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**Working Paper**

# Does Employment Protection Reduce the Demand for Unskilled Labor?

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DISCUSSION PAPER SERIES

IZA DP No. 1290

# Does Employment Protection Reduce the Demand for Unskilled Labor?

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Forschungsinstitut
zur Zukunft der Arbeit
Institute for the Study
of Labor

# Does Employment Protection Reduce the Demand for Unskilled Labor?

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Discussion Paper No. 1290
September 2004

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# ABSTRACT

## Does Employment Protection Reduce the Demand for Unskilled Labor?*

Perhaps it does. We propose a model in which workers with little education or in the tails of the age distribution – the inexperienced and the old – have more chance of job failure (mismatch). Recruits' average education should then increase and the standard deviation of starting age decrease when strict employment protection raises hiring and firing costs. We test the model using annual distributions of recruits' characteristics from a 1975-95 panel of plants in Belgium, the Netherlands, Italy, the UK and the US. The model's predictions are supported using the Blanchard-Wolfers index of employment protection as well as our alternative index.

JEL Classification: J21, J83

Keywords: employment protection, labor demand, unskilled workers, firm panel data

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## 1. Introduction

This paper presents a model of the impact of employment protection legislation on the demand for unskilled labor, which we then test using data from a panel of firms. The model relies on lack of information on the part of firms about worker characteristics. This lack of information, combined with a posited greater likelihood of "failure" on the job of unskilled workers, makes firms choosy about hiring unskilled workers when employment protection laws raise dismissal costs. Acemoglu and Angrist (2001) first put forward this type of model to analyze the impact of disability discrimination legislation on the demand for disabled workers. The model provides a rationale for why employment protection *matters more* for less skilled workers[1]. The paper is therefore primarily concerned with the distribution of employment opportunities rather than overall employment levels.

Our concern with distributional issues follows the changing emphasis in the literature. The literature on labor demand began by taking a homogeneous labor assumption (starting with Nickell, 1986, then developed for example by Bentolila and Saint-Paul, 1994, and more recently by Blanchard and Wolfers, 2000). Such an assumption rules out looking for the differential impact of employment protection laws across skill groups. However, evidence has begun to mount that employment protection laws impact adversely on young workers (Scarpetta, 1996) and the long-term unemployed (OECD, 1999; Nickell and Layard, 1999, 3063), even if the majority workforce are little affected. A series of papers by Kahn (2000), Jimeno and Rodriguez-Palenzuela (2002), and Bertola et al (2004) analyzing 20-30 year panels of OECD country data find that young workers and older workers, particularly males, fare less well in high unionization and employment protection environments. Moreover, Modesto (2004) has provided a formal treatment of how youth and old-age unemployment may increase with employment protection legislation, particularly if wages are inflexible – or even increase due to increased insider power of incumbent workers. Our paper offers an empirical contribution to this debate, using firm-level data.

Our model emphasizes the inflexibility caused by employment protection law rather than that resulting from union bargaining (emphasised in Kahn, 2000, and Bertola et al, 2004). According to our management informants, union density within the firms had generally been stable over time; hence this factor can be taken as a fixed effect. In the empirical work, we control for union density at the national level, and for movements in average manufacturing wages (to pick up the pressure of national collective agreements). Whether our emphasis is appropriate can be tested on an empirical basis. In fact, as will be seen, our predictions for the employment protection variable are generally borne out.

The firms in our dataset have subsidiaries both in highly regulated countries such as Italy and Belgium, and less regulated countries such as the UK and US. Figure 1 contrasts the labor environments in two of these countries, Italy and the UK. There are striking differences in job opportunities for the unskilled worker categories such as the under 25s and the over 60s. As can be seen, the 20-24 unemployment rate in Italy is around 30%. This figure is three times that in the UK, despite Italy's large training and subsidized work programs (bottom row). Correspondingly, the 20-24 employment/population rate in Italy is currently only about 40% and falling, while that in the UK is around 70%. Similarly, at the other end of the age

[1] In a similar vein, Koeniger et al (2004), put forward a model in which employment protection reduces the firm's outside option, so permitting unions to negotiate higher wages. Higher employment protection costs for unskilled than skilled workers should then mean that strict employment protection helps unions to reduce the skilled/unskilled wage differential – which they find. Our model provides a rationale for higher employment protection costs for unskilled workers.

spectrum, among over 60s, Italy offers few jobs, with an employment/population ratio of only 20%, compared to the UK's 35% (and a US figure of 45%). Admittedly, Italian over-60s might be content not to work – their low unemployment rate indicates little search for work, and Italy has large state-funded early retirement programs (OECD, 1996, 208). But the marked lack of jobs for older workers still needs explanation. The labor market in Italy evidently works well for prime-age groups, but not for others. This type of relatively uneven performance is the motivation for our paper.

Our use of firm-level data to test for employment protection effects is a form of "insider econometrics", to use Ichniowski and Shaw's (2003) term. We use fieldwork interviews to generate a detailed understanding of recruitment in four multinationals with subsidiaries in the US and several European countries (see Daniel and Siebert (2003) and Morton and Siebert (2001) for similar cross-country company comparisons). This fieldwork is combined with detailed econometric hypothesis testing using the firm-level datasets we assemble. In particular, by tracking these firms over approximately a 20-year period we gain both time and country variation, which allows a fixed effects econometric framework to hold unobservables constant. The more usual aggregate country comparisons have problems of consistently defining employment and unemployment (for a survey, see Addison and Texeira, 2003). Here, our data on recruits' characteristics are perhaps more consistent across countries, since age and education are easier to define and measure, and company personnel records form a common statistical source.

A study such as ours has to face the difficulties of measuring the force of employment protection legislation (on which, see Bertola et al, 2000). Further, we desire a time-varying measure. A well-known recent measure is that constructed by Blanchard and Wolfers (2000), based on the OECD's (1999) scoring of the strictness of employment protection legislation in member countries. We begin with this measure, as do Bertola et al (2004) and Koeniger et al (2004). However, sensitivity tests need to be conducted. For these tests, we have elected to use our own measure, which has been constructed independently of Blanchard and Wolfers, and uses somewhat different assumptions[2].

To preview our results, we find that strict employment protection – both on the Blanchard-Wolfers measure and on ours – is associated both with higher average education and with less dispersion in the starting age of recruits. We interpret these results to mean that firms become choosier about hiring from the less educated as well as the young and old age groups, who are more of a risk than prime-age workers, when employment protection raises dismissal costs. The latter result might underlie the UK's comparatively healthy age dispersion of jobs shown in Table 1.

Our plan is as follows. In the next section, we present the model of labor demand, and discuss the econometric specification. In the third section, we discuss our data. In the fourth section we present the regression estimates. The final section offers a summary and conclusions.

[2] There is also a time-varying measure of the strictness of employment protection legislation based on employer views (Di Tella and McCulloch, 2004), but this series is short, 1984-1990. Alternatively, there is a series on product market regulation (Nicoletti and Scarpetta, 2003) – which is related to labor regulation but, of course, is not the same.

## 2. The Model

Theory. We analyze the demand for unskilled and skilled production labor (excluding management grades). We proxy 'skill' by two characteristics: the starting age of recruits, A, and the education of recruits, E. These are the characteristics for which data is consistently available from personnel records in our study plants. We do not have data on the firms' capital stock, and it is simplest to assume that capital stock differences are predetermined, and absorbed into the firm fixed effect. [3]

Let us begin with a revenue function, R, which we define as price, p, multiplied by efficiency units of labor, $\ell$. Hence total revenue is $R = p\ell$. We assume that efficiency units of labor are linearly related to the number of workers, L, multiplied by a worker efficiency function g (A, E), as follows:

$$\ell = Lg(A, E) \qquad (1)$$

with $g_A \gtrless 0$, $g_{AA} < 0$, $g_E > 0$, $g_{EE} < 0$, and $g_{AE} > 0$ (to allow substitution between A and E).

Let us explain our assumptions for the efficiency function, taking first the starting age (A) argument. We make the efficiency of labor an inverted-U function of recruits' starting age. Our reasoning is that young recruits (under 25) have less experience, and hence can be expected to be less skilful, given education. Older recruits (over 55) also have disadvantages, perhaps out-of-date skills or negative selection, in that an older worker looking for a job may have proved unsatisfactory in previous work. It is thus possible to have "too much experience". This assumption is motivated by the fall-off in employment/population rates for older workers shown in Table 1 – but, in any case, it is generally accepted that most firms do not hire older workers (for discussion, see Heywood et al, 1999). The worker efficiency function, measured in product price (p) terms, is given in Figure 1, with $g_A$ starting > 0, and becoming < 0, and with $g_{AA} < 0$. Admittedly, careful hiring and training procedures within the company can reduce the curvature of the g function, at a cost. Also, in some cases, for example in firms which emphasize training, the curvature will perhaps be more pronounced because older workers are less trainable. All we need for our argument is some degree of curvature, so that there is a starting age range which the firm prefers as in Figure 1.

