2002) show that there will be a reservation shock level y^* above which the firm will promote the worker and below which the firm will terminate the employment relationship.

In this model, the flow return to the firm of a new job having value V_0 is:

(1) $rV_0 = (y_0 - w_0) - c_0\lambda F(y^*) + \lambda \int_{y*}^{\infty}[V(y) - V_0]dF(y)$

The flow value to the firm of a continuing job is:

(2) $rV(y) = [y - w(y)] + s[V_0 - V(y)]$

The flow value of a new temporary job to an already trained worker is:

(3) $rV_0^e = w_0 + \lambda F(y^*)(V_u - V_0^e) - sV_0^e + \lambda \int_{y*}^{\infty}\{V^e[w(y)] - V_0^e\}dF(y)$,

where V_u is the value of being unemployed and $V^e(w(y))$ is the value of a permanent job with productivity level y. Finally, the flow return to the worker of being employed in a permanent job is:

(4) $rV^e[w(y)] = w(y) - s\, V^e[w(y)]$.[6]

[6] The flow values shown in equations (1)-(4) are the sum of current period income and the expected capital gain or loss due to possible productivity shocks or exogenous retirement. As an example, equation (4) shows that the current income for a worker on a permanent jobs is wage earnings, and the expected capital loss is the probability of retirement s times the value of the permanent job.

With these value functions, assume Nash bargaining for both temporary and permanent jobs. For permanent jobs, the firm's status quo value is V_0-c, which is the value of posting a new vacancy minus the firing costs, while the worker's status quo value is V_u, the value of being unemployed. For temporary jobs, the firm's status quo value is V_0-c_0, which is the value of creating a new vacancy minus the firing cost from a temporary job, while the worker's status quo value is still V_u. The reservation productivity y* is defined implicitly as:

(5) $V(y^*) = V_0 - c_0$.

That is, the reservation productivity makes the firm indifferent between promoting the worker into a permanent job and firing the worker, paying the temporary job firing costs and announcing a new vacancy.

With symmetric Nash bargaining, we have the following conditions for the worker's value of temporary and permanent jobs:

(6) $V_0^e - V_u = c_0$ and

(7) $V^e\,[w(y)] - V_u = V(y) - V_0 + c.$ [7]

Thus, the worker's gain to promotion into a job with productivity y is:

(8) $V^e\,[w(y)] - V_0^e = V(y) - V_0 + c - c_0 = V(y) - V(y^*) + c - 2c_0$,

[7] As explained by Blanchard and Landier (2002), the left hand sides of equations (6) and (7) represent the worker's gain to, respectively, a temporary and a permanent job, while the right hand sides are the firm's corresponding gains.

and the worker's expected gain to promotion given a promotion is:

$$(9)\ E\{V^e\,[w(y)] - V_0^e | y \geq y^*\} = E\,[V(y) - V(y^*) + c - 2c_0 |\, y \geq y^*].$$

Using this framework, we can now contrast the gains to promotion for experienced vs. inexperienced workers. Under my assumptions about timing, the only difference between hiring an experienced vs. an inexperienced worker is that the firm must pay training costs for the latter at the beginning of the temporary job. Denote these costs as H. Thus, the net productivity for hiring an inexperienced worker is (y_0-H). In this setup, we must modify the firm's and worker's values of a temporary job relative to hiring an experienced worker. Under competition, the value to the firm of hiring an inexperienced or an experienced worker for the temporary job must be the same. Since the inexperienced worker is instantaneously trained, the promotion decision becomes identical for the two types of workers and so does the status quo income given promotion. Thus, the only way for the firm to be indifferent between hiring an experienced vs. and inexperienced worker is for the latter to accept a wage that is reduced by the full training costs (the usual general human capital result). Thus, the inexperienced worker gains more upon promotion than the experienced worker.

The scenario outlined above assumes that the inexperienced worker receives training instantaneously upon being hired into the temporary job. If training is not instantaneous, it is possible that a productivity shock could occur before the worker is trained. Even in such a scenario, the more experienced worker is more likely to have been trained before starting a temporary job than a less experienced worker. If one can only be promoted upon receiving training and a favorable productivity draw, then the less experienced worker will have a larger expected gain in wages upon promotion than the more experienced worker. Alternatively, it may be possible for one to be promoted before being trained. Of course, for such a decision to be profitable for the firm, the productivity shock threshold needs to be higher for a currently untrained worker than a currently trained worker. In such a case, the firms which promote inexperienced workers will have had on average more favorable draws than those which promote more experienced workers, and this difference in firm selectivity will tend to raise the observed return to promotion for inexperienced relative to that for experienced workers. This is the case because promotion for the untrained worker also entails training, which raises the value of unemployment; in contrast, for a trained worker, the value of unemployment stays the same upon promotion.

The basic prediction of this model is that the more experienced workers are more likely to be promoted into permanent jobs but to receive a smaller wage gain than less experienced workers due to the selection effect. More highly-trained temporary workers require a smaller positive productivity shock to warrant promotion into a permanent job. Moreover, the Nash framework is used for analytical convenience. The key result from the framework is that workers can appropriate at least a portion of mandated firing costs, and a Nash bargaining model isn't necessary for such an outcome. For example, in a monopoly union setting, we might expect higher firing costs to make firm demand less elastic with respect to wages, thus raising union wage demands. And in a monopoly union setting, we would also expect firms with a more positive productivity draw to have a higher willingness to pay.

IV. Data and Descriptive Patterns

I use the ECHP data for 1995-2001 for the following countries to study the structure of the wage premium for permanent employment contracts: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Spain, and the United

Kingdom.[8] This is a panel data base that follows individuals over the 1994-2001 period. The questions were harmonized as much as possible in order to produce a data base that would provide comparable information across countries.[9] Beginning in 1995 for all of these countries except Finland and in 1996 for Finland, the ECHP asked each employed wage and salary worker whether his/her job was characterized by a fixed term contract. Specifically, each employed respondent is asked: "What type of employment contract do you have in your main job?" The possible responses are: 1) permanent employment; 2) fixed-term or short-term contract; 3) casual work or no contract; 4) some other working arrangement. I include only those with responses 1) or 2), that is, those that state they have a permanent or a temporary employment contract.

Table 1 shows mean values for the incidence of temporary employment among wage and salary workers by country and gender, for ages 16-65. The figures are weighted using the ECHP's sampling weights which I adjusted so that each country receives the same total weight. About 9-11% of the sample has a temporary contract. Moreover, women have a higher incidence of temporary employment in each country than men do, and temporary jobs are especially prevalent in Spain. Finland and Portugal also have a relatively high incidence of temporary jobs as well.[10]

Table 2 shows the mean of the log of hourly earnings expressed in purchasing power parity units in 2001 US dollars by country, gender and contract type.[11] In all cases except for women in the United Kingdom (where pay is the same across contract type) permanent jobs pay more than temporary jobs, usually considerably so. For example, for men, there is an average

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8 Of the fifteen countries in the ECHP, these are the only ones with data on contract type (i.e., permanent vs. temporary), with repeated observations on the same person, and complete data on the explanatory variables.

9 For further description of the methods and sample characteristics of the ECHP, see the Eurostat web site: http://circa.europa.eu/irc/dsis/echpanel/info/data/information.html .

10 Earlier work has shown that the ECHP data on the incidence of temporary employment contracts match up well with published sources such as the OECD. See Kahn (2010).

11 The ECHP provides purchasing power parity rates for each country in each year, allowing one to transform the earnings data into US purchasing power units for that year. These transformed earnings variables were then corrected for US inflation by using the Personal Consumption Expenditures deflator for the US, taken from www.bea.gov. I excluded observations with hourly earnings less than $1 or greater than $300 in 2001 purchasing power parity units. These exclusions amounted to about 0.2% of the sample. Because of the necessity of excluding those with implausible wage figures, I needed to express wages in a common currency and correct for inflation.

