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Explaining Variations in Wage Curves: Theory and Evidence

Erling Barth,[1] Bernt Bratsberg,[2] Robin A. Naylor,[3] and Oddbjørn Raaum[4]

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Abstract

This paper presents both theoretical analysis and econometric evidence for the United States, Great Britain and Norway on the extent to which hourly wages of different groups of workers are sensitive to local labour market conditions. We focus on differences by union status. Our theoretical framework captures both a turnover-based efficiency wage mechanism and one originating in union-firm bargaining. Under fairly general conditions, we show that wages are less sensitive to local unemployment the higher is the bargaining power of the union. In accordance with this theoretical prediction, we find that the absolute value of the elasticity of wages with respect to unemployment is higher in the nonunion sector than in the union sector for all three countries. We interpret the evidence as giving support to an efficiency wage interpretation of the wage curve.

Keywords: Wage curves, efficiency wages, bargaining

JEL-Code: J31, J41, J51

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1. Introduction

The effect of local unemployment on pay determination has become the focus of substantial analysis, particularly with the publication of 'The Wage Curve' by Blanchflower and Oswald (1994). It is not surprising that this should be so, as real wage flexibility plays a crucial role in determining how economies adjust to adverse shocks. The extent to which wages fall in response to increasing unemployment is a crucial determinant of the magnitude and persistence of the impact of labour demand or supply-side shocks (see, for example, Layard *et al*. (1991) or Blanchard and Katz (1997)). Furthermore, as Card (1995) has observed, a wage curve interpreted as an equilibrium relationship between wages and unemployment – that is, a quasi-supply function for labour - is a key potential element in the construction of an aggregate labour market model when combined with a simple labour demand schedule.

The central and surprising finding reported by Blanchflower and Oswald (1994) concerns the remarkable uniformity across countries in the estimated elasticities of earnings with respect to unemployment. The estimates imply that a rise in unemployment of ten percent will lead to a one percent fall in wages in almost all countries, regardless of differences across countries in factors such as institutional arrangements. This is particularly surprising in the context of cross-country studies suggesting that the levels of centralisation and coordination in wage bargaining *do* exert an influence on macroeconomic performance (see, for example, Calmfors and Driffill (1988) and Layard *et al*. (1991)) and the wage distribution, Blau and Kahn (1996). Moreover, consistent with this evidence on the importance of institutional arrangements is the finding reported by Blanchflower and Oswald (1994) that the estimated responsiveness of wages to unemployment seems to vary across union and nonunion workers. This issue of differences by union status in the elasticity of the wage curve is the focus of our paper.

There are three main motivations for our focus on union/nonunion differences in the elasticity of wages with respect to unemployment. Our first motivation is empirical and concerns the issue of how robust is the observation that the impact of local labour market conditions varies across union and nonunion workers. Our second motivation is policy-related. There is currently a significant debate on the role and relevance of trade unions in Europe (see, for example, Boeri *et al*. (2001)). In an important sense, the process of convergence toward European monetary unification has placed renewed emphasis on the role of unions as a mechanism for national wage coordination. The future role of unions at both the micro and the macro level is likely to be shaped significantly by policies of the European Union and it is important that the policy debate is informed by research evidence on the effects of unions on labour market outcomes.

Our third motivation is theoretical: we are interested in what the empirical evidence can tell us about the theoretical micro-foundations of the wage curve. Card (1995) argues that readers of Blanchflower and Oswald (1994), ". . . will find the efficiency wage model a leading contender for explaining the wage curve, yet . . . with no real evidence in favour of the model . . . " and suggests that more thought should be given to developing the implications of an efficiency wage model for differences in the slope of the wage curve across different groups of workers. In the current paper, we pursue this suggestion by developing an efficiency wage model of the wage curve in which we incorporate wage bargaining between firms and unions. Other models have integrated bargaining and efficiency wages (see, for example Hoel, 1989 and Rødseth, 1992), but these do not focus on how wages respond to variations in unemployment. The innovation in our model is that the wage curve we derive has two independent sources: it can be derived *either* from the efficiency wage mechanism, which we specify, *or* as a result of wage bargaining, or from the combination of the two. In other words, we combine in a single model both the efficiency wage (EW) and the wage bargaining (WB) micro-foundations of the wage curve, where Blanchflower and Oswald (1994) considered these as separate and unrelated models potentially underpinning the wage curve.

Our approach enables us to investigate how the predictions of the EW model vary across union and nonunion sectors, whilst allowing for an independent influence of unions on the wage curve elasticity. This last point is important empirically. As Card (1995) points out, the bargaining model does not offer a convincing explanation of the wage curve in countries, like the United States, in which unionisation is low. Furthermore, if wages are insensitive to local labour market conditions in the union sector, where bargaining determines the outcome, it is difficult to imagine that the average wage curve of the economy can result from the impact of local unemployment on bargaining outcomes. Moreover, if wage curves differ between the union and the nonunion sectors, we have a simple explanation for why wages tend to be less sensitive to *local* labour market conditions in countries with high union density and powerful unions. Of course, high union density tends to go hand in hand with more centralised wage setting with wages responding to changes in aggregate, rather than local, unemployment.

In countries with a strong tradition of wage bargaining there may be both EW and WB mechanisms operating to shape any resulting wage curve. Our theoretical framework nests the two explanations and allows us to make predictions regarding possible differences in wage curves across union and nonunion sectors even when efficiency wage characteristics are common to both sectors. We know of no reason to suppose that EW mechanisms characterise the nonunion sector only. As we show in section 3 of the paper, our framework suggests that wages are likely to be less sensitive to local unemployment the higher is the bargaining power of the union. This is the main prediction we address in our empirical analysis, where we investigate the slope of the wage curve - that is, the unemployment elasticity of wages - in 3 countries: the United States, Great Britain and Norway. We also develop a number of hypotheses and predictions in an attempt to discriminate between EW and WB models of the wage curve.

The rest of this paper is organised as follows. In the next section, we review the existing empirical evidence on wage curves in union and nonunion sectors. In section 3, we draw on the efficiency wage model of Campbell and Orszag (1998) and build a theoretical model in which bargaining power generates rent-sharing and local labour market conditions affect wage outcomes both via turnover costs and via conflict payoffs. The core empirical analyses are described and presented in sections 4 and 5. In section 6, we examine whether the relationship between wages and unemployment is better described by a Phillips curve than by a wage curve, and in section 7 we address the robustness of empirical results in subsamples of young and old workers. Section 8 concludes.

2. Empirical evidence on union and nonunion wages curves

During the last decade, numerous empirical studies from various countries have estimated the relationship between local labour market conditions, typically unemployment, and the wage level (see Blanchflower and Oswald, 1994). The empirical evidence seems to identify a negative relationship between wages and contemporaneous local unemployment, and the idea of an empirical law of an elasticity of earnings with respect to local unemployment of around –

0.1, suggested by Blanchflower and Oswald, seems to be widely accepted (Card, 1995). However, when the wage curve is examined separately for different groups of workers, the empirical pattern is one of large differences in wage curve elasticities across groups (Card, 1995; Baltagi and Blien, 1998; Turunen, 1998). Blanchflower and Oswald (1994) themselves report a distinctly different slope for the wage curve in the union and nonunion sectors.

Table 1 summarises estimates of the elasticity of earnings or wages with respect to the unemployment rate in the local labour market from four studies that report such estimates separately for unionised and nonunionised workers. These studies draw on micro data sets from the United Kingdom, the United States and the Nordic countries.

Strikingly, each of the studies listed in Table 1 finds greater wage curve elasticities for nonunionised workers than for unionised workers, although differences are not always statistically significant. In particular, strong indications that nonunion wages are more sensitive to regional unemployment rates appear in the evidence from the United Kingdom. For example, Blanchflower and Oswald (1994) find that an increase in local unemployment has little effect on hourly wages of unionised workers but a severe negative effect on wages of nonunionised workers. According to the estimates in the first row of Table 1, a doubling of the local unemployment rate will raise the hourly wage of unionised workers by (an insignificant) one percent and reduce the wage of non-unionised workers by 12 percent. While Blanchflower and Oswald note such differences, they place no stress on the differential responses of union and nonunion wages to local unemployment. Instead, they observe that the union-sector wage curve elasticity tends to be estimated imprecisely, particularly in data from the United States where samples of unionised workers often are limited in size (1994, p. 159). No U.S. results are reported for hourly wages with regional fixed effects. The huge differences between union and nonunion workers are found for annual earnings, and one might expect this to reflect annual hours being less sensitive to local labour market conditions in the union than in the nonunion sector. The difference by union status is minor in Turunen (1998), but this may reflect that the NLSY sample is for young workers for whom union affiliation is less important.

Albæk *et al*. (1999) report separate elasticities for union and nonunion sectors for each of the Nordic countries. In Finland, Norway, and Sweden, the absolute values of the coefficients are higher in the nonunion sector than in the union sector, but none of the differences is statistically significant. With respect to the Danish results, they comment that the union variable is actually membership in an unemployment insurance scheme rather than union membership, which may bias the results for Denmark considerably. The reported figures are from specifications without regional dummies. They find that the wage curve disappears altogether for these four countries once fixed regional effects are accounted for. The reported figures thus represent a medium- or long-term relationship between wages and unemployment in the Nordic countries, rather than transitory wage curve effects.