As for the role of education in the efficiency function, the efficiency of labor obviously increases with education, though at a decreasing rate, as shown in Figure 2. Again, as with starting age, the impact of education on worker efficiency will depend on the firm's hiring and training expenditures, which we set to one side for the moment.

Consider next the element of labor force adjustment: quits, dismissals, layoffs (redundancies), and hires. To simplify, we assume the firm is in a steady state with $L_t = L_{t-1}$, so there are no layoffs, and the hiring rate, h, just balances the quit rate, $\delta$, plus the dismissal rate, $\theta$, i.e.:

$$h = \delta + \theta \qquad (2)$$

Dismissals are central to our model. All workers face some probability of failing, and hence of dismissal. Failure can be thought of as stemming from mismatch between the worker and the job. However, since more is known about skilled workers – who are better

[3] Admittedly, an increase in capital usage over time, in response to employment protection, could also account for recruitment of more skilled workers. Our empirical findings for the employment protection variable would then simply have to be interpreted as reduced forms.

educated and prime-age – their probability of failure is lower. Hence, dismissal rates should be lower for skilled than unskilled workers. Accordingly, we make the dismissal rate a function of A and E, i.e.: $\theta = \theta(A,E)$. We assume $\theta$ follows a U-shape, being least during the prime age range, so $\theta_A >< 0$ and $\theta_{AA} > 0$. On the other hand $\theta_E < 0$ because the more educated have a lower failure probability.

Dismissals bring firing costs, F, and expected firing costs are $\theta F$. We will assume F is the same throughout the production worker group, whether skilled or unskilled. (F costs would, of course, be higher for management grades but these are not included in our study.) Since unskilled workers are more likely to be dismissed, their expected firing costs will therefore be greater than those of skilled workers, and employment protection legislation will increase such costs more for unskilled workers.

Hiring costs, H, are also relevant. Hiring and firing costs tend to move together, because firing costs bring 'shadow' hiring costs (Bentolila and Bertola, 1990, 391). Thus, as it becomes more difficult to fire workers, more must be invested in hiring costs – choosing the appropriate worker. This effect is likely to be greater for unskilled workers, who are untried workers without a track record, once employment protection legislation makes it difficult to substitute an unsuitable worker with a suitable one (Saint-Paul and Kugler, 2000, 8). In other words, strict employment protection raises the possibility that a job can become permanently filled by a sub-standard worker. This possibility imposes an opportunity cost on the firm: the lost value of the option of filling the vacancy with an able worker. This type of expected opportunity cost must be higher for untried candidates without track records or qualifications.

In practice, therefore, we can think of employment protection legislation as increasing the expected sum of hiring and firing costs, $\theta V = \theta(H+F)$. The dashed lines in Figures 1 and 2 show worker efficiency net of expected hiring and firing costs. Employment protection legislation will increase the divergence between the two lines by increasing V.

Finally, wages will also be a function of A and E, i.e. $w = w(A, E)$. The wage function need not be the same as the worker efficiency function, though the two will be related. Obviously, wages increase with education, so $w_E > 0$. However, equity considerations, or trade union pressures, are likely to prevent much variation of wages with recruits' starting age, hence for simplicity, we assume $w_A = 0$. For simplicity, we have made wages independent of V – though as Modesto (2004) and Koeniger et al (2004) point out, higher V could be associated with higher w since protection of incumbent workers raises their bargaining power. We will return to this subject below.

The firm aims to choose L, A, and E to maximize the present value of profits, N, defined as the value of output minus wage, hiring and firing costs. We write the objective function as follows:

$$N = \sum_{t=0}^{\infty} \beta^t [pL_t g(A_t,E_t) - w(A_t,E_t)L_t - F\theta(A_{t-1},E_{t-1})L_{t-1} - H(\delta + \theta(A_{t-1},E_{t-1}))L_{t-1}] \quad (3)$$

where $\beta = (1 + r)^{-1}$ is the discount factor with r = discount rate, and p = product price. We assume $L_0 = 0$. The third term in square brackets gives total firing costs which depend on the number of workers dismissed last period, $\theta(A_{t-1},E_{t-1})L_{t-1}$ times F. The fourth term gives hiring

costs, which depend on the quit and dismissal rates (substituting from equation (2) above) times H.

We derive the first order conditions for equation (3) following Acemoglu and Angrist (2001, 922), with firms immediately adjusting to steady state employment levels, so $L_t = L$, $A_t = A$, and $E_t = E$ every period. We also assume that it takes a year for any H or F costs to arise. We can then simplify (3):

$$N = pLg(A,E) - w(A,E)L + \frac{\beta}{1-\beta}[pLg(A,E) - w(A,E)L - F\theta(A,E)L - H(\delta + \theta(A,E))L] \quad (4)$$

since $\sum_{t=1}^{\infty} \beta^t = \beta/(1-\beta)$. Differentiating (4) with respect to L gives:

$\partial N/\partial L = (pg(A,E) - w(A,E))/(1-\beta) - [F\theta(A,E) + H(\delta + \theta(A,E))]\beta/(1-\beta) = 0$, so the employment level chosen satisfies the condition:

$$pg(A,E) = w(A,E) + \beta[F\theta(A,E) + H(\delta + \theta(A,E))] \quad (5)$$

As can be seen, L drops out of this condition, because of the linear form we have given the revenue function. The conditions for A and E, given below, are therefore in per-worker terms, with L determined outside the model.

For the age choice, we have:

$$\partial N/\partial A = L(pg_A - w_A) + \frac{\beta}{1-\beta} L(pg_A - w_A - \theta_A V) = 0, \text{ where } V = F + H.$$

Hence, $pg_A = w_A + \beta\theta_A V$. (6)

In other words, the marginal revenue product of labor by starting age must equal the "full" marginal cost of labor by age, including expected hiring and firing costs. In terms of Figure 1, where we have chosen a simple flat wage-age line, the choice of starting age, A*, is given by the maximum of dotted $pg(A, \overline{E}) - \beta\theta V$ line.

There is a similar condition which the optimum education choice, E*, must meet:

$$p\, g_E = w_E + \beta\theta_E V. \quad (7)$$

Figure 2 illustrates the position.

It is also necessary to consider the possibility of substitution between education and starting age, as shown in Figure 3. The positively sloped section of the isoquant indicates the region where starting age is too high, reducing worker efficiency. The firm will always aim to operate to the left of this point, the "ridge line" – though in practice the line will not be well defined, since the worker age-efficiency function in Figure 1 will have a broad top. With no employment protection, cost minimization requires the factor combination indicated by point X.

Now let us consider the impact of employment protection – higher V costs – on selection of worker characteristics. From (6) we see that V can increase or decrease the marginal cost of older workers, depending upon whether $\theta_A$ is $< 0$ or $> 0$ (see Appendix 1). But since $\theta_A$ is likely to be small, we would not expect employment protection legislation to much affect the <u>average</u> starting age that management selects – though there may be some fall since education substitutes for age, and education is likely to increase, see below. However, strict employment protection legislation will reduce the <u>dispersion</u> of starting ages. This effect can be seen most simply from Figure 1. Raising V increases the curvature of the

dotted pg $(A, \overline{E})$ – $\beta\theta V$ line. With a low V, the line is not very curved, so the firm will be indifferent about A, because the penalty associated with worker failure due to choice of the wrong starting age, $\theta_A$, will be low. Now suppose V increases. When this happens, it will become more important for the firm to 'get it right', that is, to choose specifically the prime age group for which $\theta V$ is lowest. Thus, high V should reduce the dispersion of A.

The position is different for education, in equation (7). Here, we see that an increase in V lowers the marginal cost of a more educated worker, because $\theta_E$ is negative (Appendix 1). Stricter employment protection legislation should therefore unambiguously tilt management decisions in favor of more educated recruits. Consequently we predict an increase in recruits' average education as V increases. Moreover, as Figure 3 shows, an increase in V is likely to make the full cost of education cheaper relative to starting age, leading to substitution of education for starting age, and a movement from X to Y. At the same time, unlike the case for starting age, there is no reason to expect increases in V to reduce the dispersion of recruits' education. Dispersion depends on the penalty associated with making the wrong education choice, which in turn depends on the curvature of the g, $\theta$ and w functions. However, the curvature of these functions does not depend upon V.