0.325 log point gap favoring permanent jobs, while for women, the average permanent job wage advantage gap is 0.245 log points. France, Spain and the Netherlands show especially large pay gaps favoring permanent contracts. Of course, Table 2 doesn't control for individual measured characteristics such as human capital or sector, and it also doesn't adjust for unmeasured person effects. The next section describes the regression design that attempts to estimate the effect of obtaining a permanent job at the individual level.

Tables A1-A3 provide some further descriptive detail on the incidence of temporary employment by age-gender-country group (Table A1) as well as wage levels for temporary and permanent employment by age-gender-country group (Tables A2-A3). Table A1 shows that temporary jobs are far more common among younger than older workers. While the ECHP doesn't have information on the respondents' full work history, I note that for a subsample of workers of age less than or equal to 22 years old, 27-28% of these workers were in temporary jobs. This is a considerably larger incidence than those for workers under 35 years old as shown in Table A1. Moreover, in France and Spain, 51-78% of workers 22 years of age and younger were in temporary jobs. These findings suggest that many workers in Europe do start their careers in temporary jobs.

Tables A2 and A3 show that the permanent job wage advantage tends to be larger among older workers, seemingly in contradiction to the theoretical model presented earlier. However, the Tables do not of course control for other factors affecting wages, both measured characteristics as well individual fixed effects. The following section presents the methodology through which these factors can be controlled.

V. Empirical Procedures and Basic Regression Results

The basic empirical setup for testing the predictions about the wage impact of permanent employment is to estimate the following individual fixed effects model of the determinants of the log of hourly earnings:

(10) Ln Wage=f(Age, Agesq, Edlow, Edmid, Tenure, Tenuresq, Temp, Temp*Age, Temp*Agesq, Temp*Immig, Temp*Edlow, Temp*Edmid, Temp*Tenure, Temp*Tenuresq, Industry dummies, Regional Unemployment Rate, Occupation dummies, Year dummies, u),

where for each employed wage and salary worker Ln Wage is the log of hourly earnings (defined above), Age is age in years, Agesq is Age squared, Edlow and Edmid are, respectively, dummy variables for low (ISCED levels 0-2) and middle levels (ISCED level 3) of schooling with high levels of schooling the omitted category (ISCED levels 5-7), Temp is a dummy variable for a temporary employment contract, Immig is a dummy variable equaling one if the respondent was either born in a foreign country or is not a citizen, Tenure is number of months of tenure with one's current employer, Tenuresq is Tenure squared, and u is a disturbance term.[12] Equation (10) was estimated separately for men and women pooling the data across countries with the standard errors clustered by country. I use the ECHP sampling weights adjusted so that each country receives the same weight. Since I use individual fixed effects (formally, the within estimator, in which variables are defined as deviations from the person-specific mean in an unbalanced panel), time-invariant variables such as country dummies or immigrant status are not included. I use the full estimation sample including those who don't change contract type (i.e. permanent or temporary job status) in order to obtain more efficient estimates of the time-varying parameters such as Age or Tenure. Since schooling can change this is a time-varying factor. Moreover, Tenure doesn't increase one year for each year in the panel, since people change jobs. Thus, Tenure effects are separately identified from the impact of Age. In addition, in some supplementary analyses, I estimate equation (10) separately by country and gender with standard errors clustered at the individual level.

[12] As discussed below, combining non-citizens and those born in another country into one category is necessary in order to use the sample of 13 countries shown in Tables 1 and 2. Also as discussed below, I additionally in some analyses restrict the models to countries with information on the respondent's birthplace.

The key explanatory variables in (10) are those relating to temporary employment, specifically, its main effect and its interactions with Age, Agesq, Tenure, Tenuresq, and immigration status. I also include interactions with the education variables, although the theory outlined above doesn't distinguish across levels of formal schooling. The theoretical analysis predicts that temporary employment will have positive interaction effects with Age and Tenure, and a negative interaction effect with the immigrant dummy variable on the idea that the young, those recently hired, and immigrants have less experience in the domestic labor market or the current firm than older workers, those with longer Tenure, and natives. Since women have less labor market experience than men, the same reasoning predicts that the women will face a larger discount in temporary vs. permanent jobs than men do. However, to the extent that there is occupational or industrial segregation by gender, even permanent jobs for women may not be as well protected or as relatively high paying as those for men. Prior research has emphasized the relative concentration of immigrants, women and the young in temporary jobs when employment protection is more strict since these groups have on average less labor market experience in the host labor market (Kahn 2007).

The controls in equation (10) include basic human capital variables, Tenure, Industry and Occupation (see the Appendix for the actual Industry and Occupation categories), as well as the Regional Unemployment Rate. Regional unemployment rate information was collected from the European Labour Force Survey and matched to the regional indicators in the ECHP data.[13] Because employment in a temporary job can affect one's future industry, occupation and job tenure, I also estimate the basic model excluding these variables. Such estimates can be viewed as reduced forms of the impact of temporary employment relative to permanent employment.

As mentioned, immigrants are defined as those who were either born in a foreign country or are not citizens. The ECHP has enough information to define this variable for each of the 13 countries in Tables 1 and 2. However, to focus further on the immigrant labor market, I analyze

[13] I am grateful to Alison Davies and Rhys Powell for their help in acquiring the European Labour Force Survey regional unemployment rate data. Since the ECHP did not collect regional information for Denmark or the Netherlands, I used the national unemployment rate for those countries.

a subset of countries for which the ECHP has data on the respondent's birthplace and time since migration to the current country. These include Belgium, Denmark, France, Ireland, Portugal, and Spain.[14] On this subsample, I am able to test whether the impact of being in temporary job differs depending on whether one was born in a foreign country in the European Union or outside the European Union. Moreover, I also test whether time in the current country affects the returns to a permanent vs. a temporary job. The logic of the model described earlier implies that the wage effect of a permanent job should be greater for those born outside of the EU and for more recent migrants. Unlike much research on immigrant assimilation profiles, the ECHP allows one to follow the same immigrant over time. However, there may still be a bias in estimating such profiles to the extent that there is selective outmigration (Lubotsky 2007).

The validity of the fixed effects approach depends on the assumption that changes in individuals' unmeasured productivity are not correlated with movements from temporary to permanent or permanent to temporary jobs, controlling for the other variables in the model. However, as in the research studying the union wage impact that estimates wage changes among workers who change their union status (e.g. Card 1996), we cannot rule out unmeasured productivity changes for individuals as a cause of movements to or from permanent employment. The model suggested that such moves would be caused by firm-related developments pertaining to productivity shocks and that workers without prior training would require a larger productivity shock in order to be promoted into a permanent job. Below, I present some direct information on training and wages received by workers on temporary jobs, suggesting the relevance of the theoretical model, even if one cannot rule out time varying individual effects as an alternative explanation. In addition, I include only those with jobs in a given year, which contributes to the unbalanced nature of the panel. It is of course possible that those without wage offers or those

[14] I exclude the United Kingdom from this analysis because the number of respondents for whom the foreign birthplace is identified is too small. Specifically, there were only 90 such individuals in the United Kingdom sample, compared to 318-1913 in each of the six countries analyzed.

with offers below their reservation wages dropped out of the labor market and that we therefore only observe successful transitions to employment from employment.[15]

Table A4 shows mean values for the regression samples. Men in the ECHP outearn women by 0.14 log points (recall that wages are measured in real purchasing-power corrected units); women are slightly more educated than men, with women having a higher incidence of high education (omitted) group and a lower incidence in the two lower groups; women are slightly more likely to be in temporary jobs and have shorter current Tenure.