In sum, existing evidence suggests that wage responsiveness to local labour market conditions varies between union and nonunion workers, but the results for the United States and the Nordic countries are far from conclusive. First, we would like to provide more solid evidence, and, second, to draw inferences regarding the underlying theory of the wage curve from examining differences in union-nonunion wage curve elasticities across the United States, Great Britain and Norway.

3. Combining rent-sharing and efficiency wage explanations of the wage curve

In this section, we develop a theoretical model in which local labour market conditions potentially affect wages through either efficiency wage or wage bargaining channels, or both. Blanchflower and Oswald (1994) present three distinct models that all generate a wage curve response in the local labour market: a region-specific implicit contract model, an efficiency wage model, and a bargaining model. Card (1995) is dismissive of the implicit contract model as an explanation of the wage curve, pointing out that key predictions of the model are inconsistent with Blanchflower and Oswald's own empirical evidence.

The primary objective of our paper is to seek evidence on the relative strengths of the remaining two theoretical explanations of the wage curve. Accordingly, our theoretical framework nests a simple efficiency wage model with endogenous turnover and a standard bargaining, or rent-sharing, model as limiting cases. In this framework, the EW and WB mechanisms interact as local unemployment affects turnover costs and thus the magnitude of rents over which the union and the firm bargain. We use this framework primarily in order to study how the sensitivity of wages to the level of local unemployment - *the wage curve* - is influenced by the bargaining power of the union.

One might expect that combining efficiency wage and bargaining models of the wage curve would necessarily produce the prediction that the wage curve elasticity will be greater in the presence of unions. This is because where unions are absent there is only the efficiency wage effect but where unions are present there is both an efficiency wage and a bargaining effect of unemployment, both leading to lower wages at higher levels of unemployment. The novelty of our analysis, however, is that we find that combining the efficiency wage and bargaining models generates a third effect which we describe as the rent effect of local unemployment and which causes the wage curve to be *less* elastic in the presence of unions, thereby offsetting the separate efficiency wage and bargaining effects. The intuition for the rent effect is that higher unemployment leading to lower turnover costs generates greater rents in the firm which the union will share in the form of a moderated reduction in the wage level. This is the reason why the wage curve is likely to be less elastic in the presence of unions. We now develop a formal theoretical framework in which to explore these issues.

The efficiency wage mechanism is due to endogenous turnover costs, given by

(1) $$C = H\left(\frac{W}{(1-u)\overline{W}}\right)^{-\eta}, \eta > 0$$

where u is the local unemployment rate and $\overline{W}$ is the average local wage rate. C can be interpreted as replacement costs per worker defined by the product of hiring costs per worker, H, and the probability that a worker quits (see Campbell and Orzag (1998)). In equation (1), the separation probability is assumed to depend on the current wage relative to the alternative wage, $\overline{W}$, weighted by $1-u$, which measures the probability of becoming re-employed. The parameter $-\eta$ is the elasticity of quits with respect to expected relative pay.

The firm's profit is given by:

(2) $$\Pi = R(L) - CL - WL$$

where R(L) is total revenue. For simplicity, the local union is assumed to care only about the bargained wage level.

We assume that during a potential conflict over an agreement (e.g., a strike), the firm receives a zero payoff, while the union receives $\varphi(1-u)\overline{W}$, where $\varphi \in [0,1]$ measures the fraction of the expected alternative wage that workers may obtain as income during a strike. If $\varphi = 0$, workers recieve no income during a strike. In this case unemployment does not affect workers' conflict pay and there is no bargaining microfoundation to the wage curve. As in

Blanchflower and Oswald, the wage curve associated with bargaining arises because higher unemployment lowers workers' threat points. From (2), the Nash-maximand is given by:

(3) $$N = (W - \varphi(1-u)\overline{W})^{\beta}\, \Pi^{1-\beta}$$

where β is the bargaining power of the union.

The first order condition for the maximum of (log) N is:

(4) $$\frac{\beta}{W - \varphi(1-u)\overline{W}} = \frac{(1-\beta)L(C'_{w}+1)}{R - CL - WL}$$

All firms are assumed to be identical and in equilibrium we have $W = \overline{W}$. Inserting this equilibrium condition into (4) gives:

(5) $$W = b(y - C) + (1-b)\eta C$$

where

(6) $$0 \le b = \frac{\beta}{1-(1-\beta)\varphi(1-u)} \le 1$$

and y=R/L is revenue per worker. Note that if $\varphi = 0$, then b = β.

Consider first the limiting cases, which describe the efficiency wage and pure bargaining models, respectively, in equation (5).

(i) *The pure efficiency wage (EW) model of the wage curve*

The pure efficiency wage model occurs when $\beta = 0$. From (5), we see that the equilibrium wage level in the pure EW model is given by:

(7) $$W^* = \eta C,$$

where the superscript denotes the pure efficiency wage benchmark outcome. The efficiency wage is proportional to the turnover-cost per worker. *The benchmark efficiency wage elasticity of wages with respect to unemployment* is then

(8) $$\lambda^* = \frac{\partial W^*}{\partial u}\frac{u}{W^*} = -\eta \frac{u}{1-u} < 0$$

as in Campbell and Orzag (1998). The EW unemployment elasticity, λ^*, is proportional to the elasticity of turnover with respect to local unemployment and is decreasing in η.

Notice that since $W^* = \eta C$, the bargained wage given by (5) can be interpreted as a weighted average of the rents per worker, net of turnover-costs, and the efficiency wage. We know from the above that the weight, b, is equal to β when the alternative wage is independent of local unemployment, i.e., when $\varphi = 0$.

(ii) *The pure wage bargaining (WB) model of the wage curve*

The pure bargaining model arises under the special case that $\eta=0$. In this case, it follows from (1) and (5) that the equilibrium bargained wage (W^B) with exogenous turnover (and associated costs) is given by:

(9) $W^B = b(y-H)$.

b is defined in (6) and is a function of u. From (9) and (6), local unemployment affects the bargained wage through its influence on the expected alternative wage and thereby on the conflict pay-off of the union. Thus, under pure bargaining, the "benchmark" elasticity of the derived wage curve is given by the *elasticity of the alternative wage with respect to unemployment*. From (9) and (6), this is given by:

(10) $$\lambda^B = -\frac{(1-\beta)\varphi u}{1-(1-\beta)\varphi(1-u)} \leq 0$$

which is our formalization of the bargaining model mechanism discussed by Blanchflower and Oswald (1994). From (10), it follows that the absolute value of the elasticity of the pure bargained wage with respect to unemployment is decreasing in φ; that is, the smaller the proportion of expected alternative wages obtained during a strike, the smaller is the impact of unemployment on wages.

(iii) *Efficiency wage and wage bargaining models of the wage curve combined*

In the pure bargaining model of the wage curve, we assumed that turnover was exogenous to the level of unemployment. Consider now the case with endogenous turnover. In this case both bargaining and efficiency wage mechanisms influence wages. First, from (1), we define the *elasticity of net revenue with respect to unemployment:*

(11) $$\lambda^R = \frac{\partial(y-C)}{\partial u}.\frac{u}{y-C} = \eta\frac{u}{1-u}\frac{C}{y-C} > 0$$

This represents the rent effect of local unemployment on bargained wages. This rent effect is positive since higher unemployment decreases turnover costs. Consequently, it is increasing in the absolute value of the turnover elasticity, η.

Using the definitions of $\lambda*$, λ^B, and λ^R given in (8), (10) and (11) respectively, the combined wage curve elasticity derived from (5) and (6) can be written as:

(12) $$\lambda = p\lambda^* + (1-p)\lambda^B + \frac{b(1+\eta)(y-C)}{W}\lambda^R$$

where $p=(W^*/W)\leq 1$. Recall that W^* denotes the pure efficiency wage and that W denotes the equilibrium bargained wage in the combined model. The formulation for the combined wage curve elasticity given in (12) enables us to view the total wage curve elasticity as combining three elements: the pure efficiency wage effect, the alternative wage effect and the rent effect.

The rent effect is a novel feature of the combined model, as it does not appear in the limiting cases of the existing bargaining (η=0) or efficiency wage models (β=0). The intuition behind the rent effect is that as unemployment increases, turnover costs fall and the rent to be shared rises. Since the bargaining model is a rent-sharing mechanism, increased rent feeds into higher wages. The two first terms of λ are both negative, while the last term is positive. One implication of this is that if λ^B (i.e., the alternative wage effect) is sufficiently small relative to the rent effect, then the wage curve elasticity under union bargaining will be smaller (in absolute terms) than in the nonunion case.

It follows from equation (12) that:

(13) $$\lambda = \lambda^* < 0 \quad for \quad \beta = 0,$$

(14) $$\lambda = \lambda^R > 0 \quad for \quad \beta = 1.$$

Thus, with no rent sharing ($\beta = 0$), we are back in the pure efficiency wage case. In the monopoly union case ($\beta = 1$), the wage curve elasticity equals the pure rent-effect of unemployment, which is positive. When the union is free to set whatever wage it prefers, it collects all the rents: that is, W = y - C. From (13) and (14), it is clear that comparing a nonunion case with a strong-union case, the elasticity of wages with respect to unemployment is not only higher under the latter, but it is actually positive.