Our predictions for the impact of employment protection legislation on our two characteristics, A and E, thus form an interesting contrast. Strict employment protection legislation (high V) should leave the average A of recruits undisturbed, but lower the dispersion of A. Exactly the converse should be true of E.

Admittedly, in deriving these results we have adopted certain simplifications. In particular we have ignored the possible countervailing impact of high V in raising worker efficiency – in particular, the possibility of an H argument in the g (A, E, H) function. In other words, greater hiring expenditures, H, could so stimulate (via better choice) the productivity of uneducated workers as to offset the costs associated with their higher probability of failure. However, it is implausible that such a full offset should occur. If it did, why did not management choose higher H in the first place, without being forced by employment protection legislation? In any case, we subject the matter to test below.

Specification. We form the observations for the average and dispersion of recruits' starting age and education as follows. The analysis relates to male recruits whose contracts become permanent/open-ended within a year at the plant, since our best continuous data series relate to this group. For each of the eleven plants in each year (the plants are described in more detail below), we calculate the mean and the standard deviation of recruits' starting age and education. The data points thus represent average behavior for each plant in each year.

This method reduces the data on some 2,400 recruitment events to approximately 140 plant-time data points, depending upon missing values. We use arithmetic weights in our estimation procedures, with weights based on the number underlying each plant's distribution in that year, to allow for the fact that sometimes the number hired in a year is small.

Our statistical model in its general form is:

$$Q_{it} = \sum_{i=1}^{10} a_i + \sum_{i=1}^{10} b_i t + \sum_{i=1}^{10} c_i EPL_{it-1} + \sum_{i=1}^{10} d_i X_{it-1} + e_{it} \qquad (8)$$

Where $Q_{it}$ = the average or standard deviation of recruits' education or age in the i-th plant in the t-th year, i = 1,2...,10, t = time trend; $a_i$ = constant term for the i-th plant; $EPL_{it}$ =

employment protection index in the i-th plant's country and t-th year; $X_{it}$ = a vector of other controls; $e_{it}$ = the error term. This model is completely unrestricted, with different coefficients for each plant. We then use F-tests to test whether it is permissible to restrict some or all of the coefficients to equality. This F-test procedure can be used to test whether, for example, the coefficients of the UK plants as a group can be restricted to equality – or whether some other grouping is permissible, for example, of sister-plants.

A restricted form of equation (8) is

$$Q_{it} = \sum_{i=1}^{10} a_i + bt + cEPL_{it-1} + dX_{it-1} + e_{it} \qquad (9)$$

This is the basic fixed effects form, with only the constant fixed effect term, $a_i$, differing among plants, all other coefficients being the same. F-tests generally show that we can accept the restrictions implicit in (9).

In equation (9), the fixed effects, $a_i$, are meant to account for omitted variables specific to the firms, but which are constant over time. For example, plants in richer countries such as the US should have access to a supply of better-educated workers, which will obviously affect hiring decisions. By contrast, the time trend variable, t, is intended to capture effects specific to each time period, and the same across firms. An example of such effects is the reduction in unskilled labor demand such as might result from international trade competition and/or skilled-labor-using technical progress, to which all our plants have presumably been subject.

A further point is the simultaneity of starting age and education. Education and starting age will tend to be substitutes, at least when younger workers are hired. (The standard deviations of starting age and education are simpler – we can take these variables to be independent of each other, and also of the average values of starting age and education.) In fact, as we shall see, the system seems to be recursive[4]. First, the firm chooses recruits' education independently of starting age. Then, second, the firm chooses starting age dependent on education, and the two turn out to be good substitutes. To address this issue, we use simultaneous equations techniques to estimate the average age and average education equations.

Measuring employment protection.

As we have already noted, it is difficult to capture a many-dimensioned force such as employment protection in a single time-varying variable (Bertola et al, 2000). Firing costs are influenced by many rules governing unfair dismissal, layoffs for economic reasons, severance payments, minimum notice periods, administrative authorization for dismissals, and prior discussion with representatives of unions or labor market administrations. In addition, for the US in particular, there is judge-made law raising the costs of dismissal (Autor, 2003), even though there are weak statutory provisions.

Nevertheless, some progress has been made. Specifically, Blanchard and Wolfers (2000) have constructed cardinal measures of the strictness of employment protection legislation for several countries; including the ones we are interested in here. Their measure is based on the OECD's country rankings of strictness of employment protection in the 1980s

[4] A recursive model may be consistently estimated using equation-by-equation ordinary least squares (Greene, 2003, 397), but not if the covariance matrix of the equation disturbances is non-diagonal, as appears to the case for some of our specifications.

and 1990s (OECD, 1999). They then use the index developed by Lazear (1990), who quantified firing costs as the amount of severance and notice period measured in monthly wages owed to a dismissed worker after 10 years of service, to stretch the series back to the 1970s.

The resulting index for our time period is shown in Figure 4. As can be seen, it ignores possible increasing US case-law strictness of protection. It also ignores the UK's decrease in strictness under Thatcher's Conservative administration. However, on the good side, the index aims at a cardinal measure of dismissal costs (it does not simply rank countries), it covers the countries and the time periods we need, and it is independent of our own data calculations. Therefore, we use this index (as have Bertola et al (2004) and Koeniger et al (2004)) as a foundation.

We will also subject the results to sensitivity tests, and in particular demonstrate results using our alternative index of employment protection. Our index is based on somewhat different assumptions to the Blanchard-Wolfers index. It is also based on the OECD (1999) method, but for our US states, it incorporates both changes in legislation, and in relevant court practices. For the US, exceptions from the employment-at-will doctrine introduced in the majority of US states throughout the past two decades may increase employment protection – not via legislation but rather the threat of potentially costly litigation – which we have calibrated using the Rand study on termination litigation in California (Dertouzos et al., 1988).[5]

The resulting index is shown in Figure 5. The marked difference between the two indices in the treatment of the US can be seen. Our index gives strictness of employment protection at the state level, and also shows an increase of employment protection based on case law, while the Blanchard-Wolfers index does not. There are other detail differences as well. Thus, allowing for country fixed effects, our index only explains 0.46 of the Blanchard and Wolfers index, so the two indices are reasonably different.

The controls. Let us now turn to the control variables, X. In the first place, our plants produce four different products, and we may expect these to have different requirements for high skilled relative to low skilled workers. The plant fixed effect term helps to control for this factor. In addition, our plants can be formed into groups producing the same product (as subsidiaries of the same multinational), which allows further control.

A wage variable is also needed. Although in the development of our model we abstracted from wage effects, wages may rise if employment protection shelters incumbent workers, and so unions push up their wages (e.g., see Modesto, 2004). Since it is unskilled workers who are at risk, a measure of wage compression would be best, but this variable is not available over time. Hence, we simply include the average hourly manufacturing wage. Increases in this variable should have employment protection-like effects, causing firms to

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[5] Our index is constructed based on the OECD (1999) index for individual dismissal of workers with regular contracts, applying OECD weights. It includes scores for procedural inconveniences (procedures and delay to start notice) notice and severance pay for no-fault individual dismissal and difficulty of dismissal (definition of unfair dismissal, trial period, compensation and reinstatement). It is then combined with an index of the strictness of regulation of temporary employment, again based on OECD (1999), and smoothed over time.

become more choosey, and so increasing education requirements, and reducing the standard deviation of starting age.

Similarly, strong unions are likely to play a role in promoting and enforcing employment protection and other labor regulation, as well as in pushing up wages generally, all of which will tilt labor demand in favor of skilled workers. Again, this effect should be seen in a decrease in the starting age standard deviation, and an increase in the education average. As noted above, union density at the plant level is not available over time, so we make do with national-level figures.

We also include the tax wedge for the country (total taxes divided by GDP), on the argument that when taxes increase labor cost, this may reduce the relative demand for unskilled workers. This argument requires wage inflexibility; otherwise tax increases will be shifted back toward the worker, with little effect on labor demand.

A further variable is unemployment. For example, it might be that in bad times, the relative demand for unskilled workers decreases, since firms tend to hoard skilled labor then (Reder, 1955; Devereux, 2000). Conversely, in good times, skilled labour takes time to train, so firms must take unskilled workers, and the relative demand for unskilled workers increases. However, for our firms there might also be counteracting cyclical shifts in unskilled labour supply[6]. Our firms are mainly in non-durable manufacturing which has less cyclical employment variation than durable (McLaughlin and Bils, 2001). Thus, in slack times many unskilled workers will be searching for a job in non-durables (they are laid off from durables), causing a relative fall in unskilled wages and maintaining unskilled recruitment in non-durables. To allow for such effects, we incorporate as controls both the plant's employment deviations from trend, and also the national unemployment rate.