Tables 3 and 4 contain basic individual fixed effects regression results for the determinants of the log of hourly earnings for the pooled sample of 13 countries, separately for men and women. Column (1) of Tables 3 and 4 shows the effects of employment in a temporary job without controlling for Industry, Occupation, Tenure or for interactions between temporary employment and other variables. The effect is -0.0322 for men and -0.0149 for women, with both effects significant. The corresponding Ordinary Least Squares (OLS) estimates, with country dummies and an immigrant dummy additionally included, are -0.1435 (standard error 0.0315) for men and -0.1146 (standard error 0.0263) for women.[16] Recall from Table 1 that the mean differential between wages in permanent and temporary jobs was 0.325 log points for men and 0.245 log points for women, which are substantially larger than the OLS regression coefficients in magnitude. Thus, most of the difference in the mean wages between permanent and temporary jobs is due to individual characteristics, with both measured and unmeasured characteristics having an important effect. The gross wage differential between permanent and temporary jobs is interesting in a descriptive sense in the same way that the gross union-nonunion wage differential describes one dimension of wage inequality. However, from an

[15] Of those who had temporary jobs in the previous year, 82-85% were employed the next year. The corresponding transition to employment of those with permanent jobs was 95-96%.

[16] As noted, the fixed effects estimates are identified from those who changed permanent job status. Most of such cases involved moves from temporary to permanent jobs: using sampling weights, I compute that 63-65% of those who switched status changed from temporary to permanent jobs. These workers experienced 8-9% real wage gains, in contrast to real wage gains of 4-7% for those who switched from permanent to temporary jobs. Somewhat surprisingly, stayers' wages rose by 4-6%. Of course, these means don't control for other factors affecting wages, but they do suggest that most of the explanatory power comes from the large wage gains of workers who move from temporary to permanent jobs, which is the setting for the model presented above.

individual worker or firm's point of view, the differential purged of the effects of measured and unmeasured characteristics is more relevant. And the fixed effects results in Tables 3 and 4 do indicate at least a modest return to a permanent job for both men and women, with a slightly larger effect for men.

Columns (2) and (3) of Tables 3 and 4 show that the estimated return to a permanent job largely holds up when I control for current Tenure and also Industry and Occupation. Specifically, the temporary employment coefficient with controls for Tenure, Industry and Occupation included is 82-89% as large in absolute value as it is without these additional controls, although the effect for women now just misses being statistically significant. The male result remains significant, however. The column (3) results suggest that there is a modest average return to permanent employment controlling for Tenure, Industry and Occupation, as well as unmeasured person effects.

Columns (4)-(6) show the results of the key interactions between temporary employment and Age, immigrant status, and current job Tenure. The results are qualitatively similar for men and women. Specifically, temporary employment has less negative effects on the wages of older workers, non-immigrants and those with longer current Tenure. Each of these results was predicted by reasoning outlined earlier about training, temporary employment and permanent employment. For men, the temporary employment interactions with the Age variables and immigrant status are significant, while for women the interactions with Age and Tenure are significant.[17] To illustrate the magnitude of these interaction effects, consider the fully specified model in column (6) of Tables 3 and 4. For men, the interactions between temporary employment and Age imply that the impact of permanent employment is about 0.10 log points more negative at Age 16 than at the mean Age of 39 years, while for women temporary employment has an effect on wages that is about 0.06 log points more negative at Age 16 than at the mean Age of 37. Column (6) of Tables 3 and 4 implies that interaction effects with Tenure

[17] For women, the interaction effects with immigration status are 1.44-1.77 times their standard errors in absolute value.

are smaller in magnitude than those with Age: for men, the wage effect of temporary employment is 0.0044 log points more negative at zero tenure than at two years' Tenure, a long duration for a temporary job (roughly 75% of the temporary workers have current Tenure no longer than two years); for women, the effects of a temporary job 0.0123 log points more negative at zero Tenure than at two years' Tenure. Moreover, as predicted, temporary employment has more negative effects for immigrants with effects of -0.059 for men and -0.0365 for women.

All told, the impact of getting a permanent job is moderately higher for those with less experience in the domestic labor market or firm. Combining the interaction effects for Age, immigrant status and Tenure, we can conclude that young immigrant men (Age 16) just starting their jobs have a 0.16 log points higher return to getting a permanent job than natives of average Age with two years of current Tenure; for women, the corresponding exercise yields an effect of 0.12 log points.[18]

In addition to these interaction effects, Tables 3 and 4 also contain information about the impact of temporary employment by education level and gender. For men, the less educated and those with a middle level of schooling have a modestly smaller return to getting a permanent job (.01-.03 log points) than those with high levels of schooling (i.e. the interactions Temp*Edlow and Temp*Edmid have positive coefficients, with the interactions with low schooling levels being significant). For women, however, these interactions are very small in magnitude and insignificant.

As noted, the return to getting a permanent job is on average slightly larger for men than women, even though the basic model implied that those with less experience should receive a higher return to permanent employment. Since women on average have less experience than

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[18] The interaction models of columns 5 and 6 of Tables 3 and 4 imply much less negative temporary job wage effects for older, native, less educated workers with longer Tenure. For example, the models imply that for 50 year old, less educated natives with 2 years of Tenure, temporary jobs pay 0.004 log points more than permanent jobs for women and 0.01 less for men. Both of these estimates are statistically insignificant. However, they do suggest for workers with experience in the domestic labor market and their own firms, being promoted into or moving into a permanent job has very small wage effects. Perhaps the additional job security afforded by a permanent job is valuable to these workers, who evidently have gained skills even on their temporary jobs, an issue I explore below.

men, one might have expected a higher return to permanent employment for women than for men. However, to the extent that there are glass ceilings in employment for women, wages in permanent jobs may constrained for them (Arulampalam, Booth, and Bryan 2007). I note that the slightly larger male permanent job wage effect persists after controlling for largely one digit Industry and Occupation (column 3 of Tables 3 and 4); thus, for the glass ceiling phenomenon to explain this difference, there must be gender segregation within these categories, a plausible result. For example, Anker (1998, p. 102) found that for a sample of OECD countries with data from roughly 1990 the gender occupational segregation index was 0.38 when occupations were defined at the one digit level but fully 0.63 when they were defined at the three digit level.[19] Thus, it is possible that women are on a different track from men. However, among women, the least experienced still obtain the highest return to promotion to a permanent job.[20]

Up to now, I have included all countries with data on the key variables and defined immigrant as either being foreign born or not being a citizen. For a subsample of the ECHP countries, one can discern the respondent's actual birthplace, allowing for a more detailed look at the role of immigration. These countries include Belgium, Denmark, France, Ireland, Portugal and Spain. Tables 5 and 6 show fixed effects log hourly earnings results for these countries where I take a more detailed look at immigration than is possible in the full sample of 13 countries. Specifically in Tables 5 and 6 I study the role of being foreign born inside vs. outside the European Union (EU) as well as the role of time in the current country. I expect that immigrants born in the EU and those who have been in the current country for a longer time to have better knowledge about employment practices than immigrants born outside the EU or who

[19] The gender occupational segregation index (or index of dissimilarity) is the fraction of women or men who would have to change jobs in order to achieve perfect gender integration.

[20] In an earlier version of this paper (Kahn 2013), I report estimates of the basic wage models separately by country-gender group. The results, while less statistically significant than the pooled results in Tables 3-4 were largely consistent with them. Specifically, temporary employment usually had positive interactions with Age, negative interactions for immigrants, and, for women, positive Tenure interactions. For men, the Tenure insteractions had no consistent pattern, as perhaps suggested by their insignificant effects in Table 3. Since the separate country-by-country regressions control for Year effects and Age, they in effect adjust for conditions at labor market entry.

have recently arrived. Using the same logic as discussed earlier, I expect a smaller return to permanent employment for immigrants born in the EU vs. outside the EU as well as for immigrants with more years since migration (Ysm).[21] The findings for men shown in Table 5 largely support these hypotheses; however, for women, the results shown in Table 6 do not support these hypotheses about immigrants.