The strong assumption of a monopoly union is convenient to derive simple theoretical predictions, but hardly plausible. The more interesting question concerns how the unemployment elasticity changes with the bargaining power of the union. We can show that as union bargaining power increases, (i) the alternative wage effect becomes weaker, and (ii) the rent effect becomes stronger, pulling λ upwards and flattening out the wage curve as the union becomes more powerful. It is straightforward to see that the rent effect must be increasing in union bargaining power. To see how the alternative wage effect responds to a change in bargaining power consider equation (10). From (10), it follows that the alternative wage effect, λ^B, is *increasing* (that is, decreasing in absolute value) in the bargaining power of the union as:

$$(15) \qquad \frac{\delta\lambda^B}{\delta\beta} = \frac{\varphi u}{(1-(1-\beta)\varphi(1-u))^2} \geq 0$$

The intuition behind this result is straightforward. As the bargaining power of the union increases, more weight is given to the inside factor, *y-H* in (5), and less weight to the outside expected alternative wage component.

Consider equation (12) once more. From (12), the derivative of λ with respect to β is given by:

$$(16) \qquad \frac{\partial\lambda}{\partial\beta} = -(\lambda^B - \lambda^*)\underset{-}{\frac{\partial p}{\partial\beta}} + (1-p)\underset{+}{\frac{\partial\lambda^B}{\partial\beta}} + (\eta+1)(y-C)\underset{+}{\lambda^R}\,\underset{+}{\frac{\partial \frac{b}{W}}{\partial\beta}}$$

For $\lambda^* < \lambda^B$, λ is unambiguously increasing with the bargaining power of the union, causing the wage curve to flatten as the union becomes more powerful. The first term represents an increased weight on the alternative wage effect compared to the efficiency wage effect of unemployment, when β rises. The second term reflects the influence of increased bargaining power on the alternative wage effect. The third term is the increased impact of the rent-sharing effect as the rent sharing parameter increases.

From (16), it follows that introducing bargaining in a pure efficiency wage situation, may *initially* lead to more sensitivity towards unemployment if the efficiency wage effect is small and the share of alternative income in workers conflict payoff is very large, such that $\lambda^* > \lambda^B$. In this case λ may become more negative as β increases, since we have one negative influence counteracting the two positive effects. The condition for the efficiency wage effect to be smaller or equal to the alternative wage effect may be written as:

$$(17) \qquad \beta \leq \tilde{\beta} = 1 - \frac{\eta}{(1+\eta)\varphi(1-u)}$$

For β smaller than $\tilde{\beta}$, the first term of (16) is negative. We note a number of properties of the critical level. First, $\tilde{\beta}$ decreases with η; the larger the efficiency wage effect (in absolute terms), the smaller is $\widetilde{\beta}$. Second, $\widetilde{\beta}$ decreases with the share of alternative wage accruing to workers during a conflict, φ. For

(18) $$\varphi \leq \frac{\eta}{(1+\eta)(1-u)}$$

any positive β is greater than $\widetilde{\beta}$ and λ is always increasing in β. If condition (18) is not met, there is a region, bounded from above by $\widetilde{\beta}$, in which λ *may* initially decrease with β.

The interpretation of this result is as follows. Initially, for small values of β, the alternative wage effect may be larger than the efficiency wage effect. Since the rent-sharing parameter is small for small values of β, the positive influence of the rent sharing effect of unemployment is limited. Increasing β means putting more weight on the alternative wage effect, which then tends to pull the wage curve elasticity (-λ) upwards. However, we found that the alternative wage effect approaches zero as β increases and the union gains more protection from external market pressures. The sign of λ^*-λ^B changes from negative to positive at $\beta = \widetilde{\beta}$. From then on all three effects contribute to a wage which is less sensitive to local unemployment as a consequence of a more powerful union.

Our model predicts that introducing bargaining in a pure efficiency wage situation, may *initially* lead to more sensitivity towards unemployment if the efficiency wage effect is small and the share of alternative income in workers conflict payoff is very large. However, two other forces work in the other direction. As bargaining power increases, the alternative wage effect diminishes and the rent-sharing effect goes up. We find the region of falling λ to have little empirical relevance. First, there are considerable wage curve effects in nonunion sectors, see for example, previous evidence reported above. Second, it seems unlikely that workers obtain a very large share of alternative income during a strike. Third, there are counteracting forces from the last two terms of (16) which means that the potential minimum of λ is well below $\widetilde{\beta}$. Finally, a very low level of β is not of much empirical interest when comparing union with nonunion sectors. It is likely that, even in a nonunionised sector there may exist rent-sharing mechanisms and union recognition with bargaining introduces non-trivial levels of union influence on wages. Consequently, we view the result that bargaining might increase the sensitivity of wages with respect to unemployment as unlikely to have significant empirical relevance.

As we indicated at the beginning of this section, one might have expected that combining efficiency wage and wage bargaining models would have produced a wage curve elasticity which was a weighted sum of the elasticities in the two models. The main result from our theoretical analysis is, however, that the encompassing model also produces another effect of unemployment; namely the rent-sharing effect. This effect is always positive since higher unemployment produces lower turnover costs and thus more rents to be shared. And, most important, as bargaining power increases*, the rent-sharing effect increasingly dominates* the other two effects. Eventually, as the bargaining power parameter approaches unity, the positive rent-sharing effect is the *only* mechanism through which unemployment affects wages.

Let us summarise the key predictions arising out of the theoretical framework. First, if wage curves are generated solely through collective wage bargaining, then we would not expect to find wage curves in the nonunion sector. Second, if efficiency wage mechanisms underpin the wage curve, then we would expect to find wage curves in the nonunion sector. Third, assuming that efficiency wages are just as likely to characterise union as nonunion workplaces, then we would not be surprised to find an efficiency-wage generated wage curve in the union sector. However, we would expect any efficiency-wage based wage curve to be smaller in absolute magnitude in the union than in the nonunion sector, partly as a result of the rent-effect which we have identified in the current paper. Fourth, we would expect wage curves to be less elastic in countries with stronger unions and with more pervasive union wage bargaining. These are the predictions to which we address our empirical analysis in the subsequent sections of the paper.

3.1 Age, experience and turnover

In this paper, we focus on differences in the wage curve by union status. As noted above, Card (1995) observed how in principle patterns of variation in the wage curve elasticity across groups could be used to choose between competing theories of the wage curve. The pure efficiency wage model we have presented predicts that the absolute magnitude of the wage curve elasticity will be increasing in η, the elasticity of the quit rate with respect to expected relative pay. The quit rate is higher for younger and less experienced workers and also more sensitive to relative wages. Therefore, η is inversely related to the age (experience) of the worker. Thus, the pure EW model predicts a more elastic wage curve for younger, less experienced workers. The impact of η on the unemployment elasticity in the union sector is smaller.

Consider this in the context of the combined model. Equation (13) can be re-written as:

(19) $$\lambda = (1-p)\lambda^B + p\frac{u}{1-u}\left[b - \eta(1-b)\right].$$

In this formulation, the wage curve elasticity, λ, is defined as a weighted average of λ^B, the elasticity of the alternative wage with respect to unemployment, and of the term in square brackets, where the weights are given by *1-p* and *p*, respectively. Recall that *p=W*/W*, and hence varies with β. Equation (19) lends itself to a straightforward diagrammatic representation: see Figure 1.

In the nonunion case (*b*=0), $\lambda^{Nonunion} = \lambda^* = -\eta u/(1-u)$, as given by equation (8). In the union case, λ^{Union} is given by (19), or by the weighted average of λ^B and f in Figure 1. A number of properties emerge from consideration of equation (19) and its representation in the Figure. First, as argued above for the case of the pure EW model, $\lambda^{Nonunion}$ falls as η increases. Second, the effect on λ^{Union} of a change in η is ambiguous: λ^{Union} first rises and then falls in η. Third, at least for large η, the absolute magnitude of the difference between $\lambda^{Nonunion}$ and λ^{Union} is increasing in η. We address these testable propositions within the unionised EW model in the empirical sections below.

3.2 Robustness

The predictions of wage setting models are frequently sensitive to the more or less explicit assumptions underlying the theoretical set-up. Do our results rely on our specific bargaining model and efficiency wage mechanism, or can we consider them to be more general? Consider the more general, although ad hoc, wage setting mechanism, where

(20) $$W = w^e + r\left[y(E) - w^e\right]$$

where w^e denotes the equilibrium efficiency wage, E is an efficiency indicator and r is a rent-sharing parameter, $0 < r < 1$. Any efficiency wage model where productivity (value added) is enhanced by unemployment, for instance a Shapiro-Stiglitz (1984) type of model, produces the efficiency wage effect of unemployment, comparable to λ^* in our setup. The impact on wages is found by taking the derivative of W with respect to w^e in (20). More importantly, for any positive r, *any* such model involves *the rent effect* as well, the impact of which is to be found by taking the derivative of W with respect to E above. Again, we note that the combined impact from the efficiency wage mechanisms implies a positive relationship between the rent sharing parameter r and the elasticity of wages with respect to unemployment. The larger fraction of the rent accruing to workers, the larger is the impact of the positive rent-effect of unemployment on wages.