**3. The Data**

The sample includes data from four major manufacturing multinationals, most with plants in the US, the UK, and a country in continental Europe. The companies were chosen because they had subsidiaries in both regulated and unregulated countries, were large enough to regularly hire workers, and had 15-20 years past personnel record data. The industries involved are ice-cream manufacturing (Italy, the UK, and Missouri for the US), distilling (Italy, the UK, and California for the US), food processing (Netherlands, the UK, and Maryland for the US) producing mainly margarine, and pharmaceuticals (Belgium and the UK) producing penicillin products (for details, see Daniel and Siebert, 2003).

Basic employment and labor costs data for the resulting sample are given in Table 2. As can be seen from the table, labor costs per production worker tend to be lowest in the UK plants. The pharmaceuticals pair shows the biggest difference; with labor costs in the Belgian plant being more than twice UK costs, due to higher Belgian labor taxes and extended collective agreements. Nevertheless, unit labor costs are similar in the two plants ($165 to $175 per $000 sales) indicating that the Belgian plant is securing a level of labor productivity which is twice as high as that of its UK counterpart. We would expect such differences to feed through to the hiring process, with the Belgian plant concentrating more on prime-age, educated workers than its UK counterpart.

[6] For an early discussion of how supply shifts may affect relative wages and employment, see Perlman (1958).

Table 2 also shows unit labor costs appear to differ quite widely for the distillers' plants (from \$0.049 to \$0.099 per liter), and food processing plants (\$49 to \$74 per ton), which may undermine the competitive assumption (the more expensive plants should have been eliminated over time). However, it is difficult to calculate the labor productivity factor underlying unit labor costs. Moreover, the exchange rates used are problematic. Therefore, we believe that these differences should not be taken as strong evidence against the competitive assumption.

Information on mean values of the labor demand variables is given next, in Table 3. The first row shows that the data period is about 20 years, 1975-95, in most cases, though the ice-cream plant in Italy yielded only 12 years of observations, 1985-97. Hence, we have achieved a serviceable time series. Admittedly, the following rows show that the average number of hires per year is quite low in some plants, so the means and standard deviations will be unreliable[7]. To help circumvent this problem, as noted above, we use the underlying number of observations as arithmetic weights.

As for the starting age variable, the mean values are in the late twenties for most plants. Thus, school-leavers are generally not hired[8]. The high starting age values show the emphasis on previous experience amongst this group of large plants. Nevertheless, the standard deviation of starting age is smaller in plants in Italy, the Netherlands and Belgium, where employment protection legislation is stricter, as expected.

The education variable rows show the US plants to have highest average education, at around 12 years, almost 2 years higher than the Italian plants. At first glance, this pattern runs counter to our hypothesis that strict employment protection should result in more emphasis on education. However, country laws on the school-leaving age, which is low in Italy, could affect the average – as well as differences in country wealth. These factors should be picked up in the fixed effect (in addition, there have been changes in school-leaving laws, for which we include a school-leaving age dummy). Table 3 also shows the standard deviation of education of new hires, which varies from 1 to 2.5 years. There is no particular pattern in the cross section, comparing countries, nor do we expect any. We will now explore these relationships more systematically.

**4. Results**

The main results: The main results are given in Table 4. For each dependent variable, we report results using the Blanchard-Wolfers index as well as our index. Since it is necessary to pick up plant-specific unobservables, as noted above, we include fixed effects in all specifications. F-tests show that we can assume the same coefficients for all plants, apart from the fixed effects. Although we experimented with various groupings based on plant ownership, these were not significant.

Estimation of the standard deviation equations is by ordinary least squares using underlying observations as arithmetic weights. However, because we expect starting age and

[7] For consistency over time, we concentrate on permanent males, defining "permanent" to include workers whether hired on a temporary basis or not, who became subsequently employed on an open-ended basis within a year. Where such hires fell below 2 in any year, we recorded a missing observation.

[8] However, an exception is the Italian distilling plant, which has starting age averaging only 23.7. Special factors seem to be operative in this plant, which recruits extensively among relatives of current employees. Such extra knowledge of applicants could allow age and education criteria to be lowered. Again, we rely on the fixed effects term in (9) to control for these special factors.

education to be determined jointly, the average age equation and the average education equation are estimated as simultaneous equations by three stage least squares, again including underlying observations as arithmetic weights.[9] In fact, as shown by the Durbin-Wu-Hausman test, average age is endogenous in the education equation (column 5) while average education is exogenous in the age equation (column 1) using the Blanchard-Wolfers index. It thus appears that we have a recursive system, with the personnel office making a decision first on an applicant's education, which is then traded off against starting age.

On the other hand, using our own index (columns 2 and 6), both education and age appear to be exogenous, which is puzzling since we expect simultaneity. However, at least this result means that we can use ordinary least squares, which provides a cross-check on the results using three stage least squares. (As a further cross-check, we also give in Appendix 2 reduced form equations for average starting age and education. ) The important point, as we will see, is that average education responds significantly positively to employment protection, whatever the specification.

Findings for employment protection are given in the top two rows. For the average starting age variable using the Blanchard-Wolfers index (column 1), the coefficient on employment protection, -1.60, is negative, but insignificant. If we use our own index (column 2), the coefficient on employment protection is also negative and insignificant, -0.50. These findings are in accordance with our model, which does not predict a strong link between employment protection and recruits' average age. At the same time, we see that education is strongly substitutable for starting age, which is plausible. This result is most marked using the Blanchard-Wolfers index: a one -year increase in recruits' education is associated with a 3.25-year decrease in starting age. Personnel offices evidently trade off education against starting age.

Turning to the standard deviation of starting age, we see the predicted contrast. Employment protection legislation significantly reduces the standard deviation of starting age in both specifications. This reduction is most marked using the Blanchard-Wolfers index (column 3), with a coefficient of -9.12. This reaction is in line with our model's predictions. The elasticity, taken at the means, is –0.99 (= –9.12×1.0/9.2). Thus, moving Italy's employment protection level from 4 down to the average of 1, a change of –120% (= (1-4)/½(1+4)), would imply an increase of almost 120% in the standard deviation of starting age. Such a change would widen Italy's age standard deviation to about 11 years – approximately UK levels (Table 3). Using our own index (column 4), employment protection still significantly reduces the standard deviation of starting age, though to a lesser extent. Here the elasticity is –0.53 (=–3.73×1.3/9.2), which implies moving Italy's employment protection level down to the average, would widen Italy's age standard deviation to about 8 years.

Turning next to average education, there is the predicted opposite pattern in both specifications. Using the Blanchard-Wolfers index (column 5), average education responds significantly to the employment protection index, with a coefficient of 2.35. The elasticity is 0.21 (= 2.35×1.0/11.4). This elasticity is smaller than that for the standard deviation of starting age, but this is appropriate since education levels cannot vary much. Thus, moving Italy's employment protection level once again by –120% would imply a reduction of 25% in

[9] The instrument we used for education was the school-leaving age variable, which can reasonably be excluded from the age equation. The instrument we used for starting age was the average age of the company's worker stock which, for its part, can reasonably be excluded from the education equation.

the Italian firms' education levels, which is quite enough (given Italy's already low education levels). Using our index (column 6), the elasticity is, 0.17 (=1.47×1.3/11.4), implying a reduction of about 20% in the Italian firms' education level if Italy's employment protection level moves down to the average. The positive link between employment protection and average education remains in the reduced form specification (Appendix 2).

Finally, the last two columns show that the standard deviation of education does not respond significantly to employment protection using either measure. This result is also consistent with our model, which makes no predictions for the standard deviation of education.

The pattern of results for the remaining variables also gives confidence in the model, though the results are somewhat stronger using the Blanchard-Wolfers index. Thus, when using this index, the time trend variable (column 3) shows that the standard deviation of starting age has been trending downward, at –0.34 years per year. At the same time, the average education of recruits (column 5) has been trending significantly upward, at 0.08 years per year. (The education trend is not simply a consequence of the school-leaving age increase that occurred in some countries during the period, since we have the school-leaving age control, which is significant.) These results are sensible. They indicate that management has been becoming choosier over the years, raising education, and more tightly defining starting age, which could reflect global competition and/or skill-using technical progress raising skill requirements.