First, looking at the male results in Table 5, column 2 shows that the effect of a temporary job for those born outside the EU is 0.17 log points more negative than for the native born, a statistically significant result that is large in magnitude. Moreover, the effect of being foreign born for those born within the EU is 0.15 log points more positive than this result, a difference that is also statistically significant. In fact, relative to the native born, being foreign born within the EU leads to a temporary job wage effect that is only 0.022 log points more negative (i.e. the sum of the Temp*Foreignborn and Temp*Foreignborn in the EU coefficients), a difference that is not statistically significant. Thus, regarding the gains to permanent employment, foreign born males who originated from within the EU resemble the native born much more closely than they do foreign born males originating from outside the EU. Moreover, the wage disadvantage of a temporary job is less the longer a foreign born male has been in the current country, although the Temp*Ysm and Temp*Ysm squared coefficients are not significant individually or as a pair. As was the case for the full sample of 13 countries, the effect of a temporary job continues to be significantly less negative for older men in the subsample of six countries shown in Table 5. However, the Tenure interactions are small, insignificant and opposite in sign to what I found for the larger sample. Finally, adding further interactions between Ysm and Ysm squared and being born in a foreign EU country led to statistically insignificant results.

For women, Table 6 shows very small and statistically insignificant interaction effects between temporary employment and being foreign born as well as temporary employment and

[21] Since I estimate individual fixed effects models using the longitudinal feature of the ECHP data, the analysis of Ysm doesn't suffer from "single cross-section" problem identified by Borjas (1985); however, the results for Ysm, as noted earlier, may be affected by selective outmigration.

being foreign born from the EU. Moreover, temporary employment actually has a negative interaction effect with Ysm. While the Temp*Ysm interaction is marginally significant, the two interactions together are not significant in either column (4) or column (5) of Table 6. In addition, as was the case for men, adding further interactions between Ysm and Ysm squared and being born in a foreign EU country led to statistically insignificant results. Thus, for women, our hypotheses about immigration are not borne out. It is possible that labor force selection issues are more salient for immigrant women than immigrant men, and these could have masked an effect of getting a permanent job for immigrants. But even for this subsample of six countries, the Age and Tenure interactions with temporary employment for women are similar to what they were in the full sample of 13 countries.

VI. Temporary Employment, Training and Wages

As discussed above, it is possible that my findings are consistent with differential selection into permanent jobs caused by differential changes in unmeasured productivity. For example, it is possible that temporary workers with long tenure or experience in the labor market have declining productivity relative to those who are not in temporary jobs. Even if such workers get promoted into or obtain permanent jobs, their future wages may reflect this hypothetical downward relative productivity path, resulting in a smaller permanent job wage premium even in a fixed effects model. To shed light on this possibility, I examined wage levels and the incidence of company-sponsored formal training for those on temporary jobs. The results are shown in Appendix Tables A5 and A6. First, Table A5 shows that for both men and women, workers who are older, more highly educated, or who have longer tenure on a temporary job earn significantly higher wages than otherwise. These findings suggest that workers are learning useful skills on temporary jobs. The effects for immigrants, however, are statistically insignificant.

Table A6 shows the results of analyses for workers on temporary jobs of the determinants of whether one has received during the term of the ECHP panel any company-sponsored or subsidized education or training, a variable which is available for 11 of the 13 countries analyzed above (see the note to Table A6).[22] Standard errors are clustered at the individual level to reflect the fact that the dependent variable is the result of training observations across waves up to the given year. I define the dependent variable this way so that one tell if a temporary worker has ever received training (at least where we can observe it). The results are very similar to the wage results in Table A5. Specifically, temporary workers who are older, more highly educated, or have longer current Tenure are more likely than otherwise to have received some form of company-sponsored tenure.

The model on which Tables 3 and 4 are based posits that less experienced workers need a larger firm-specific productivity shock to justify promotion into a permanent job, and the wage and training results in Tables A5 and A6 are consistent with this model. It is possible that the less experienced workers on temporary jobs who located permanent jobs also experienced person-specific increases in productivity that were larger than those received by more experienced temporary workers who got promoted. In some sense the model predicts this as well since less experienced workers have less training. While one cannot rule out the impact of time-varying unobserved variables in addition to the firm productivity shocks hypothesized above, the analysis of wages and training among temporary workers provides some evidence in favor of the model discussed above, at least with respect to Age and Tenure. For immigrants, however, there is little evidence of a differential relative to natives in wages or training on temporary jobs even though I have found some evidence of a higher premium for immigrants for obtaining a permanent job.

VII. Implications for Wage Inequality: The Case of Spain

[22] When I restricted the basic fixed effects wage regression model to these 11 countries, the results were very similar to those in Tables 3 and 4.

The analysis so far has suggested that there is a modest overall premium associated with obtaining a permanent job and that this premium varies considerably across skill groups. The model suggested that firing costs associated with permanent jobs were responsible for this pay premium, which, like union wage effects, can influence overall wage inequality. Such an impact is more likely the more equal the sizes of the temporary and permanent job sectors, and Table 1 shows that temporary jobs are much more prevalent in Spain than elsewhere. Specifically, the incidence of temporary jobs in Spain is 30-33% compared to an average of only 6-9% in the other countries. In a standard decomposition of the variance of log wages into within group and between group components, where in this case the groups are temporary and permanent workers, only for Spain does the permanent-temporary wage differential have a noticeable effect on the overall wage variance. Specifically, consider the following decomposition of the variance of log wages for men or women:

$$(11)\quad Var(y)=aVar(y_p)+(1-a)Var(y_t)+a(ybar_p-ybar)^2+ (1-a)(ybar_t-ybar)^2,$$

where y is log wages for all (male or female) workers, a is the fraction of workers in permanent jobs, y_p is log wages for permanent jobs, y_t is log wages for temporary jobs, and ybar, $ybar_p$ and $ybar_t$ are sample means for y, y_p and y_t respectively (again for the male or female sample). In equation (11), the first two terms comprise the within sector variance component, while the second two add up to the between sector component. The latter is larger the closer the fraction of permanent jobs is to .5 and the larger permanent-temporary average wage differential.

Performing the decomposition in equation (11) for, for Spain, one finds that 14% of the female and 18% of the male log wage variance is accounted for by the average wage differential between permanent and temporary jobs (the between effect), while for the other countries, the between sector effect ranges from 0 to 6 percent, with an unweighted average of only 3%. The smaller effects in the other countries are due to the much smaller size of the temporary jobs sector. Based on these results, I pursued the role of the returns to a permanent job in affecting overall wage inequality in Spain. Because of both measured and unmeasured worker characteristics, the overall permanent-temporary job wage differential is an overestimate of the causal effect of employment in permanent jobs. To focus on these causal effects, I estimated the basic fixed effects wage model separately for Spain, and the results are shown in Table A7. The findings are qualitatively similar to those in Tables 3 and 4. To assess the role of the returns to permanent employment, I computed for each worker in Spain, the following:

$$\text{(12)} \qquad \text{Temporary job wage effect} = b_{temp}*Temp + b_{tempage}*Temp*Age + b_{tempagesq}*Temp*Agesq + b_{tempimm}*Temp*Immig + b_{temptenure}*Temp*Tenure + b_{temptensq}*Temp*Tenuresq$$

where the terms with b and subscripts refer to fixed effects regression coefficients, and Agesq and Tenuresq are age and tenure squared respectively.

Equation (12) shows the contribution of the temporary job wage effect to an individual's predicted wage. It is zero for anyone with a permanent job but of course can be nonzero for those with a temporary job. I then compute the standard deviation across the sample in the temporary job wage effect to obtain the contribution of the returns to a temporary job to wage inequality. For men, the estimates in Table A7 imply that this standard deviation is 0.035-0.036 log points. This is about 7% of the overall standard deviation of male wages in Spain of 0.493. For women, the estimates are 0.022-0.023, or about 4.5-4.6% of the standard deviation of female wages of 0.494. These effects are modest but still positive and suggest that the segmentation produced by a dual labor market of permanent and temporary jobs can influence wage inequality.[23] To study the specific role of heterogeneity in the permanent job wage effect, I re-

[23] I do not take account of the estimation error in the coefficients in forming the standard deviation of the predicted contribution to wage inequality. Consideration of this additional source of variance implies that these modest contributions are likely to be overestimates.

estimated the models in Table A7 excluding interactions with the temporary employment indicator. I then computed:

d_{temp}*Temp,

where d refers to the coefficient on temporary employment, for each individual and computed standard deviation of (d_{temp}*Temp) across the entire sample. In the noninteraction model, the effect of temporary jobs contributes 0.020 log points to the male standard deviation and 0.004 to 0.005 to the female standard deviation of log wages.[24] Recall that in the interaction models, the contribution ranged from 0.035-0.036 for men and 0.022-0.023 for women. Thus heterogeneity in the temporary job wage effect accounts for an increase of 0.015-0.016 log points for men and 0.018 log points for women. Again, this is a modest effect.