We use the above set up with turnover costs for two reasons. The first is that a simple expression for λ is already available in the literature (Campbell and Orzag, 1998) and that the model has a strong standing in the theoretical literature, see for example Weiss (1991). It is also closely related to the equilibrium wage distribution literature, Burdett and Mortensen (1998) and Green, Machin and Manning (1996), even though the quit-function here is a major simplification. The second is that we tend to believe that endogenous turnover is an important feature of most firms and establishments. Barth and Dale-Olsen (1999) find in an empirical study that establishment turnover, over and above that required to adjust the stock of employees, is sensitive to the wage policy of the establishment. In any case, we argue that our main result does not rely on this specific efficiency wage mechanism, but only on the assumption that productivity is affected by the level of local unemployment.

Of course, the rent-sharing coefficient, *r*, may itself be affected by unemployment, as in our model above, in which the feedback effect through alternative wages affects the outcome. Other mechanisms are also plausible. For example, unemployment may affect the ability of one of the parties to put up a credible ultimatum threat, thus affecting the bargaining power parameter. The main point to make here, however, is that our main result survives even with an exogenous *r* and does not rely on the specific modelling of the alternative income effect.

Our model assumes decentralized wage setting, with wage increments in one firm affecting other workers and firms through turnover and its associated costs. Under centralised bargaining, firms may incur given turnover costs for lower wages. Workers too may benefit from coordinated wage moderation either in terms of higher after-tax wages, if we take into account that unemployment benefits are financed by income taxes - see, for example, Holden and Raaum (1991) - or if coordination gives workers other incentives to internalise the employment effects of higher wages.[1] Once coordinated wage setting takes place, the empirical wage curve may look different, even if the same basic mechanisms work on turnover and alternative income.

From our theoretical analysis, we conclude that the wage curve effect is likely to be smaller in an unionised environment than in a nonunion environment. Thus, within countries we expect to find a lower wage elasticity with respect to unemployment in the union sector than in the nonunion sector. Moreover, because of the higher levels of unionisation and of centralisation and coordination of bargaining in Norway, we would also expect to observe lower wage curve elasticities in Norway than in the United States and Great Britain.

3.3 *A two-sector model*

The model developed in the preceding section analyses the case of homogeneous firms where no mobility exists between union and nonunion jobs. In this section, we argue that our main result prevails also in a two-sector setting. With mobility between sectors, the workers' outside option consists of either unemployment or of employment in one of the two sectors. Consider now the effect of increased unemployment. The process leading to our equilibrium wage (equation (5) above) may be thought of as occurring in two stages. First, firms or establishments respond to a higher unemployment rate, given wages elsewhere, with a reduction in their own wage level: there is a wage cut in each firm. In the second stage, there is a feedback effect from reduced average wages in all other firms, which in turn leads to an even lower wage in each firm, and so on. In a two-sector setting, the feedback effect is influenced by wage reactions in both sectors. The expected outside wage, ω^j, can be modelled as:

$$\omega^j = (1-u)\left[\alpha^j \overline{W^u} + (1-\alpha^j)\overline{W^n}\right] \tag{21}$$

where the α^j (*j=u,n*) is the probability of getting a job in the union sector.

In this case, there are feedback effects across the two sectors. First, there is a direct effect from the *(1-u)* term. Second, there is the indirect effect from the reaction of other firms to unemployment and so on. The feedback effect works through the average wages in equation (21). For the nonunion sector, the fact that the union sector wages enters the expression for the alternative wage introduces *less* sensitivity of wages to unemployment than if only nonunion

[1] There is an extensive literature on why employers and unions may have incentives to coordinate wage setting,

sector wages were included in the calculation of the outside option. The reason is that the relatively unemployment-inelastic union wage enters the expression for the outside wage. And similarly, the feedback effect for the union sector is more wage-elastic since the nonunion wage response enters the calculation of outside wages. The feedback effect across sectors thus tends to moderate the original difference between the two sectors. This mechanism, however, is a second-order effect, working *only if there is a difference in elasticities* in the first place. Although feedback effects between sectors reduce differences, they cannot eliminate the difference in wage elasticities between sectors. Hence, our previous results hold also for a two-sector model.

4. U.S., British and Norwegian data

In the light of the theoretical analysis presented above, we estimate wage curves for three different economies. Because the theoretical model explicitly models the wage-setting process of profit maximising unionised and nonunionised firms, the empirical analyses are of private sector workers and the dependent variable is the log of the hourly wage of the worker.

The U.S. estimates are based on two data sources: the Outgoing Rotation Group samples from the Current Population Survey (CPS) and the National Longitudinal Survey of Youth (NLSY). The CPS is a monthly survey of about 50,000 households, and each month one-quarter of the respondents (the outgoing rotation groups) provide detailed employment information including current wages. Although the survey has been conducted for more than 50 years, the question about union membership has been asked on a monthly basis only since January 1983. Accordingly, our CPS samples are drawn from the January 1983 through December 2000 surveys. The NLSY is an ongoing longitudinal survey of a nationally representative sample of 12,686 youths aged 14-22 in 1979. Our samples are drawn from the first 18 waves (1979 through 1998) of the survey.

The classification of workers into union and nonunion sectors differs somewhat across the two U.S. data sources. In the CPS samples, we classify union workers as those who hold membership in a trade union or an employee association "similar to a union." Unfortunately, the NLSY has not asked respondents about union membership on a consistent basis, so we rely

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see, e.g., Hoel (1989), Moene and Wallerstein (1993) and Layard *et al.* (1991) for careful analyses.

instead on information about whether wages on the particular job were set by collective bargaining. Besides the very different age distributions of respondents, differences in union classification may be expected to generate differences in estimates of unemployment elasticities in the two data sets.

We limit the U.S. regression samples to non-agricultural, private-sector workers 16 to 64 years of age, and exclude those who are self-employed or usually work one hour or less per week. We also drop observations with hourly wages below $1 per hour (constant 1982-84 dollars) or with missing data on variables used in the analysis. The hourly wage is defined as the rate of pay for hourly employees and as weekly earnings divided by usual hours worked per week for salaried workers. The sample restrictions leave total samples of 2,353,202 observations (of whom 52.7 percent are male and 12.1 percent are members of a trade union) from the CPS and 89,210 observations (53.1 percent male and 16.0 percent union) from the NLSY. We merge the micro samples with monthly data on unemployment in the state of residence, defining the unemployment rate most relevant to the prevailing labour contract as the average local unemployment rate over the 12 months prior to the wage observation. The unemployment rates are collected from the Local Area Unemployment Statistics (LAUS) program of the Bureau of Labor Statistics.

The British data are taken from the first seven waves (1991-1997) of the British Household Panel Survey (BHPS), a nationally representative survey. The full survey contained information on just over 10,000 individuals. The sample used for the estimations reported in the current paper is based on employees aged between 16 and 64 employed in the private sector. Combining information across the waves, the total sample size is 21,109 workers. Workers are classified as employed in the union sector if a union is reported as present at their establishment. Where a union is reported as not present, the worker is classified as employed in the nonunion sector. Thus, the union/nonunion distinction is defined on the basis of the characteristics of the establishment, not of the individual worker. This is consistent with the theoretical model in which wage-setting is based not on individual bargaining but on union-firm bargaining. We would expect the individual's wage to reflect the bargaining status of the union at the establishment, not the individual's personal union membership status. For a further discussion of this issue, see Barth, Raaum and Naylor (2000), and Andrews, Stewart, Swaffield and Upward (1999). Separate regressions are run for union and nonunion sectors and for male and female workers. The dependent variable is the log of hourly pay. Unemployment refers to the percentage of the labour force in the relevant Travel-to-Work-Area (TTWA) who are unemployed at the time of the survey interview.

The Norwegian data is register based. It is drawn from the register of employer and employees at Statistics Norway. The sample includes 111 463 observations of wage earners from 1991 to 1997. Representative samples are drawn independently each year. The analysis is conducted for full-time workers. The wage is based on daily wage for full-time workers, corrected for predicted hours (from the level of living survey 1995) for full-time workers based on gender, education, region, industry and age. The information on gender, schooling, (Mincer) experience, region, industry and seniority are register-based and highly reliable. The sample includes persons from 16 to 64 years of age. Extreme wage observations are removed. Information on union membership is taken from individual wage reports and thus includes only union members who pay their dues through the employer. This is common practice for the large majority of union members in Norway. The Norwegian unemployment figures are collected at county level.

In sum, we find the data from the three countries comparable, but there are also several differences. The outcome variable is the hourly wage, the set of controls includes broadly comparable variables, age bands cover the same groups and the time periods are similar. The union worker definition differs somewhat, as the CPS and the Norwegian data use individual membership, while the split is based on union existence at the establishment in the NLSY source and the BHPS. The aggregation level differs as we study states in the United States, counties in Norway and TTWA's in Great Britain. The focus, however, is on differences by union status *within* countries. All in all, the comparability is sufficiently high to make interesting cross-country observations.