The main other significant controls are for business conditions, that is, the unemployment and deviations from employment trend variables. Both variables point to a rise in recruiting standards when business is good. Thus, we see that lower unemployment is linked to an increase in average education. Similarly, a positive employment deviation from trend is linked to a reduction in the standard deviation of starting age. The simple argument, as we noted above, is for good business conditions (low unemployment) to favor lower recruitment standards, as firms run out of hoarded skilled labour. However, as we also noted above, supply shifts could explain our contrary findings, since few unskilled workers may be looking for jobs in non-durable manufacturing when unemployment is low, given the likely strong expansion in (better-paying) durable manufacturing at such times. Scarcity of unskilled workers could thus explain the apparent rise in recruiting standards in our sector when times are good.

The union density and the tax wedge variables produce mixed results. We argued above that both should have employment protection-like effects. Hence, we would expect negative coefficients on these variables in the standard deviation of starting age equation, and positive in the average education equation. Using the Blanchard-Wolfers index, union density significantly reduces the standard deviation of starting age, as expected, but it is insignificant in the average education equation. Using our index, the coefficient on union density is insignificant in both equations. However, our union density data are at the country level, not the plant level, since a time series of union density at the plant level is not available. Most of the plants had closed shops even in the 1990s (over 90% union density – see Table 2), so union density is likely to have been pretty constant over time, making our country series inappropriate. As for the tax wedge, this variable is significantly positively linked with average education as expected. However, it is insignificant in the age standard deviation equation, in both specifications. The effects of union density and the tax wedge are not clear cut, therefore, though some results go in the expected direction.

Finally, the manufacturing pay variable is trying to be significant in the expected directions. In other words, there are signs that personnel offices become more choosey when manufacturing pay levels are high, so recruits need to be more educated, and at the same time the standard deviation of starting age tends to fall.

Sensitivity tests: First, we exclude the three US plants, and look only at plants in the European subset of countries using both indices. This test aims to show whether these employment protection results are robust to a big change of sample. In addition, the test circumvents the possible problems of the Blanchard-Wolfers US employment protection index, which omits judge-made employment protection (see above).

Summary results are given in Table 5, Panel A[10]. We report the coefficients of both the employment protection indices and the time trend, the time trend being interesting because it should show a common country tendency towards a shrinking market for unskilled labor. Taking first Blanchard-Wolfers employment protection measure in the upper panel, the first row repeats Table 4's results for reference. The second row gives the results for the Europe-only sample. We see the same pattern: higher employment protection reduces the standard deviation of starting age, and increases average education, leaving the other two dependent variables unaffected. Hence our results are not much affected by the change in sample. Turning to our own measure, the pattern does not hold for the standard deviation of starting age equation for the Europe only sample. Hence, changing the sample size gives more confidence in the Blanchard-Wolfers measure.

Now turn to the time coefficients in the second row. The main feature here is that both employment protection indices produce a significant negative time trend for the standard deviation of starting age, and a significant positive time trend for average education for the overall sample. These findings indicate that management has been becoming choosier over the years, which could reflect global competition raising skill requirements, as noted above. However, using either index, there is no significant time trend in the Europe-only sample. Increasingly stringent hiring standards appear only to affect the US plants over this time period, according to this specification. This result could be due to the fact that US plants have had more room to raise their standards as global competition bites – standards in continental European plants already being quite high.

Our second test groups observations into 5-year averages, as shown in Panel B. The advantage of such grouping is that the number of observations underlying the dependent variables (means or standard deviations) is increased, so weighting is unnecessary. Also, reducing the number of datapoints per plant to 3 or 4 over time might more truly reflect the amount of information we have on employment protection, given that the indexes are static for long periods at a time. As can be seen, there is still a negative link between employment protection and the standard deviation of starting age, though the elasticity is reduced to about –0.2. Also, there is still a positive link between employment protection and education, this time with an increased elasticity of 0.64 to 0.85. However, in this formulation there is now a positive link between employment protection and the standard deviation of education, which is not expected by our theory. Nevertheless, the main results survive.

Our final test probes the employment protection indices. Specifically, we test for the strength of the association between both measures of employment protection and average

[10] Full details are available from the authors on request.

worker tenure in the plants. Our argument here is simply that high tenure in a plant should indicate high employment protection in that plant. The tenure variable is constructed using the average tenure of the workforce in a given year. Admittedly, average plant tenure changes only slowly in response to changes in the legal environment. Nevertheless, we think the exercise can still provide a check on the indices.

The results are given in Table 6, which shows a strong association between employment protection and plant tenure using the Blanchard-Wolfers index. The association is somewhat weaker using our alternative index. The coefficient on employment protection using the Blanchard-Wolfers index is 9.64 and highly significant, implying an elasticity, taken at the means, of 1.09 (=9.64x1.3/11.5). While labor turnover is not the focus of our inquiry, both equations behave quite well, with the expected negative link between plant employment deviations and tenure (positive deviations mean more hiring, and should lead to a lowering of average tenure), and positive time trend. However, the equation using our measure throws up a negative relation between union density and tenure, which is hard to explain. These results therefore give particular confidence in the Blanchard-Wolfers employment protection index.

**5. Conclusions**

This paper develops and tests a model in which employment protection costs can influence the type of labor characteristics demanded by employers. The model postulates that workers with little education, or in the tails of the age distribution, have more chance of failure (mismatch), and thus of imposing hiring and firing costs on the firm. Consequently, such workers are less likely to be recruited when strict employment protection raises hiring and firing costs. In particular, the model predicts that recruits' *average* education should increase and the *standard deviation* of starting age should decrease when employment protection becomes stricter. As Table 4 shows, our model's predictions are borne out using two alternative measures of employment protection, with the results being somewhat stronger using the Blanchard-Wolfers index. Strict employment protection indeed reduces the variability in starting age, and raises education requirements, independent of the employment protection measure used. Hence, there are strong indications that employment protection affects the steady-state distribution of recruits' characteristics – raising education requirements, and reducing starting age dispersion.

The adverse distributional impact of employment protection legislation, implied by our results, has become increasingly apparent from recent OECD country studies which disaggregate by age, as noted in the introduction. We offer a further disaggregation by education, and a different firm-based methodology to arrive at the same conclusion. Our study shows that employment protection is generally bought at a cost to the inexperienced, the old, and the uneducated – the have-nots.

Caveats and directions for further research must be noted. In the first place, our results depend upon the measurement of employment protection. We have shown that the employment protection indicators we use generally behave sensibly when explaining patterns of workforce tenure in our sample. That is, average worker tenure increases strongly in plants/time periods with strict employment protection (Table 6). The results also survive when we radically alter the sample by dropping all the US observations using the Blanchard-Wolfers index (but only partly using our index – see Table 5). Further work is necessary on measurement of employment protection. In particular, our post-1975 period has little variation in employment protection, since the system is essentially mature. A more powerful

analysis would be possible with data extending back into the 1960s, when the continental European countries were setting up their employment protection measures – though it would be an unusual firm which kept personnel record data extending back so far.

A second point arises about the generalizability of our results. Our data are reasonably consistent across countries because they come from subsidiaries of four multinationals, which impose a reporting uniformity. Also, all the subsidiaries are in a similar industrial sector, nondurables manufacturing, and we consider the hiring only of males into contracts which are open-ended (or become so within a year). These restrictions reduce extraneous noise. However, small firms, the service sector (including government), and the market for temporary workers are all excluded from consideration. There is work to be done to fill this gap. In particular, it is important to understand how the increasing use of temporary contracts allows some firms to contract around employment protection and experiment with unskilled groups. The large company, industrial sector, to which our results apply, is only a small part of the whole.

Nevertheless, our results show how employment protection legislation can influence recruiting decisions at the level of the company. Company time series studies bring their own difficulties in terms of missing information. However, the company is where the employment decisions are made. Rather than relying on country aggregates such as employment-population or unemployment rates, which are the subject of many factors, we are therefore able to provide specific tests of important employment decisions.