VIII. Robustness Checks

I attempted several alternative specifications, all of which led to similar results to those in Tables 3 and 4 (results available upon request). First, I restricted the sample to those age 22-60 to abstract from schooling and retirement decisions. Second, I estimated the models with no weighting and also with no individual sampling weights but constraining each country to have the same total weight. Third, I estimated the models excluding either the education or the age variables (recall that Tables 3 and 4 show that the findings are not sensitive to exclusion of tenure).

IX. Conclusions

[24] The Temp coefficients from the models without interactions were -0.0431 (model including Tenure and Tenure squared main effects but not Industry and Occupation) and -0.0434 (model additionally including Industry and Occupation) for men, and these effects were both significant at the 1% level; for women, the corresponding estimates were -0.0117 and -0.0155, with the latter effect significant at 10%.

In this paper, I have used ECHP data to investigate the premium workers command in permanent jobs relative to temporary jobs across thirteen European countries. A basic framework to understand this issue comes from Blanchard and Landier's (2002) research in which it is assumed that firms start workers in temporary jobs. In the presence of higher firing costs for permanent jobs relative to temporary jobs, firms will be reluctant to promote workers unless the economic circumstances of the firm warrant it. Once promoted, workers are able to appropriate some of the firing costs, since these raise the value of continuing the employment match once the worker is promoted. Thus, an important determinant of the wage premium in a permanent job is the value of the match relative to breaking it up and starting over with a temporary employee.

I hypothesized that before being promoted into a permanent job, inexperienced workers must receive training in the temporary job to which they have been hired. In equilibrium, their starting wages in the temporary job will be below the level of starting wages for experienced, trained workers starting a temporary job. Because of the wage discount at the beginning of the temporary job for less experienced workers, the wage gain conditional on promotion to a permanent job will be greater for them. In this framework, we expect to observe a higher incidence of permanent employment among more experienced workers because they have had more opportunities to be in firms that receive a favorable productivity shock, and the exit probability from permanent jobs is relatively low.

I tested the prediction that the permanent job wage premium falls as labor market experience rises using longitudinal data from the ECHP. Taking into account individual fixed effects, I found that among men the wage premium for a permanent vs. temporary job is indeed significantly lower for older workers and native born workers; for women, the wage premium for a permanent job was found to be lower for older workers and those with longer current job tenure. Moreover, there is some evidence that the wage return to a permanent job was especially high for immigrant men born outside of the European Union; in contrast, the premium was much smaller for those born within the EU. These findings all suggest that the gain to promotion into permanent jobs is indeed higher for those with less experience in the domestic labor market; moreover, previous findings that immigrants and the young are more likely to be in temporary jobs than the native born and older workers are consistent with the view outlined above as well (OECD 2002; Kahn 2007). I also found that among those with temporary jobs, those with more experience (proxied by age and tenure) earn higher wages and are more likely to have received employer sponsored or subsidized schooling or training. These findings are consistent with the model which emphasized training as a key factor explaining heterogeneity in the permanent wage premium.

An implication of these results is that policies that contribute to dual labor market structures consisting of a protected, permanent sector and an unprotected, temporary sector also contribute to wage inequality. I documented a modest contribution of temporary employment to wage inequality in Spain, the country with the highest incidence of temporary jobs in the sample. However, lifetime wage inequality may be less affected to the extent that training on temporary jobs leads to permanent employment. Nonetheless, given the current stagnation in European labor markets and the extremely high rates of unemployment among youth in several countries,[25] it may take a long time before new entrants can realize the gains to being promoted into protected jobs.

[25] For example, among males age 15-24, unemployment rates in 2012 were 48.4% in Greece, 33.7% in Italy, 36.4% in Portugal, and 54.4% in Spain; for females in this age group, the corresponding figures were 63.2% in Greece, 37.5% in Italy, 39.2% in Portugal, and 51.8% in Spain. See OECD (2013), pp. 246-7.

Appendix: Occupation and Industry Dummy Variables Categories

Occupations

Legislators, senior officials and corporate managers
Managers of small enterprises and other managers
Physical, mathematical, life science, health and engineering science professionals
Teaching professionals
Other professionals
Associate professionals (technicians in physical, engineering, life and health sciences)
Teaching associate professionals and other associate professionals
Clerical workers
Personal service workers
Sales workers
Skilled agricultural and fishery workers
Craft workers
Operatives
Elementary occupations (the omitted category from the regressions)

Industries

Agriculture, forestry and fishing
Mining, electricity, gas and water supply
Nondurable manufacturing
Durable and other manufacturing
Construction
Wholesale and retail trade
Hotels and restaurants
Transport, storage and communications
Finance
Real estate
Public administration
Education
Health and social work
Other community, social and personal services (the omitted category from the regressions)

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Table 1: Temporary Employment as a Fraction of Total Employment, Wage and Salary Workers

	Men		Women	
	Incidence	Sample Size	Incidence	Sample Size
Austria	0.045	10251	0.060	7212
Belgium	0.075	5120	0.132	4267
Denmark	0.041	6032	0.065	5830
Finland	0.120	6502	0.175	6641
France	0.080	14981	0.093	12548
Germany	0.063	18207	0.089	13266
Greece	0.078	8589	0.105	5325
Ireland	0.040	7818	0.074	5588
Italy	0.066	17201	0.081	11095
Netherlands	0.027	14053	0.050	9518
Portugal	0.101	14191	0.145	10771
Spain	0.304	17882	0.326	9615
UK	0.034	12759	0.037	12297
Total	0.085	153586	0.108	113973

Source: ECHP, 1995-2001. Sample weights are adjusted so that each country receives the same total weight.

Table 2: Mean Log Hourly Earnings, Permanent and Temporary Jobs, Wage and Salary Workers

	Men					Women				
	Permanent Job		Temporary Job		Difference	Permanent Job		Temporary Job		Difference
	Mean Log Hourly Earnings	Sample Size	Mean Log Hourly Earnings	Sample Size	(Perm-Temp)	Mean Log Hourly Earnings	Sample Size	Mean Log Hourly Earnings	Sample Size	(Perm-Temp)
Austria	2.031	9849	1.837	402	0.194	1.842	6769	1.740	443	0.103
Belgium	2.096	4769	1.882	351	0.214	2.024	3707	1.913	560	0.111
Denmark	2.152	5769	1.984	263	0.167	2.069	5448	1.943	382	0.127
Finland	1.888	5743	1.714	759	0.174	1.759	5460	1.630	1181	0.129
France	2.167	13778	1.732	1203	0.435	2.042	11293	1.626	1255	0.416
Germany	2.138	17061	1.846	1146	0.293	1.870	12054	1.722	1212	0.148
Greece	1.801	7867	1.597	722	0.204	1.645	4695	1.487	630	0.158
Ireland	2.178	7474	2.020	344	0.158	1.998	5135	1.948	453	0.050
Italy	1.984	15950	1.753	1251	0.230	1.909	10103	1.701	992	0.207
Netherlands	2.194	13718	1.768	335	0.427	2.069	9070	1.739	448	0.330
Portugal	1.433	12743	1.302	1448	0.130	1.353	9188	1.147	1583	0.206
Spain	2.091	12378	1.636	5504	0.455	1.976	6437	1.577	3178	0.399
UK	2.117	12324	1.980	435	0.138	1.940	11834	1.940	463	-0.001
Total	2.024	139423	1.699	14163	0.325	1.883	101193	1.638	12780	0.245

Source: ECHP, 1995-2001. Sample weights are adjusted to that each country receives the same total weight.
Hourly earnings are in 2001 US purchasing power corrected dollars.