5. Empirical wage curves

Wage curve estimates by gender and by union status are reported in Table 2. The unemployment elasticities are based on the static model of Blanchflower and Oswald[2];

(22) $w_{irt} = \alpha + \lambda \log u_{rt} + \gamma X_{irt} + d_r + f_t + v_{irt}$,

where w_{irt} is the (log) wage rate of individual i in region r observed in period t, X_{irt} is a set of individual controls (such as age, education, and industry affiliation), d_r and f_t are unrestricted fixed regional effects and period effects, respectively, and v_{irt} denotes the error term. Basically, the unemployment elasticity, fixed regional effects, period effects and coefficients of control variables are estimated using a series of stacked cross-sections.[3] Applying region-by-period clustering, any period-specific correlations in error terms within regions are taken into account when estimating the standard error of the unemployment elasticity, see Moulton (1990), Card (1995).

In all countries, the estimates of the male unemployment elasticity are negative and statistically significant in the samples of nonunion workers. The wages of male union workers are, however, less affected by the local unemployment rate. In the CPS data for the United States and in Norway, the point estimates are close to zero for union males, while the negative NLSY estimate is not statistically different from zero.[4] The British estimate for union males is negative, but not statistically different from zero. As the precision of the estimated unemployment elasticities in Great Britain and Norway are low, we cannot reject the null hypothesis that union and nonunion estimates are equal. The point estimates, however, are indeed lower (in absolute value) in the union sector. In the CPS data, the difference in unemployment elasticities between male union and nonunion workers is clearly statistically significant.

For females, the nonunion elasticities are all negative, but much lower in absolute value compared to males (in the same country). In Great Britain, wages of females in the nonunion sector are not affected by local labour market conditions. Moreover, only the U.S. estimates are significantly different from zero. Among female union workers, wages are basically unrelated

[2] A dynamic version based on the panels of regions is discussed in the next section.
[3] In other words, the panel element—that the same individuals are observed in more than one year—of the NLSY, British and Norwegian samples is ignored.
[4] The difference between estimates from the CPS and NLSY samples may result from differences in age composition of the two samples—a topic we return to below.

to regional unemployment except for the CPS data where the unemployment elasticity is highly positive.

The various estimates in Table 2 form independent observations of union/nonunion differences in wage responsiveness to regional unemployment. We therefore consider a test based on differences between the average coefficients from the nonunion and union sectors across all samples. The average male nonunion elasticity across the four samples is –0.0738 with a standard error of 0.0121. The average male union elasticity is –0.0292 with a standard error of 0.0162. The difference between the average coefficients for the nonunion and union sectors, –0.0446, is significantly different from zero with a t-value of –2.21. Similarly, for the female sample, the difference has a t-value of –2.13. Considering the averages from both male and female samples, the difference of –0.0441 has a t-value of –2.91. The meta analysis thus rejects the hypothesis of equal wage responsiveness between nonunion and union sectors.

In sum, the empirical evidence supports the key prediction of the efficiency wage model: nonunion wages are adversely affected by worsened conditions in the local labour market and more so than union wages. In fact, there is no statistically significant downward sloping wage curve for union workers in any of the three countries.

The U.S. estimates reported in Table 2, based on the CPS samples, are substantially smaller than those of Blanchflower and Oswald (1994; 1995). According to Table 2, a doubling of the local unemployment rate will reduce wages of nonunionised males by 4 percent and females by 1.6 percent. Some of the discrepancy with prior studies may be because ours are estimates of the responsiveness of hourly wages while Blanchflower and Oswald typically study annual earnings. Card (1995) and Bratsberg and Turunen (1996) find smaller wage curve elasticities when the dependent variable is the hourly wage. For male nonunion workers, the estimate in Table 2 is in line with those of the Card and Bratsberg/Turunen studies.

In the U.S. data, the female union sector wage curve is found to be upward sloping. Our theoretical model offers two explanations. First, workers in female dominated *unionised* firms have higher bargaining power. There is no evidence to back up such an assumption, quite on the contrary, there is evidence that wages are lower in firms with more females – see, for example, Hellerstein et al (1997). Second, and perhaps more plausible, when quits respond weakly to changes in the expected relative wage, the rent-sharing effect (shown to be positive in equation 16) dominates the efficiency wage effect at low levels of bargaining power. Barth and Dale-Olsen (1999a,b) provide evidence that the elasticity of quits is lower for women than for men, and Sicherman (1996) finds that, at low levels of tenure, women are less likely than men to leave the firm.

A related empirical literature addressing union and nonunion wage structures in the United States has noted that union wages show less cyclical variation than nonunion wages (Lewis, 1963; 1986; Moore and Raisian, 1980; Hendricks, 1981). The lack of cyclicality in union wages is often attributed to the generally longer duration of union contracts. When we alternatively use the lagged 3-year average state unemployment rate in the wage regression, we find that wage curve elasticities become slightly more negative in both sectors (results not shown). Equally important, differences between unemployment effects in union and nonunion sectors are even larger than those in Table 2. We therefore conclude that the empirical pattern that wage curve responses are larger in the nonunion sector than in the union sector is not driven by differences in contract length in the two sectors.

Compared to previous British studies, we find that the estimated male unemployment elasticities are in line with those reported in Blanchflower and Oswald (1994), around -0.1 when the sample consists of both nonunion and union workers. Our nonunion estimate for males is about -.11 compared to Oswald and Blanchflower's estimate of –0.17 for nonunion plants in the Great Britain, obtained from the British Social Attitudes Survey. Again, part of this difference may stem from the fact that we use hourly rather than annual or weekly earnings information. Additionally, it may be that the wage curve elasticity has diminished in the Great Britain since the time at which the Blanchflower and Oswald estimates - based on data for the mid to late 1980s - were obtained.

The Norwegian regressions reveal a significant wage curve for all males, but splitting by union membership shows that local labour market conditions affect only male nonunion wages. We find no significant wage curve for the remaining Norwegian samples. Such findings are consistent with Albæk *et al*. (2000), who conclude that, when fixed regional effects are controlled for, there are no wage curve effects in the Nordic countries. Albæk *et al*. interpret their finding as resulting from the rather high degree of centralization of bargaining in these countries, and argue that real wage flexibility is obtained at the national level rather than at the regional level. Because the standard wage curve regressions include time dummies, aggregate wage flexibility is effectively swept out of the analysis. Moreover, in Norway the yearly central wage increments in the union sector effectively form a floor of wage growth in the nonunion sector as well, which may reduce regional wage responsiveness even in the nonunion sector. The results for Norway in Albæk *et al.* were based on data from 1989 to 1993, a period characterized by a high degree of centralization and rather small nominal wage increments. In contrast, the data underlying the present paper also include observations from 1993-97, a period during which local wage settlements formed a considerably larger part of total wage growth. Our conjecture is that this is the main reason why we now observe a significant wage curve for males in the nonunion sector in Norway. Our results for Norway conform reasonably well with the results reported by Dyrstad and Johansen (1999), who report a long run elasticity of wages with respect to regional unemployment of –0.02 for a pooled sample of union and nonunion workers in the manufacturing sector.

Regional fixed effects are included in all our regressions, following the arguments by Blanchflower and Oswald that wage curve estimates otherwise will be biased by a correlation between permanent regional wage determinants and the unemployment rate. When we omit the regional dummies from the regressions, we find that estimates indeed are sensitive to the inclusion of fixed regional effects.[5]

When we compare the group-specific unemployment elasticities across countries, it appears that the nonunion male elasticity is higher (in absolute value) in Great Britain and Norway than it is in the United States. This is surprising in light of the much larger union sectors in Great Britain and Norway, which we would expect to have spillover effects on nonunion wages. When union wages do not respond to local labour market conditions, the total effect on nonunion wages might be expected to be lower when spillover is dominant. A possible explanation for the unexpected result is that the static model is mis-specified and that it ignores potential dynamics in the process of wage formation. If the relationship between past

[5] Consistent with the findings for the United States of Blanchflower and Oswald (1994), the coefficient of log unemployment tends to be positive when regional dummies are excluded from the regression model. Moreover, at least for the nonunion samples, inclusion of regional fixed effects in the regression model is necessary to uncover a negative wage curve relationship. For three out of four U.S. subsamples, the coefficient of unemployment turns significantly more negative (falls to zero in the case unionised males) when the regression includes regional effects—indicating positive long-term relationships between regional unemployment and wages for these samples. Similarly, in the samples from Great Britain estimated wage curve elasticities are more negative when the regression accounts for regional effects, although differences between estimates obtained with and without regional effects are not statistically significant. When regional dummies are removed from the equation, we find a more elastic wage curve for nonunion than union workers of both genders in Norway as well.

and current wages differs by union status and/or across countries, the conclusions drawn from the static model estimates in Table 2 will be misleading. We therefore proceed by exploring wage dynamics in greater detail.

6. Wage dynamics – A Wage or a Phillips Curve?

The static wage curve model in section 5 is based on the premise that the entire impact of changes in unemployment on the level of wages takes place within one period (that is, one year). Blanchflower and Oswald (1994) claim that data for many countries, including the United States and Great Britain, support this hypothesis as the estimated coefficient on the lagged wage typically is close to zero. There is little consensus on this issue, however. For example, Card (1995) suggests that, " . . . the reports of the death of the Phillips curve are premature. More evidence on the dynamic relation between wages and unemployment will probably be required before economists disavow Phillip's hypothesis," (p. 795). Indeed, Blanchard and Katz (1997) and Staiger *et al.* (2001) conclude that wage dynamics are present in the United States, and consequently, that the static model of Blanchflower and Oswald is mis-specified. Similarly, Bell, Nickell and Quintini (2000) find that dynamics are important in the United Kingdom, that is, that regional labour market conditions affect wages over a number of years.