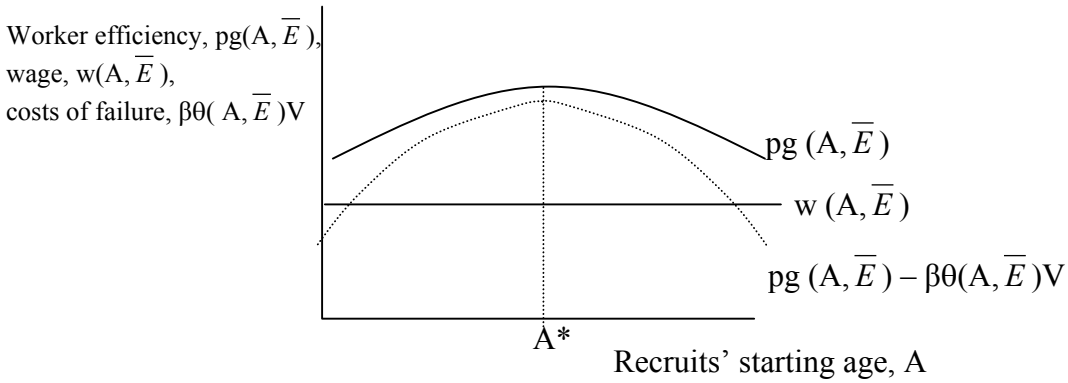


**Figure 1: Production worker efficiency and starting age (given education)**

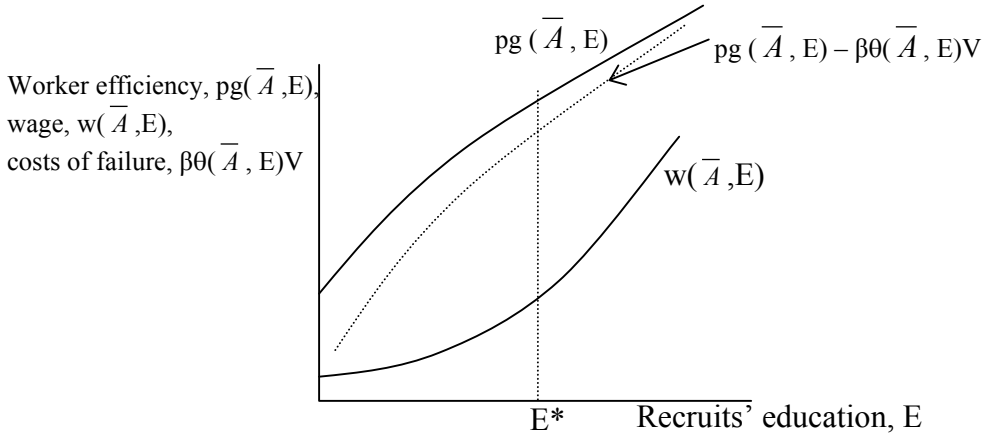


**Figure 2: Production worker efficiency and education (given starting age)**

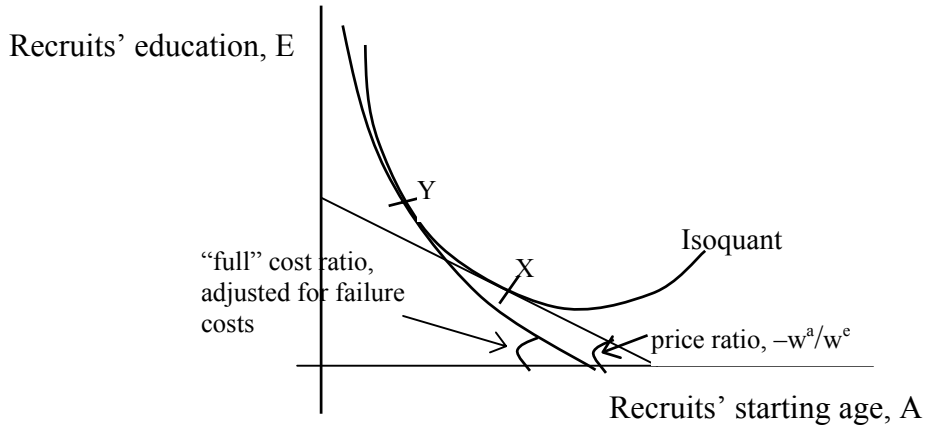


**Figure 3: Substitution between Education and Starting Age**

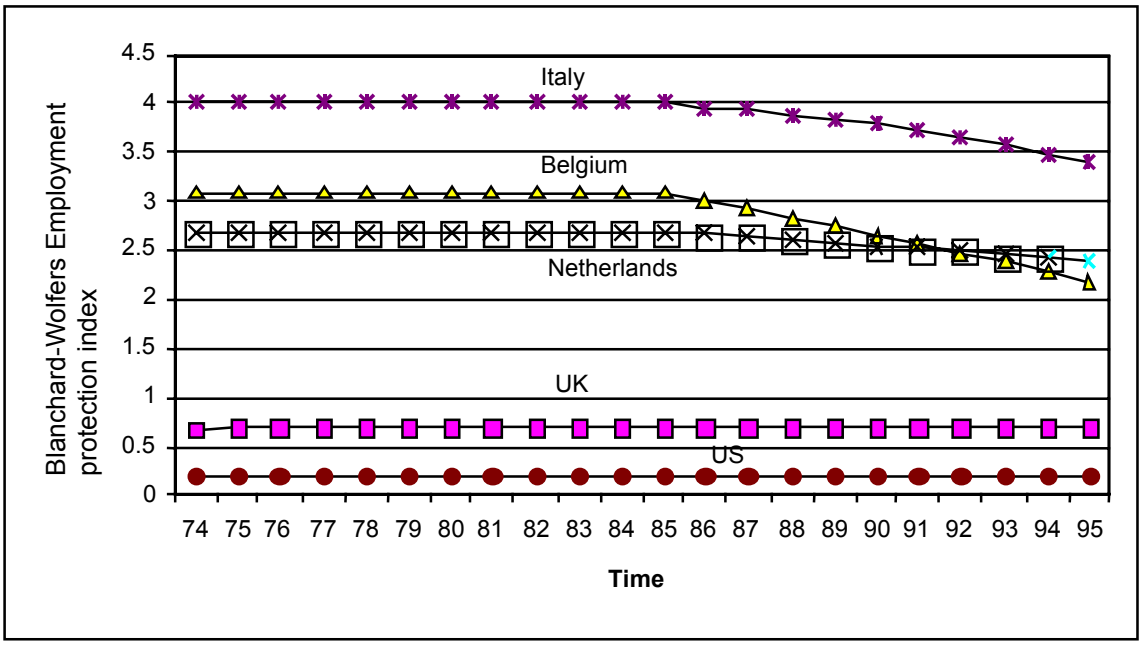


**Figure 4: Blanchard-Wolfers Employment Protection Index**

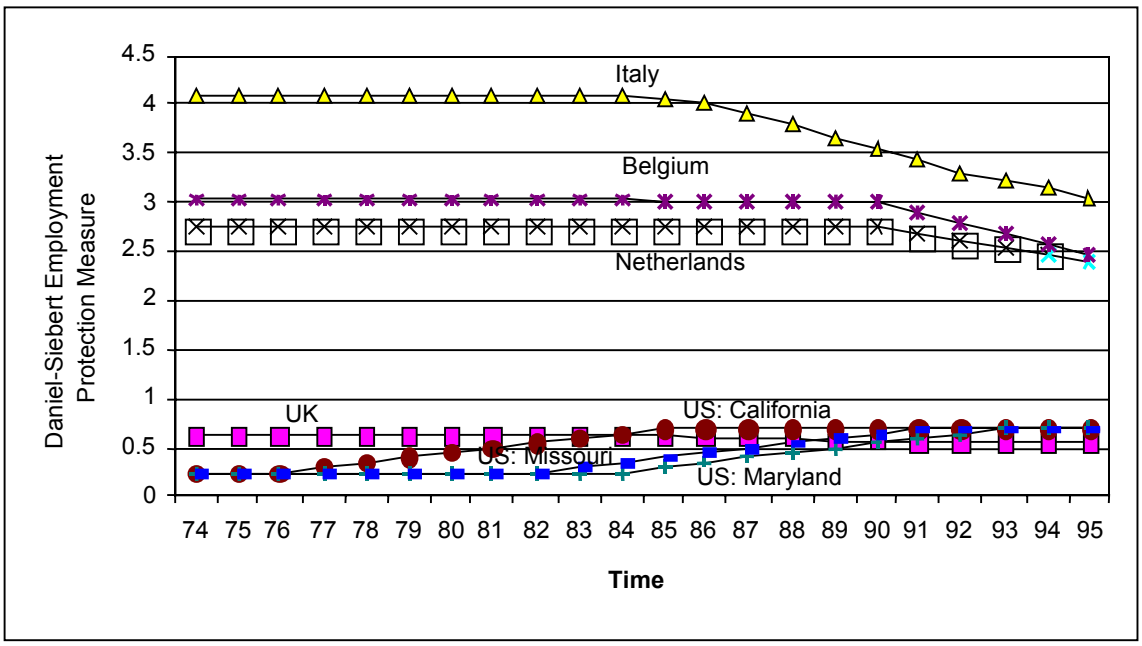


**Figure 5: Daniel-Siebert Employment Protection Index**

**Table 1: Labor Market Aggregates in Italy and the UK**

| | | Italy | | UK | |
|---|---|---|---|---|---|
| | | 1985-90 | 1995-00 | 1985-90 | 1995-00 |
| 20-24 age group: | Unemployment | 30 | 29 | 13 | 11 |
| | Employment/population | 47 | 40 | 72 | 69 |
| 25-54 age group: | Unemployment | 7 | 9 | 8 | 5 |
| | Employment/population | 64 | 63 | 76 | 79 |
| 59-64 age group: | Unemployment | 2 | 4 | 9 | 6 |
| | Employment/population | 22 | 18 | 34 | 35 |
| Adult and youth training + subsidized employment, participants as % of laborforce | | n.a. | 7.3 | n.a. | 2.1 |

**Source:** OECD, 2001, tables on Standard Labor Market Indicators, and Public Expenditures on Labor Market Programs.

**Notes:** Unemployment, and employment/population are percentages, averaged for the periods 1985-90, or 1995-2000. All figures are for male and females taken together.