Table 3: Selected Log Hourly Earnings Individual Fixed Effects Regression Results, Men

	(1)	(2)	(3)	(4)	(5)	(6)
Age	0.0803**	0.0777**	0.0765**	0.0786**	0.0760**	0.0749**
	(0.0078)	(0.0079)	(0.0075)	(0.0079)	(0.0081)	(0.0077)
Agesq (/100)	-0.0626**	-0.0575**	-0.0564**	-0.0607**	-0.0555**	-0.0544**
	(0.0069)	(0.0067)	(0.0063)	(0.0069)	(0.0068)	(0.0064)
Edlow	-0.0377**	-0.0356**	-0.0346**	-0.0393**	-0.0372**	-0.0362**
	(0.0086)	(0.0080)	(0.0081)	(0.0077)	(0.0071)	(0.0073)
Edmid	-0.0330**	-0.0315**	-0.0299**	-0.0336**	-0.0320**	-0.0306**
	(0.0084)	(0.0081)	(0.0078)	(0.0077)	(0.0073)	(0.0071)
Temp	-0.0322**	-0.0271**	-0.0286**	-0.3173**	-0.3174**	-0.3163**
	(0.0082)	(0.0075)	(0.0072)	(0.0968)	(0.0976)	(0.0931)
Unemrate	-0.0017	-0.0017	-0.0016	-0.0017	-0.0017	-0.0016
	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)
Tenure (years)		0.0058**	0.0059**		0.0057**	0.0058**
		(0.0011)	(0.0010)		(0.0011)	(0.0011)
Tenuresq		-0.0004**	-0.0004**		-0.0004**	-0.0004**
		(0.0001)	(0.0001)		(0.0001)	(0.0001)
Temp*Age				0.0155**	0.0156**	0.0154**
				(0.0048)	(0.0049)	(0.0047)
Temp*Agesq (/100)				-0.0200**	-0.0204**	-0.0200**
				(0.0062)	(0.0063)	(0.0060)
Temp*Immig				-0.0571+	-0.0569+	-0.0590+
				(0.0267)	(0.0266)	(0.0276)
Temp*Edlow				0.0279+	0.0292+	0.0289+
				(0.0149)	(0.0151)	(0.0153)
Temp*Edmid				0.0149	0.0156	0.0167
				(0.0215)	(0.0217)	(0.0214)
Temp*Tenure					0.0023	0.0024
					(0.0043)	(0.0042)
Temp*Tenuresq					-0.0001	-0.0001
					(0.0003)	(0.0003)
Year dummies?	yes	yes	yes	yes	yes	yes
Occupation dummies?	no	no	yes	no	no	yes
Industry dummies?	no	no	yes	no	no	yes
Sample size	153586	153586	153586	153586	153586	153586
r squared	0.131	0.132	0.136	0.132	0.133	0.136

+ p<0.10, * p<0.05, ** p<0.01

Standard errors clustered at the country level. Sample includes Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, Spain, and the UK. Immig is defined as being either foreign born or a noncitizen.

Table 4: Selected Log Hourly Earnings Individual Fixed Effects Regression Results, Women

	(1)	(2)	(3)	(4)	(5)	(6)
Age	0.0738**	0.0714**	0.0699**	0.0726**	0.0703**	0.0688**
	(0.0063)	(0.0065)	(0.0062)	(0.0063)	(0.0065)	(0.0062)
Agesq (/100)	-0.0530**	-0.0487**	-0.0474**	-0.0515**	-0.0476**	-0.0462**
	(0.0057)	(0.0061)	(0.0058)	(0.0056)	(0.0060)	(0.0057)
Edlow	-0.0380**	-0.0356**	-0.0326**	-0.0373**	-0.0347**	-0.0318**
	(0.0090)	(0.0089)	(0.0087)	(0.0089)	(0.0088)	(0.0087)
Edmid	-0.0327**	-0.0315**	-0.0284*	-0.0316**	-0.0303**	-0.0272*
	(0.0091)	(0.0093)	(0.0095)	(0.0092)	(0.0093)	(0.0096)
Temp	-0.0149+	-0.0108	-0.0122	-0.1805*	-0.1707*	-0.1757*
	(0.0074)	(0.0073)	(0.0072)	(0.0646)	(0.0624)	(0.0612)
Unemrate	-0.0021	-0.0022	-0.0021	-0.0021	-0.0022	-0.0021
	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)
Tenure (years)		0.0047*	0.0049*		0.0038+	0.0040+
		(0.0021)	(0.0021)		(0.0020)	(0.0020)
Tenuresq		-0.0003+	-0.0003+		-0.0002	-0.0002+
		(0.0001)	(0.0001)		(0.0001)	(0.0001)
Temp*Age				0.0092*	0.0081*	0.0082*
				(0.0036)	(0.0035)	(0.0035)
Temp*Agesq (/100)				-0.0116*	-0.0103+	-0.0103+
				(0.0050)	(0.0049)	(0.0049)
Temp*Immig				-0.0310	-0.0346	-0.0365
				(0.0216)	(0.0214)	(0.0206)
Temp*Edlow				0.0029	0.0048	0.0059
				(0.0152)	(0.0154)	(0.0149)
Temp*Edmid				-0.0037	-0.0032	-0.0029
				(0.0144)	(0.0141)	(0.0137)
Temp*Tenure					0.0135**	0.0133**
					(0.0040)	(0.0040)
Temp*Tenuresq					-0.0009**	-0.0009**
					(0.0003)	(0.0003)
Year dummies?	yes	yes	yes	yes	yes	yes
Occupation dummies?	no	no	yes	no	no	yes
Industry dummies?	no	no	yes	no	no	yes
Sample size	113973	113973	113973	113973	113973	113973
r squared	0.143	0.144	0.148	0.143	0.144	0.149

+ p<0.10, * p<0.05, ** p<0.01

Standard errors clustered at the country level. Sample includes Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, Spain, and the UK. Immig is defined as being either foreign born or a noncitizen.

Table 5: Selected Log Hourly Earnings Individual Fixed Effects Regression Results, Foreign Born vs. Natives, Men

	(1)	(2)	(3)	(4)	(5)
Temp	-0.3882*	-0.3937*	-0.3992*	-0.3967*	-0.3946*
	(0.1252)	(0.1231)	(0.1263)	(0.1245)	(0.1155)
Temp*Age	0.0194*	0.0198*	0.0206*	0.0205*	0.0202**
	(0.0052)	(0.0051)	(0.0054)	(0.0053)	(0.0049)
Temp*Agesq (/100)	-0.0255**	-0.0260**	-0.0271**	-0.0270**	-0.0266**
	(0.0060)	(0.0057)	(0.0059)	(0.0056)	(0.0051)
Temp*Foreignborn	-0.1089	-0.1700+	-0.1682+	-0.3343	-0.3308
	(0.0600)	(0.0812)	(0.0810)	(0.2302)	(0.2239)
Temp*Edlow	0.0388+	0.0388+	0.0407+	0.0405+	0.0405+
	(0.0170)	(0.0168)	(0.0169)	(0.0170)	(0.0167)
Temp*Edmid	0.0203	0.0191	0.0202	0.0202	0.0200
	(0.0290)	(0.0297)	(0.0304)	(0.0300)	(0.0274)
Temp*Foreignborn in the EU		0.1478+	0.1463+	0.1338+	0.1370+
		(0.0667)	(0.0686)	(0.0610)	(0.0646)
Temp*Tenure			-0.0024	-0.0025	-0.0023
			(0.0075)	(0.0075)	(0.0073)
Temp*Tenuresq			0.0001	0.0001	0.0001
			(0.0005)	(0.0005)	(0.0005)
Tenure			0.0070*	0.0071*	0.0073*
			(0.0020)	(0.0021)	(0.0018)
Tenuresq			-0.0004*	-0.0004*	-0.0004*
			(0.0001)	(0.0001)	(0.0001)
Years since migration (Ysm)				-0.0065	-0.0060
				(0.0069)	(0.0077)
Ysm squared				0.0001	0.0001
				(0.0001)	(0.0001)
Temp*Ysm				0.0122	0.0114
				(0.0132)	(0.0129)
Temp*Ysm squared				-0.0002	-0.0002
				(0.0002)	(0.0002)
Test H0: Temp*Foreignborn+Temp*Foreignborn in the EU=0		0.4137	0.3963	.4125	.4219
Year dummies?	yes	yes	yes	yes	yes
Industry and Occupation dummies?	no	no	no	no	yes
Sample size	66024	66006	66006	66006	66006
r squared	0.157	0.157	0.158	0.158	0.163

+ p<0.10, * p<0.05, ** p<0.01

Countries include Belgium, Denmark, France, Ireland, Portugal, and Spain.