Our study of wage dynamics is based on average wages at the regional level. We first estimate a series of annual, cross-sectional versions of equation (22). These regressions include the regional dummies of the equation but omit time dummies and regional unemployment. The estimated coefficients of the regional dummies (region *X* year) thus constitute a panel data set of average regional wages adjusted for differences in observable wage determinants over time and across regions.

A simple test for a Phillips curve versus a wage curve specification is suggested by Blanchflower and Oswald (1995) and also used by Blanchard and Katz (1997).[6] Consider the following extension of the aggregated model

$$w_{rt} = \alpha + \rho w_{rt-1} + \lambda \log u_{rt} + d_r + f_t + v_{rt} \qquad (23)$$

[6] Card (1995) suggests an alternative test discussed below.

where w_{rt} is the adjusted regional log wage as described above. The autoregressive coefficient, ρ, captures the dynamics of the model. Note that if ρ=1, equation (23) is a standard Phillips curve, while if ρ=0, the equation is a *static* wage curve. Also note that the intermediate case ($0<\rho<1$) represents a *dynamic* wage curve, in which the impact of changes in (regional) unemployment lasts for more than one period. Consequently, the long-run unemployment elasticity is given by $\lambda/(1-\rho)$. An implication is that comparisons of the (static) unemployment elasticities, such as those in Table 2, are misleading if wage dynamics (i.e., the ρ's) differ across groups.

There is a long-standing issue in the empirical literature on wage dynamics. As stressed by Blanchard and Katz (1999, p.71), there are important differences in the observed wage dynamics in the United States and Europe: "The coefficient on the error correction term for the United States is close to zero with point estimates that are typically wrong signed . . . Put another way, the Phillips curve specification . . . appears to provide a good description of the data. In most European countries, however, the error correction term comes in with a significant, and right signed coefficient".

Blanchard and Katz interpret the U.S./European dissimilarities in wage dynamics as resulting from underlying differences in the direct effect of productivity on wages, although they do not provide an elaborate analysis based on a micro model. The theoretical model of the present paper may, however, generate results in line with the observations of Blanchard and Katz. Consider the following simple modification to the turnover cost function in equation (1), with turnover costs proportional to lagged wages:

(24) $$C = H\left(\frac{W}{(1-u)\overline{\overline{W}}}\right)^{-\eta} = hW_{-1}\left(\frac{W}{(1-u)\overline{\overline{W}}}\right)^{-\eta}$$

where $0<h<1$. Two types of mechanisms may be behind such a relationship. First, turnover costs may be linked to last period's wages, e.g., through contractual arrangements, or secondly, the fraction of workers actively searching for a new job may depend on the level of lagged outside wages. We have from (5):

(25) $$W = b(y-C) + (1-b)\eta C$$

It follows that the elasticity of the current wage with respect to the lagged wage is given by:

(26) $$\rho = \frac{(1-b)\eta C - bC}{b(y-C)+(1-b)\eta C} = 1 - b\frac{y}{W}$$

The elasticity of wages with respect to lagged wages, ρ, is positive and is decreasing in β. The minimum of ρ occurs in the monopoly union case; $\beta=1$ and $\rho=C/y\text{-}C$. The maximum is reached in the nonunion case where $\rho=1$. In other words, the nonunion case generates a Phillips curve relationship, while the union case is characterised by a dynamic wage curve (if any at all).

Thus, we obtain the following empirical predictions. First, in the log wage regression of equation (23), the coefficient of lagged log wages should be smaller in the union sector than in the nonunion sector. Second, the coefficient of log unemployment should be less negative for union workers than for nonunion workers. Given the stronger presence of unions in Europe compared to the United States, these predictions are consistent with the observations by Blanchard and Katz described above. We proceed with a more direct test of the model, comparing the results from simple dynamic models that distinguish between union and nonunion sectors.

In Table 3, we report estimates of equation (23) using regional/state means. As expected, the static regional models (in which ρ is restricted to equal 0) produce unemployment elasticities very similar to those we find in the micro level regressions of Table 2. The dynamic specification produces different results for the three countries. Starting with the United States, we first note that, when we pool union and nonunion workers, both the estimated wage dynamics (i.e., the coefficient of the lagged wage) and the effect of local unemployment are lower in our data than in Blanchard and Katz (1997, Table 2, column 4). Further checks (not reported) reveal that this is due primarily to the sample period; when we restrict samples to 1991 and earlier (as in Blanchard and Katz), coefficient estimates become more in line with those of the prior study.[7] The coefficient of the lagged wage is significantly less than unity, but far away from zero, for both men and women. For the United States, both the Phillips curve and the static wage curve are rejected in the pooled samples. The log-run unemployment elasticity is much higher than reported by the static model, –0.122 versus –0.028 for males and –0.112 versus –0.010 for females.

When we estimate the dynamic wage model separately by union status and gender, results support two central predictions of the theoretical model. First, looking at the U.S. long-run elasticities, we see that the linkage between wages and local unemployment is significantly more negative in the nonunion sector than in the union sector. There is no wage curve for union males and wages of female union workers are positively related to the state unemployment rate. In other words, the key finding of the previous section that wages in the nonunion sector are more sensitive to local labour market conditions than are wages in union sector is unaffected by the addition of wage dynamics to the empirical model. Second, the results reveal that the elasticity of wages with respect to lagged wages, ρ, is larger in the nonunion sector than in the union sector.[8] This result is consistent with the prediction in equation (26) that the elasticity declines with higher bargaining power of workers. Thus, accounting for wage dynamics amplify the union-nonunion differences in unemployment elasticities uncovered by the static model.

In Great Britain, wage dynamics appear to differ from those in the United States as the coefficient on the lagged dependent variable is close to zero for all groups. This gives rise to a static wage curve, consistent with the findings of Blanchflower and Oswald, but seems at odds with results reported by Bell *et al.* (2000) based on data from the British New Earnings Survey. Bell *et al.* study regional wages over the period 1976-1997 and find an average long-run elasticity of regional wages with respect to unemployment in the range –0.09 to –0.13 and the lagged dependent variable coefficient is between 0.50 and 0.75.[9] Unfortunately, our seven-year panel is likely too short to avoid downward bias in estimates of wage dynamics (see Nickell, 1981). Furthermore, wage formation may have changed in Britain, in that the dynamics are

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[7] The implication is that both the autoregression in hourly wages and the link to local unemployment have weakened over time. Contributing less importantly to discrepancies with the prior study are differences in samples, model specifications, and aggregation of industry controls.

[8] Part of the union/nonunion differences may be the result of greater bias from measurement error in the lagged dependent variable in union samples. When we alternatively use the lagged *pooled* wage (from columns 1 and 4) in the regressions—so that the extent of measurement error is similar across union and nonunion regressions—results again show that the coefficient of the lagged wage is significantly lower for union workers.

[9] A major conclusion of Bell *et al.* is that failure to control for unobserved region-specific trends generates an upward bias in the coefficient estimate of the lagged wage. When we follow Bell *et al.* and include regional trends in the regression, we too find that the lagged dependent variable coefficient becomes smaller. Because we fail to uncover wage dynamics in the UK data in the first place, however, accounting for regional trends does not affect our conclusions for Great Britain. In the U.S. data, coefficient estimates of the lagged wage fall by approximately 16 percent in the nonunion sector and 40 percent in the union sector when we include state-specific trends, but, again, none of the conclusions of this section are impacted by this experiment. To the contrary, differences between the union and nonunion sectors are reinforced.

different in the 1990s compared to the previous two decades. Separate estimations for the three seven-year periods of the regional panel used by Bell *et al.* (2000) yield coefficient estimates of the lagged dependent variable in the wage equation of -0.26, -0.22 and 0.34,[10] suggesting that sluggishness has increased rather than decreased in Britain. From this we conclude that our panel from the United Kingdom is too short to obtain reliable estimates of the wage dynamics. Because similar results are obtained from analyses of different data, the lack of dynamics in our samples is not a peculiar feature of our data, nor does it indicate significant change in wage dynamics in the 1990s.

For our purposes, the main point is that the dynamic model suggests an even larger difference between the nonunion and union sector than the difference observed in the static model. For females, there exists, as for female union workers in the United States, a positively sloped wage curve but the precision of the estimates is too low to make any strong conclusions.

As for Great Britain, the Norwegian estimates of wage dynamics are based on a relatively short time period and are therefore less precise. The dynamics differ from those of the other two countries, however, as the estimate of ρ is negative in each of the Norwegian samples. The negative coefficient of the lagged wage may reflect the impact of centralised wage setting in which wages within regions tend to adjust to an economy-wide average. In particular, regions that experience low wages in one period (perhaps because of rising local unemployment) may experience a catch up in wages in the next period due to adjustment regulations in central wage settlements.