## Table 2: Labor in the Study Plants, Mid 1990s

| | Ice-cream | | | Distillers | | | Food Processing | | | Pharma-ceuticals | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | IT | UK | US | IT | UK | US | NL | UK | US | BL | UK |
| Employ-ment, study plant[a] | 824 | 828 | 298 | 146 | 314 | 121 | 359 | 385 | 199 | 305 | 861 |
| % part-time | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 25 | 4 |
| % temp. | 18 | 24 | 24 | 15 | 14 | 4 | 2 | 12 | 12 | 9 | 3 |
| % union | 34 | 80 | 0 | 50 | 95 | 100 | 40 | 100 | 100 | 90 | 90 |
| Average tenure (yrs) | -- | 12.1 | 4.8 | 15.8 | 9.0 | 9.3 | 10.5 | 11.8 | 10.8 | 7.9 | 10.2 |
| Pay ($000s)[b] | 25.1 | 27.9 | 23.9 | 21.5 | 26.0 | 32.9 | 34.9 | 33.8 | 27.9 | 44.2 | 26.8 |
| Labor Cost[b] per prod. worker ($000s) | 35.0 | 32.1 | 36.4 | 44.1 | 29.5 | 47.3 | 53.1 | 39.0 | 40.9 | 67.9 | 29.7 |
| Unit Labor Cost ($)[b] | 0.220 per litre | 0.221 per litre | 0.224 per litre | 0.071 per litre | 0.049 per litre | 0.099 per litre | 74 per ton | 49 per ton | 58 per ton | 165 per $ 000 sales | 175 per $ 000 sales |

**Notes:** BL = Belgium, IT = Italy, NL = Netherlands, UK = United Kingdom, US = United States.
a Employment figures include production workers only.
b Figures are converted to US dollars using purchasing power parity.

## Table 3: Characteristics of Recruits, Mean Values, whole period

| | Ice-cream | | | Distillers | | | Food Processing | | | Pharma-ceuticals | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | IT[a] | UK | US[b] | IT | UK | US | NL | UK | US | BL | UK |
| Time period | 85-97 | 75-96 | 80-97 | 75-95[a] | 75-94[a] | 75-92a | 75-90 | 75-94 | 75-97 | 75-92 | 75-94 |
| Average hires per year: | | | | | | | | | | | |
| Total | 45 | 61 | 41 | 13 | 74 | 25 | 11 | 28 | 26 | 16 | 43 |
| Perman-ent males | 25 | 11 | 8 | 6 | 24 | 5 | 7 | 16 | 7 | 4 | 26 |
| Starting age of new hires (years): | | | | | | | | | | | |
| Average | 30.4 | 28.1 | 27.1 | 23.7 | 30.2 | 33.6 | 25.3 | 27.7 | 34.8 | 34.8 | 31.9 |
| Standard deviation | 5.2 | 10.0 | 7.9 | 5.1 | 10.0 | 9.1 | 5.9 | 9.1 | 9.8 | 7.1 | 11.2 |
| Education of new hires (years): | | | | | | | | | | | |
| Average | 10.6 | 11.9 | 12.1 | 9.8 | 11.4 | 12.1 | 10.8 | 11.2 | 12.4 | 10.9 | 11.0 |
| Standard deviation | 2.2 | 1.6 | 1.4 | 1.5 | 1.9 | 2.4 | 1.0 | 1.4 | 1.4 | 2.1 | 1.2 |

**Notes:** a Dates given are for the starting age series; 1975-1996 is the period for the education series for all Distillers plants.

## Table 4: Determinants of Demand for Education and Age Characteristics

| Variable (Mean) | **Average Starting Age** (30.1) | | **Standard Deviation of Starting Age** (9.2) | | **Average Education** (11.4) | | **Standard Deviation of Education** (1.6) | |
|---|---|---|---|---|---|---|---|---|
| Column | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Blanchard-Wolfers employment protection measure $_{t-1}$ (1.0) | -1.60 (-0.26) | | -9.12*** (-2.94) | | 2.35*** (3.13) | | -1.40 (-1.43) | |
| Daniel-Siebert employment protection measure $_{t-1}$ (1.3) | | -0.50 (-0.14) | | -3.73* (-1.75) | | 1.47** (2.08) | | -0.10 (-1.57) |
| Union density $_{t-1}$ (38.6) | 0.09 (0.46) | 0.07 (0.44) | -0.21** (-1.96) | -0.13 (-1.29) | 0.01 (0.28) | 0.01 (0.16) | 0.01 (0.28) | 0.01 (0.32) |
| Tax wedge $_{t-1}$ (33.1) | 0.18 (0.66) | 0.06 (0.25) | -0.11 (-0.85) | -0.06 (-0.48) | 0.11*** (2.80) | 0.10** (2.24) | 0.03 (0.65) | 0.03 (0.63) |
| Employment deviation (0.6) | 0.05 (0.64) | 0.03 (0.41) | -0.12** (-2.24) | -0.12** (-2.29) | 0.01 (0.61) | 0.01 (0.78) | -0.01 (-0.96) | -0.02 (-1.13) |
| Unemploy - ment $_{t-1}$ (8.0) | -0.44** (-1.93) | -0.35* (-1.92) | -0.07 (-0.61) | -0.13 (-1.13) | -0.08** (-1.99) | -0.07** (-1.96) | 0.00 (0.13) | 0.00 (0.05) |
| MFG pay (1.5) | -6.26 (-1.42) | -7.96 (-1.58) | -2.45 (-1.00) | -4.27 (-1.35) | 0.35 (0.44) | 1.57 (1.43) | -0.76 (-0.98) | -1.57 (-1.61) |
| Age of the worker stock (42.0) | 1.50*** (3.82) | -2.04* (-1.82) | | | | | | |
| School leaving age $_{t-4}$ (15.8) | | | | | 0.44*** (3.31) | 0.30* (1.79) | 0.04 (0.24) | 0.13 (0.83) |
| Education (11.4) | -3.25** (-2.05) | -1.39*** (-3.56) | | | | | | |
| Age (30.1) | | | | | -0.06 (-1.13) | -0.05*** (-2.88) | | |
| Time trend | 0.36 (1.50) | 0.23 (1.51) | -0.34*** (-3.31) | -0.22** (-2.36) | 0.08** (2.16) | 0.06* (1.77) | 0.02 (0.46) | 0.02 (0.77) |
| Plant fixed effects | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| $R^2$ | 0.48 | 0.56 | 0.54 | 0.52 | 0.59 | 0.59 | 0.36 | 0.37 |
| Durbin-Wu Hausman test for endogeneity | Ed. exog - Prob > F = 0.20 | Ed. exog. - Prob > F = 0.47 | N/A | | Age endog. - Prob > F = 0.01 | Age exog. - Prob > F = 0.50 | N/A | |
| Observations | 148 | 148 | 153 | 153 | 148 | 148 | 147 | 147 |

**Notes:** Estimates are by ordinary least squares, apart from column (1) and (5), for which age and education have been estimated simultaneously, using three stage least squares. The sample is males on open-ended contracts, including those who became subsequently employed on an open-ended basis within a year. t-values are given in parentheses, and *, ** and *** denote significance of the t-tests at the 10%, 5% and 1% levels.