Controls include Age, Age squared, Edlow, Edmid, and Unemrate. Standard errors are clustered at the country level.

Table 6: Selected Log Hourly Earnings Individual Fixed Effects Regression Results, Foreign Born vs. Natives, Women

	(1)	(2)	(3)	(4)	(5)
Temp	-0.2212**	-0.2212**	-0.2097**	-0.2077**	-0.2189**
	(0.0316)	(0.0314)	(0.0280)	(0.0280)	(0.0299)
Temp*Age	0.0113**	0.0113**	0.0099**	0.0097**	0.0102**
	(0.0021)	(0.0021)	(0.0016)	(0.0018)	(0.0020)
Temp*Agesq (/100)	-0.0141**	-0.0141**	-0.0124**	-0.0120*	-0.0124*
	(0.0035)	(0.0034)	(0.0031)	(0.0033)	(0.0035)
Temp*Foreignborn	-0.0215	-0.0234	-0.0298	0.2248	0.2145
	(0.0324)	(0.0541)	(0.0547)	(0.1281)	(0.1286)
Temp*Edlow	0.0128	0.0127	0.0162	0.0161	0.0174
	(0.0214)	(0.0215)	(0.0206)	(0.0201)	(0.0193)
Temp*Edmid	0.0024	0.0024	0.0053	0.0052	0.0049
	(0.0224)	(0.0223)	(0.0216)	(0.0214)	(0.0215)
Temp*Foreignborn in the EU		0.0045	0.0091	0.0006	0.0002
		(0.0763)	(0.0788)	(0.0744)	(0.0697)
Temp*Tenure			0.0177*	0.0178*	0.0171*
			(0.0044)	(0.0045)	(0.0048)
Temp*Tenuresq			-0.0013*	-0.0013*	-0.0013*
			(0.0004)	(0.0004)	(0.0004)
Tenure			0.0083*	0.0084*	0.0088*
			(0.0032)	(0.0032)	(0.0030)
Tenuresq			-0.0005+	-0.0005+	-0.0005*
			(0.0002)	(0.0002)	(0.0002)
Years since migration (Ysm)				0.0109+	0.0103+
				(0.0045)	(0.0045)
Ysm squared				-0.0002*	-0.0002*
				(0.0001)	(0.0001)
Temp*Ysm				-0.0186+	-0.0175+
				(0.0079)	(0.0082)
Temp*Ysm squared				0.0003	0.0003
				(0.0002)	(0.0002)
Test H0: Temp*Foreignborn+Temp*Foreignborn in the EU=0		0.6845	0.6535	0.0162	0.0292
Year dummies?	yes	yes	yes	yes	yes
Industry and Occupation dummies?	no	no	no	no	yes
Sample size	48619	48615	48615	48615	48615
r squared	0.176	0.176	0.179	0.179	0.187

+ p<0.10, * p<0.05, ** p<0.01

Countries include Belgium, Denmark, France, Ireland, Portugal, and Spain.

Controls include Age, Age squared, Edlow, Edmid, and Unemrate. Standard errors are clustered at the country level.

Table A1: Incidence of Temporary Employment by Country by Gender by Age Group

	Men				Women			
Country	Age < 35	n	Age>=35	n	Age < 35	n	Age>=35	n
Austria	0.063	4184	0.033	6067	0.083	3317	0.038	3895
Belgium	0.142	1586	0.039	3534	0.214	1788	0.061	2479
Denmark	0.072	1841	0.027	4191	0.118	1722	0.042	4108
Finland	0.206	2226	0.073	4276	0.317	1875	0.116	4766
France	0.167	5151	0.035	9830	0.169	4366	0.053	8182
Germany	0.107	6076	0.045	12131	0.133	4620	0.068	8646
Greece	0.127	2998	0.054	5591	0.154	2414	0.063	2911
Ireland	0.065	3344	0.021	4474	0.094	3040	0.046	2548
Italy	0.119	5935	0.039	11266	0.126	4470	0.051	6625
Netherlands	0.061	4004	0.009	10049	0.067	4037	0.036	5481
Portugal	0.164	6411	0.047	7780	0.234	5122	0.058	5649
Spain	0.486	7495	0.181	10387	0.473	4936	0.172	4679
UK	0.050	5335	0.023	7424	0.044	5214	0.032	7083
Total	0.148	56586	0.049	97000	0.170	46921	0.064	67052

Source: ECHP, 1995-2001. Sample weights are adjusted to that each country receives the same total weight.

Table A2: Mean Log Hourly Earnings, Permanent and Temporary Jobs, Wage and Salary Workers, Men, by Age

	Age < 35					Age >=35				
	Permanent Job		Temporary Job		Difference	Permanent Job		Temporary Job		Difference
	Mean Log Hourly Earnings	Sample Size	Mean Log Hourly Earnings	Sample Size	(Perm-Temp)	Mean Log Hourly Earnings	Sample Size	Mean Log Hourly Earnings	Sample Size	(Perm-Temp)
Austria	1.927	3937	1.764	247	0.163	2.104	5912	1.939	155	0.165
Belgium	1.946	1372	1.840	214	0.106	2.167	3397	1.961	137	0.206
Denmark	2.048	1704	1.882	137	0.166	2.197	4065	2.108	126	0.089
Finland	1.785	1758	1.655	468	0.130	1.936	3985	1.803	291	0.133
France	1.957	4297	1.657	854	0.300	2.259	9481	1.911	349	0.348
Germany	1.967	5458	1.790	618	0.177	2.206	11603	1.901	528	0.305
Greece	1.517	2608	1.477	390	0.041	1.931	5259	1.739	332	0.192
Ireland	1.972	3102	1.907	242	0.066	2.329	4372	2.292	102	0.037
Italy	1.800	5179	1.709	756	0.091	2.068	10771	1.820	495	0.248
Netherlands	2.002	3761	1.706	243	0.296	2.291	9957	1.999	92	0.292
Portugal	1.265	5341	1.281	1070	-0.016	1.560	7402	1.365	378	0.194
Spain	1.862	3873	1.599	3622	0.263	2.187	8505	1.702	1882	0.486
UK	1.966	5066	1.830	269	0.136	2.208	7258	2.176	166	0.032
Total	1.840	47456	1.634	9130	0.206	2.119	91967	1.811	5033	0.308

Source: ECHP, 1995-2001. Sample weights are adjusted to that each country receives the same total weight.
Hourly earnings are in 2001 US purchasing power corrected dollars.