A cross-country comparison based on the long-run elasticities in Table 3 shows that point estimates of the male unemployment elasticity are highest in the United States, followed by Great Britain. Norway has a considerably lower regional long-run unemployment elasticity. Comparing estimates by union status, however, reduce differences across countries. To some extent, the overall differences across countries in male wage flexibility are explained by differences in unionisation. There appears to be a static wage curve in Great Britain and a dynamic wage curve in the United States. Our results provide support for a "Phillips-curve-like dynamic wage curve" in the United States, as suggested by Blanchard and Katz, but not in Great Britain and Norway.

[10] We are very grateful to Stephen Nickell who provided us with these estimates.

Card (1995) outlines an alternative strategy to test between the Phillips curve and the wage curve, based on a first-differenced version of the adjusted aggregated model:

(27) $$\Delta w_{rt} = w_{rt} - w_{rt-1} = \lambda_0 \log u_{rt} + \lambda_{-1} \log u_{rt-1} + \Delta f_t + \Delta \nu_{rt}$$

The Phillips curve implies that λ_{-1}=0, while the *static* wage curve implies that λ_{-1}+ λ_0=0 (i.e., no change in the regional wage component if unemployment remains constant in two consecutive periods). While the presence of both a lagged dependent variable and a regional fixed effect, together with possible serial correlation in the regional error term in equation (23), create problems - see Card (1995, p. 795) - the dynamic wage curve specification is not nested by the Card-specification. Consequently, his proposed strategy is to test the static version of the wage curve.

Starting again with the CPS data, Table 4 reveals that for the pooled sample of U.S. males, both hypotheses are rejected. Consistent with Table 3, this suggests that a dynamic wage curve is the best representation of the U.S. data. Splitting by union status reveals that the Phillips curve is rejected for nonunion males, but not for union members. However, the point estimate of the current unemployment remains insignificantly different from zero for male union workers. For U.S. females, we fail to reject the Phillips curve for nonunion workers, while the static wage curve is overwhelmingly rejected. In the female union case, however, the picture is mixed, with little support for either hypothesis.

For Great Britain and Norway, the data are not sufficiently rich – that is, the panel is too short - to provide reliable insight into wage dynamics. Norwegian males represent the only case for which standard errors are sufficiently low to make meaningful inference and here we reject the Phillips curve at the ten percent level (p=0.056, last row Table 4). However, point estimates support the static wage curve specification for all groups in both Great Britain and Norway as the sum of unemployment coefficients are close to zero. The lagged unemployment coefficients are far from zero, indicating that Phillips curve specifications do not represent the data very well.

7. Unemployment elasticities by age group

The motivation for studying wage curves separately for young and old workers is threefold. First, we wish to investigate whether the differences we have uncovered by union status can be attributed to composition effects, that the union samples consist mainly of older workers for whom wages are less affected by local labour market conditions. Table 5 shows that this is not the case. Even within age groups, union wages are typically less responsive to local unemployment rates than are nonunion wages. (The exceptions are for young males in Norway where both estimates are around –.10, for young females in Britain where the standard errors are huge and for older British males.) Across groups and countries, the average wage curve elasticity is a statistically significant –.053 for the nonunion sector and an insignificant –.008 for the union sector. The difference between the average elasticities is significant at the one percent level, with a t-statistic of 2.3.

Second, the efficiency wage framework implies that in the nonunion sector, the wages of younger workers should be more sensitive to local unemployment than wages of older workers. This follows from the assumption that young persons' turnover behaviour is more sensitive to relative wages than is turnover of older workers. Older workers' positions are more settled, either because of good matches or through a larger stock of firm-specific capital, and they are thus less likely to change jobs following adjustments in relative wages. This prediction is clearly supported by the data. Among nonunion workers, the coefficient for younger workers is consistently higher (in absolute terms), with British females again being the exception.

Third, our theory predicts that the differences between unemployment elasticities in the nonunion and union sectors are larger among younger workers. For American and British males, as well as for Norwegian women this is certainly true, but not so for the rest of the groups. However, for American females, only the wages of young nonunion workers are negatively impacted by local unemployment.

The empirical patterns are similar in the dynamic specifications, see Table 6 and 7, although the problems associated with huge standard errors in Great Britain and Norway are accentuated when samples are split by age group.

8. Conclusions

We have explored how bargaining power of workers affects the slope of the wage curve in an efficiency wage model based on endogenous turnover costs. The theoretical framework we develop predicts a negatively sloped wage curve when wages are set unilaterally by the firm.

When workers can impose a loss on the firm during bargaining through union action, wages are likely to be less responsive to changes in local labour market conditions. The intuition for this result is quite simple. When local unemployment increases, there are conflicting effects on wages. On the one hand, the efficiency wage paid by the firm decreases as wage turnover is lower with higher unemployment. This effect is augmented when there is bargaining as the fall in the alternative wage leads to a lower bargained wage. On the other hand, however, the reduced quit rate raises the revenue of the firm, net of turnover costs, and this has a positive impact on the bargained wage. The relative magnitude of these effects will depend upon the relative bargaining power of the union and the firm: but it follows from the theoretical analysis that the wage will tend to be less downward sensitive to unemployment the greater is the union's bargaining power.

A number of predictions from the theoretical model are tested by means of micro data from the United States, Great Britain and Norway. Amongst key results, we find first that male union wages are less responsive than nonunion wages to local labour conditions in all three countries. In fact, we find no statistically significant wage curves for union workers. These results suggest that efficiency wage mechanisms, rather than collective bargaining, are driving the negative relationship between wages and local labour market conditions. Second, we find that the differences between the sectors are even larger when allowing for potential wage dynamics. Third, we show that the different wage curves for union and nonunion workers cannot be explained by age composition as the difference exists even within age groups. Fourth, within the nonunion sector, younger workers have steeper wage curves. This observation represents further evidence in support of the efficiency wage explanation of wage flexibility. Fifth, there is evidence - consistent with the theoretical predictions - that Norwegian regional wages respond less to differences in local unemployment rates than is the case in the United States and Great Britain. Finally, our results suggest that there is a static wage curve for both Britain and Norway, while the U.S. results suggest a wage curve with dynamics. For the United States, both the pure static wage curve model and the pure Phillips curve specifications are rejected by the data.

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Table 1: Union and Nonunion Wage Curve Evidence

Study	Nonunion	Union	Regional Effects?	Comments
Blanchflower and Oswald (1994)	-.124 (.033)	.011 (.039)	Yes	Great Britain 1985-89 Hourly Wage
	-.086 (.005)	-.071 (.009)	No	United States 1987 Weekly Earnings
Blanchflower and Oswald (1995)	-.182 (.056)	-.093 (.053)	Yes	Great Britain 1983-91 Weekly Earnings Private Sector
	-.125 (.028)	-.076 (.061)	Yes	United States 1982-87 Annual Earnings Private Sector
Turunen (1998)	-.109 (.005)	-.029 (.012)	No	United States 1979-91 Hourly Wage NLSY
	-.061 (.010)	-.060 (.021)	Yes	
Albæk *et al*. (1999)	-.043 (.018)	-.084 (.005)	No	Denmark 1980-90 Hourly Wage Private Sector
	-.114 (.021)	-.100 (.010)	No	Finland 1989, 91, 93 Hourly Wage Private Sector
	-.076 (.021)	-.049 (.020)	No	Norway 1989, 91, 93 Hourly Wage Private Sector
	-.071 (.031)	-.051 (.013)	No	Sweden 1980, 91 Hourly Wage Private Sector

Notes: Table lists the elasticity of earnings or wage with respect to regional unemployment rate. Standard errors are reported in parentheses. Great Britain data source is BSAS and U.S. data source is CPS, unless indicated otherwise.

Table 2: Wage Curve Elasticities by Gender and Union Status

	United States		Great Britain	Norway
Data source:	CPS	NLSY	BHPS	Register data
Period:	1983-2000	1979-98	1991-97	1991-97
Unemployment measure:	Monthly state unemployment rate averaged over 12 months prior to observation	Like CPS	Travel to work area	County-level unemployment rate
Region:	State (51)	State (51)	Travel to work area (333)	County (20)
Males				
All	-.0277	-.0714	-.1010	-.0466
	(.0031)	(.0139)	(.0294)	(.0177)
Nonunion	-.0403	-.0832	-.1056	-.0660
	(.0033)	(.0153)	(.0390)	(.0242)
Union	.0104	-.0354	-.0747	-.0169
	(.0054)	(.0272)	(.0523)	(.0262)
Females				
All	-.0090	-.0259	-.0177	-.0159
	(.0030)	(.0132)	(.0316)	(.0229)
Nonunion	-.0157	-.0328	-.0033	-.0453
	(.0031)	(.0141)	(.0401)	(.0343)
Union	.0755	.0219	-.0102	.0097
	(.0085)	(.0324)	(.0563)	(.0268)
Union fraction				
Males	.164	.194	.387	.493
Females	.073	.122	.295	.498
Observations				
Nonunion males	1 036 713	38 141	7 248	39 300
Union males	203 489	9 189	4 581	38 237
Nonunion fem	1 031 429	36 775	6 541	17 138
Union females	81 571	5 105	2 739	16 988
Controls	Schooling, experience and its square, and indicator variables for marital status, two races, SMSA, part-time status, 38 industries, 11 months, 16 years, and 50 states	Schooling, experience and its square, seniority and its square, and indicator variables for two races, SMSA, part-time status, 38 industries, 11 months, 16 years, and 50 states	Age and its square and indicator variables for 11 educational qualifications, marital status, part-time work, 8 plant sizes, 9 industries, 6 years, and 216 TTWAs	Schooling, experience and its square, seniority, 27 industries, 6 years, and 19 counties

Notes: Data sources are the 1983-2000 Outgoing Rotation Groups of Current Population Survey (col. 1); 1979-98 waves of National Longitudinal Survey of Youth (col. 2); 1991-97 waves of British Household Panel Survey (col. 3); and 1991-97 register-based employer-employee data, Statistics Norway (col. 4). Samples are restricted to private-sector wage and salaried workers. Robust standard errors, clustered within region-year, are reported in parentheses. Dependent variable is log hourly wage. Complete results are contained in Tables A1-A4.