## Table 5: Determinants of Labor Demand – Sensitivity Test Results

### A Sample Comparisons

| | | Sample | Average Starting Age | Standard Dev. of Starting Age | Average Education | Standard Dev. of Education |
|---|---|---|---|---|---|---|
| Blanchard-Wolfers employment protection measure | Employment protection coefficients [elasticities] | US + Europe | 0 | –9.12*** [–0.99] | 3.03*** [0.27] | 0 |
| | | Europe only | 0 | –9.56** [–1.84] | 2.84** [0.41] | 0 |
| | Time coefficients: | US + Europe | 0 | –0.34*** | 0.08** | 0 |
| | | Europe only | 0 | 0 | 0 | 0 |
| Daniel-Siebert employment protection measure | Employment protection coefficients [elasticities] | US + Europe | 0 | -3.73* [-0.53] | 1.59** [0.18] | 0 |
| | | Europe only | 0 | 0 | 2.08* [0.28] | 0 |
| | Time coefficients: | US + Europe | 0 | -0.22** | 0.06* | 0 |
| | | Europe only | 0 | 0 | 0 | 0 |

**Notes:** Elasticities calculated at the means are shown in square brackets. Plant fixed effects included in all equations. Other variables and estimation methods are the same as in Table 4.

### B Specification test: 5-year averages

| | | Average Starting Age | Standard Dev. of Starting Age | Average Education | Standard Dev. of Education |
|---|---|---|---|---|---|
| Blanchard-Wolfers employment protection measure | Employment protection coefficients | 0 | -1.26*** [–0.21] | 7.15** [–0.85] | 1.44** [1.20] |
| | Time coefficients: | 3.78** | 0 | 1.46*** | 1.08*** |
| Daniel-Siebert employment protection measure | Employment protection coefficients | 0 | -1.22*** [–0.20] | 5.26*** [0.64] | 0.95** [0.79] |
| | Time coefficients: | 0 | 0 | 1.10** | 0.82*** |

**Notes:** Elasticities calculated at the means are shown in square brackets Estimation is by unweighted OLS for all equations since variables are averaged over 5-year periods, which increases the number of observation underlying the dependent variables making weights unnecessary. To increase degrees of freedom, only significant fixed effects are retained.

## Table 6: The Link between Plant Tenure and the Employment Protection Indices

| Variable<br>(Mean) | Average tenure<br>(11.5) | |
|---|---|---|
| Blanchard-Wolfers<br>employment protection measure $_{t-1}$<br>(1.3) | 9.64*** <br>(7.13) | |
| Daniel-Siebert<br>employment protection measure $_{t-1}$<br>(1.4) | | 7.37***<br>(8.70) |
| Union density $_{t-1}$<br>(36.9) | 0.04<br>(0.84) | -0.13***<br>(-2.91) |
| Tax wedge $_{t-1}$<br>(33.0) | -0.07<br>(-1.27) | 0.08<br>(1.21) |
| Employment deviation $_{t-1}$<br>(-0.0) | -0.09***<br>(-2.63) | -0.07**<br>(-2.31) |
| Unemployment $_{t-1}$<br>(8.4) | 0.04<br>(0.66) | 0.09<br>(1.51) |
| Time trend | 0.42***<br>(7.68) | 0.23***<br>(5.29) |
| Plant fixed effects | Yes | Yes |
| $R^2$ | 0.85 | 0.86 |
| Observations | 164 | 164 |

**Notes:** t-values are given in parentheses. *, ** and *** denote significance of the t-tests at the 10%, 5% and 1% levels.

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## Appendix 1: Comparative statics results

Differentiating equations (6) and (7) totally we derive:
$(pg_{AA} - w_{AA} - \beta V\theta_{AA})dA + (pg_{AE} - w_{AE} - \beta V\theta_{AE})dE = \beta\theta_A dV$
$(pg_{EA} - w_{EA} - \beta V\theta_{EA})dA + (pg_{EE} - w_{EE} - \beta V\theta_{EE})dE = \beta\theta_E dV$

Solving for dA/dV using Cramer's rule gives:

$$\frac{dA}{dV} = \frac{\begin{vmatrix} \beta\theta_A & pg_{AE} - w_{AE} - \beta V\theta_{AE} \\ \beta\theta_E & pg_{EE} - w_{EE} - \beta V\theta_{EE} \end{vmatrix}}{\begin{vmatrix} pg_{AA} - w_{AA} - \beta V\theta_{AA} & pg_{AE} - w_{AE} - \beta V\theta_{AE} \\ pg_{EA} - w_{EA} - \beta V\theta_{EA} & pg_{EE} - w_{EE} - \beta V\theta_{EE} \end{vmatrix}}$$

Second order conditions require the determinant of the denominator, $\Delta$, to be positive for a maximum. On the education side, we require $w_E > 0$ and $w_{EE} > 0$ (wages increase at an increasing rate as shown in Figure 2). We have no priors about $\theta_{EE}$, and so assume $\theta_{EE} = 0$. We also assume $g_E > 0$ and $g_{EE} < 0$ (diminishing returns to education). On the starting age side we assume $g_A > 0$, $g_{AA} < 0$ (an inverted U for worker efficiency by age as shown in Figure 1), $\theta_A >$ or $< 0$, and $\theta_{AA} > 0$ (failure probability is U-shaped with age, though the reaction could be near-zero for workers in the prime age group). We have no priors regarding $w_A$ and $w_{AA}$, and so assume $w_A = w_{AA} = 0$. Finally, we assume all cross-products zero, except for $g_{AE} > 0$.

We then see that dA/dV is > or < 0 since:

$$\frac{dA}{dV} = \frac{\begin{vmatrix} \beta\theta_A & pg_{AE} \\ \beta\theta_E & pg_{EE} - w_{EE} \end{vmatrix}}{\begin{vmatrix} pg_{AA} - \beta V\theta_{AA} & pg_{AE} \\ pg_{EA} & pg_{EE} - w_{EE} \end{vmatrix}}$$

$= ((pg_{EE} - w_{EE})\,\beta\theta_A - \beta\theta_E\, pg_{AE})/\Delta$ which is > or < 0 since the first term is positive or negative, depending on $\theta_A$, and the second is negative.

However, dE/dV is likely to be > 0, since by a similar procedure we find:

$$\frac{dE}{dV} = \frac{\begin{vmatrix} pg_{AA} - \beta V\theta_{AA} & \beta\theta_A \\ pg_{EA} & \beta\theta_E \end{vmatrix}}{\begin{vmatrix} pg_{AA} - \beta V\theta_{AA} & pg_{AE} \\ pg_{EA} & pg_{EE} - w_{EE} \end{vmatrix}}$$

$= ((pg_{AA} - \beta V\theta_{AA})\,\beta\theta_E - \beta\theta_A pg_{AE})/\Delta$. Here the second term depends on $\theta_A$, and so again can be positive or negative. However, the first term is positive, and grows larger with V. Where V is sizeable, therefore, we can be confident that $dE/dV > 0$.

## Appendix 2: Reduced Form Equations for Average Age and Education

| Variable (Mean) | **Average Starting Age** (30.1) | | **Average Education** (11.4) | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| Blanchard-Wolfers employment protection measure $_{t-1}$ (1.0) | -8.87$^{*}$ (-1.67) | | 3.03$^{***}$ (2.74) | |
| Daniel-Siebert employment protection measure $_{t-1}$ (1.3) | | -2.73 (-0.076) | | 1.59$^{**}$ (2.18) |
| Union density $_{t-1}$ (38.6) | -0.09 (0.45) | -0.00 (-0.01) | 0.01 (0.16) | -0.00 (-0.03) |
| Tax wedge $_{t-1}$ (33.1) | -0.42$^{*}$ (-1.88) | -0.35 (-1.52) | 0.12$^{**}$ (2.48) | 0.11$^{**}$ (2.23) |
| Employment deviation (0.6) | -0.01 (-0.07) | -0.01 (-0.09) | 0.01 (0.46) | 0.01 (0.69) |
| Unemploy - ment $_{t-1}$ (8.0) | -0.07 (-0.343) | -0.14 (-0.72) | -0.07$^{*}$ (-1.86) | -0.06 (-1.52) |
| MFG pay (1.5) | -11.08$^{***}$ (-2.68) | -11.99$^{**}$ (-2.24) | 0.84 (0.96) | 2.05$^{*}$ (1.84) |
| School leaving age $_{t-4}$ (15.8) | | | 0.55$^{***}$ (2.97) | 0.36$^{**}$ (2.08) |
| Time trend | -0.06 (-0.34) | 0.05 (0.33) | 0.07$^{*}$ (1.95) | 0.05 (1.41) |
| Plant fixed effects | Yes | Yes | Yes | Yes |
| $R^2$ | 0.49 | 0.48 | 0.57 | 0.56 |
| Observations | 155 | 155 | 148 | 148 |

**Notes:** Estimation by weighted OLS.