Table A3: Mean Log Hourly Earnings, Permanent and Temporary Jobs, Wage and Salary Workers, Women, by Age

	Age < 35					Age >=35				
	Permanent Job		Temporary Job		Difference	Permanent Job		Temporary Job		Difference
	Mean Log Hourly Earnings	Sample Size	Mean Log Hourly Earnings	Sample Size	(Perm-Temp)	Mean Log Hourly Earnings	Sample Size	Mean Log Hourly Earnings	Sample Size	(Perm-Temp)
Austria	1.767	3027	1.733	290	0.034	1.908	3742	1.753	153	0.155
Belgium	1.928	1380	1.914	408	0.014	2.093	2327	1.909	152	0.184
Denmark	2.003	1514	1.864	208	0.139	2.096	3934	2.038	174	0.058
Finland	1.682	1270	1.614	605	0.068	1.784	4190	1.648	576	0.136
France	1.922	3568	1.599	798	0.323	2.097	7725	1.671	457	0.426
Germany	1.785	4012	1.676	608	0.109	1.908	8042	1.765	604	0.142
Greece	1.485	1993	1.474	421	0.011	1.768	2702	1.514	209	0.253
Ireland	1.905	2724	1.920	316	-0.014	2.113	2411	2.023	137	0.090
Italy	1.757	3843	1.688	627	0.070	2.003	6260	1.724	365	0.279
Netherlands	1.973	3786	1.696	251	0.276	2.149	5284	1.806	197	0.342
Portugal	1.193	3883	1.147	1239	0.046	1.481	5305	1.147	344	0.334
Spain	1.808	2577	1.550	2359	0.258	2.087	3860	1.652	819	0.435
UK	1.903	4973	1.834	241	0.069	1.962	6861	2.030	222	-0.068
Total	1.773	38550	1.601	8371	0.172	1.953	62643	1.709	4409	0.244

Source: ECHP, 1995-2001. Sample weights are adjusted to that each country receives the same total weight.
Hourly earnings are in 2001 US purchasing power corrected dollars.

Table A4: Selected Mean Values of Variables Used in Regressions

	Men	Women
Log Real Wage	1.996	1.856
Age	39.307	37.927
Edlow	0.304	0.267
Edmid	0.414	0.391
Tenure	7.566	6.881
Temp	0.085	0.108
Unemrate	9.376	9.029
Sample size	153586	113973

Table A5: OLS Log Wage Regressions, Sample Includes Only Temporary Workers

	Men		Women	
Age	0.0481**	0.0438**	0.0422**	0.0340**
	(0.0039)	(0.0041)	(0.0084)	(0.0075)
Agesq (/100)	-0.0526**	-0.0480**	-0.0508**	-0.0398**
	(0.0045)	(0.0048)	(0.0111)	(0.0099)
Edlow	-0.3070**	-0.1520**	-0.4289**	-0.1966**
	(0.0461)	(0.0256)	(0.0603)	(0.0305)
Edmid	-0.2050**	-0.0724**	-0.2824**	-0.1100**
	(0.0255)	(0.0170)	(0.0273)	(0.0210)
Immig	0.0311	0.0184	-0.0292	-0.0027
	(0.0372)	(0.0364)	(0.0210)	(0.0212)
Unemrate	-0.0075**	-0.0067**	-0.0090**	-0.0088**
	(0.0021)	(0.0021)	(0.0016)	(0.0013)
Tenure	0.0157*	0.0096+	0.0275**	0.0190**
	(0.0053)	(0.0049)	(0.0061)	(0.0049)
Tenuresq	-0.0002	0.0001	-0.0010*	-0.0006+
	(0.0004)	(0.0003)	(0.0004)	(0.0003)
Year dummies?	yes	yes	yes	yes
Country dummies	yes	yes	yes	yes
Occupation dummies?	no	yes	no	yes
Industry dummies?	no	yes	no	yes
Sample Size	14163	14163	12780	12780
R squared	0.3420	0.4064	0.4378	0.5168

+ p<0.10, * p<0.05, ** p<0.01

Standard errors clustered at the country level. Sample includes Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Netherlands, Portugal, Spain, and the UK. Immig is defined as being either foreign born or a noncitizen.

Table A6: Probit Results for the Probability of Having Received Training During the ECHP Panel Period, Sample Includes Only Temporary Workers (marginal effects)

	Men		Women	
Age	0.0217**	0.0215**	0.0169**	0.0113+
	(0.0051)	(0.0050)	(0.0064)	(0.0066)
Agesq (/100)	-0.0323**	-0.0316**	-0.0227*	-0.0153
	(0.0070)	(0.0069)	(0.0091)	(0.0093)
Edlow	-0.2728**	-0.1351**	-0.3387**	-0.1672**
	(0.0180)	(0.0232)	(0.0177)	(0.0254)
Edmid	-0.1506**	-0.0523*	-0.1626**	-0.0403+
	(0.0183)	(0.0231)	(0.0190)	(0.0233)
Immig	0.0047	0.0019	-0.0560	-0.0202
	(0.0375)	(0.0367)	(0.0365)	(0.0377)
Tenure	0.0129+	0.0095	0.0419**	0.0361**
	(0.0067)	(0.0067)	(0.0078)	(0.0076)
Tenuresq	-0.0004	-0.0002	-0.0027**	-0.0024**
	(0.0005)	(0.0005)	(0.0006)	(0.0006)
Unemrate	-0.0119**	-0.0115**	-0.0058**	-0.0063**
	(0.0016)	(0.0016)	(0.0019)	(0.0020)
Year dummies?	yes	yes	yes	yes
Ccountry dummies	yes	yes	yes	yes
Occupation dummies?	no	yes	no	yes
Industry dummies?	no	yes	no	yes
Sample Size	12245	12245	10908	10908

+ p<0.10, * p<0.05, ** p<0.01
Standard errors clustered at the individual level. Sample includes Austria, Belgium, Denmark, Finland, Germany, Greece, Ireland, Italy, Netherlands, Portugal, and Spain. Immig is defined as being either foreign born or a noncitizen.

Table A7: Selected Log Hourly Earnings Individual Fixed Effects Regression Results, Spain

	Men		Women	
age	0.0738**	0.0734**	0.0928**	0.0942**
	(0.0059)	(0.0059)	(0.0086)	(0.0085)
agesq	-0.0456**	-0.0453**	-0.0619**	-0.0649**
	(0.0056)	(0.0056)	(0.0086)	(0.0085)
edlow	-0.0260+	-0.0276+	-0.0121	-0.0191
	(0.0147)	(0.0147)	(0.0199)	(0.0197)
edmid	0.0020	0.0020	0.0035	0.0026
	(0.0124)	(0.0124)	(0.0166)	(0.0164)
temp	-0.5789**	-0.5636**	-0.1764	-0.1542
	(0.0716)	(0.0716)	(0.1079)	(0.1066)
tempage	0.0267**	0.0259**	0.0055	0.0043
	(0.0039)	(0.0039)	(0.0063)	(0.0063)
tempagesq	-0.0340**	-0.0331**	-0.0033	-0.0018
	(0.0052)	(0.0052)	(0.0089)	(0.0088)
tempimm	-0.0103	0.0017	0.0684	0.0736
	(0.0528)	(0.0528)	(0.0614)	(0.0606)
tempedlow	0.0762**	0.0752**	0.0340+	0.0237
	(0.0144)	(0.0144)	(0.0184)	(0.0181)
tempedmid	0.0541**	0.0511**	0.0176	0.0082
	(0.0162)	(0.0162)	(0.0186)	(0.0184)
unemrate	-0.0009	-0.0009	0.0017	0.0025
	(0.0018)	(0.0018)	(0.0027)	(0.0027)
tempten	0.0036	0.0036	0.0128*	0.0134*
	(0.0045)	(0.0045)	(0.0060)	(0.0059)
temptensq	-0.0004	-0.0004	-0.0015**	-0.0015**
	(0.0004)	(0.0004)	(0.0005)	(0.0005)
tenure	0.0128**	0.0124**	0.0163**	0.0138**
	(0.0025)	(0.0025)	(0.0035)	(0.0034)
tenuresq	-0.0009**	-0.0009**	-0.0012**	-0.0010**
	(0.0001)	(0.0001)	(0.0002)	(0.0002)
year dummies?	yes	yes	yes	yes
occupation dummies?	no	yes	no	yes
industry dummies?	no	yes	no	yes
Sample size	17882	17882	9615	9615
r squared	0.164	0.171	0.167	0.198

+ p<0.10, * p<0.05, ** p<0.01

Immig is defined as being either foreign born or a noncitizen.