Table 3: Dynamic Wage-Unemployment Relationships

	Males			Females		
	All	**Nonunion**	**Union**	**All**	**Nonunion**	**Union**
<u>USA (CPS)</u>						
Static model						
Log unemployment (λ)	-.0346	-.0442	.0121	-.0116	-.0183	.0727
	(.0059)	(.0060)	(.0072)	(.0058)	(.0058)	(.0098)
Dynamic model						
Lagged wage (ρ)	.7207	.6836	.4396	.7537	.7443	.2262
	(.0207)	(.0222)	(.0309)	(.0206)	(.0210)	(.0350)
Log unemployment (λ)	-.0340	-.0381	.0005	-.0276	-.0289	.0507
	(.0036)	(.0040)	(.0064)	(.0034)	(.0036)	(.0103)
Long-run elast. (λ/(1- ρ))	-.1217	-.1204	.0009	-.1121	-.1130	.0655
<u>Great Britain</u>						
Static model						
Log unemployment (λ)	-.0963	-.0944	-.0871	.0403	.0409	.0397
	(.0311)	(.0407)	(.0513)	(.0342)	(.0435)	(.0663)
Dynamic model						
Lagged wage (ρ)	-.0038	.0249	.1022	-.0405	.0029	.0524
	(.0392)	(.0406)	(.0414)	(.0384)	(.0381)	(.0497)
Log unemployment (λ)	-.0989	-.1052	-.0606	.0293	.0389	-.0664
	(.0337)	(.0444)	(.0535)	(.0357)	(.0453)	(.0743)
Long-run elast. (λ/(1- ρ))	-.0985	-.1079	-.0674	.0282	.0390	-.0700
<u>Norway</u>						
Static model						
Log unemployment (λ)	-.0331	-.0550	.0083	-.0336	-.0653	.0010
	(.0189)	(.0284)	(.0291)	(.0288)	(.0397)	(.0363)
Dynamic model						
Lagged wage (ρ)	-.2792	-.0086	-.1416	-.3579	-.2646	-.2980
	(.0632)	(.0954)	(.0986)	(.0701)	(.0953)	(.1074)
Log unemployment (λ)	-.0408	-.0551	.0030	-.0263	-.0582	.0093
	(.0183)	(.0286)	(.0292)	(.0271)	(.0384)	(.0352)
Long-run elast. (λ/(1- ρ))	-.0319	-.0546	.0026	-.0194	-.0460	.0072

Notes: Estimates are based on the two-step methodology of Blanchard and Katz (1997). The dependent variable is the average regional residual from year- and group-specific log wage regressions using micro data (and sampling weights where applicable); control variables of these first-step regressions are those listed in Table 2, except for year and region indicator variables. Second-step regressions include year and region fixed effects and are weighted by the cell count of the first-step sample. There are 867 observations in the U.S. and 114 in the Norwegian second-step samples; for Great Britain, observation counts are 835 (col. 1), 782 (col. 2), 679 (col. 3), 805 (col. 4), 757 (col. 5) and 596 (col. 6).

Table 4: Dynamic Wage and Unemployment Relationships—Wage or Phillips Curve?

	Males			Females		
	All	**Nonunion**	**Union**	**All**	**Nonunion**	**Union**
USA (CPS)						
Log $u_{r,t}$ (λ_0)	-.0371	-.0398	-.0176	-.0231	-.0251	.0212
	(.0073)	(.0080)	(.0145)	(.0068)	(.0071)	(.0240)
Log $u_{r,t-1}$ (λ_{-1})	.0203	.0228	.0050	.0055	.0079	-.0390
	(.0073)	(.0080)	(.0144)	(.0068)	(.0070)	(.0239)
P-values:						
$\lambda_0 = 0$	.000	.000	.227	.001	.000	.376
$\lambda_0 + \lambda_{-1} = 0$ (Wage c.)	.000	.000	.010	.000	.000	.032
$\lambda_{-1} = 0$ (Phillips c.)	.005	.004	.727	.415	.263	.103
Great Britain						
Log $u_{r,t}$ (λ_0)	-.0736	-.1128	-.0139	-.0292	-.0329	-.0692
	(.0532)	(.0711)	(.0799)	(.0593)	(.0743)	(.1118)
Log $u_{r,t-1}$ (λ_{-1})	.0815	.1360	.0149	.0371	.0524	.0515
	(.0583)	(.0790)	(.0851)	(.0649)	(.0816)	(.1209)
P-values:						
$\lambda_0 = 0$	.167	.113	.862	.622	.658	.536
$\lambda_0 + \lambda_{-1} = 0$ (Wage c.)	.593	.241	.966	.631	.350	.549
$\lambda_{-1} = 0$ (Phillips c.)	.162	.086	.862	.568	.521	.670
Norway						
Log $u_{r,t}$ (λ_0)	-.0631	-.0631	-.0349	-.0502	-.1039	-.0054
	(.0286)	(.0398)	(.0421)	(.0442)	(.0607)	(.0539)
Log $u_{r,t-1}$ (λ_{-1})	.0593	.0611	.0325	.0410	.0904	-.0024
	(.0307)	(.0426)	(.0450)	(.0475)	(.0656)	(.0574)
P-values:						
$\lambda_0 = 0$	.030	.116	.409	.259	.090	.920
$\lambda_0 + \lambda_{-1} = 0$ (Wage c.)	.701	.879	.874	.545	.517	.871
$\lambda_{-1} = 0$ (Phillips c.)	.056	.154	.472	.390	.171	.967

Notes: Tests are based on the procedure outlined in Card (1995). The dependent variable is the first difference of the state wage effects used in Table 3. Regressions are weighted by $n_{r,t} * n_{r,t-1} / (n_{r,t} + n_{r,t-1})$, where $n_{r,t}$ denotes the group-specific cell size of region *r* in the first-step regression from year *t*. Sample sizes are as listed in Table 3.

Table 5: Wage Curve Elasticities by Age, Gender and Union Status

	United States (CPS)	Great Britain	Norway
Young Males (16-29)			
Nonunion	-.0858 (.0044)	-.2192 (.0641)	-.0947 (.0413)
Union	-.0127 (.0111)	.0011 (.0914)	-.1122 (.0380)
Old Males (30-64)			
Nonunion	-.0142 (.0039)	-.0367 (.0479)	-.0468 (.0289)
Union	.0129 (.0058)	-.0885 (.0563)	.0041 (.0294)
Young Females (16-29)			
Nonunion	-.0482 (.0042)	.0655 (.0621)	-.0567 (.0653)
Union	.0321 (.0161)	-.0062 (.0775)	.0062 (.0620)
Old Females (30-64)			
Nonunion	-.0040 (.0038)	-.0856 (.0554)	-.0097 (.0570)
Union	.0851 (.0095)	-.0346 (.0779)	.0152 (.0310)

Note: See Table 2 for details.

Table 6: Dynamic Wage-Unemployment Relationships by Age and Union Status—Males

	Young (16-29)		Old (30-64)	
	Nonunion	**Union**	**Nonunion**	**Union**
USA (CPS)				
Static model				
Log unemployment (λ)	-.0893	-.0133	-.0166	.0172
	(.0077)	(.0136)	(.0061)	(.0072)
Dynamic model				
Lagged wage (ρ)	.6297	.2418	.5604	.3533
	(.0229)	(.0336)	(.0282)	(.0328)
Log unemployment (λ)	-.0624	-.0124	-.0228	.0080
	(.0055)	(.0133)	(.0049)	(.0068)
Long-run elasticity (λ/(1- ρ))	-.1685	-.0163	-.0518	.0123
Difference	-.1522		-.0641	
Great Britain				
Static model				
Log unemployment (λ)	-.2470	.0965	-.0258	-.1088
	(.0710)	(.1180)	(.0506)	(.0603)
Dynamic model				
Lagged wage (ρ)	-.0045	.0694	.0157	.0872
	(.0434)	(.0600)	(.0415)	(.0422)
Log unemployment (λ)	-.2996	.1697	-.0524	-.1190
	(.0844)	(.1419)	(.0530)	(.0636)
Long-run elasticity (λ/(1- ρ))	-.2983	.1824	-.0533	-.1304
Difference	-.4807		.0771	
Norway				
Static model				
Log unemployment (λ)	-.0894	-.1053	-.0382	.0313
	(.0483)	(.0553)	(.0372)	(.0306)
Dynamic model				
Lagged wage (ρ)	-.0562	-.0466	-.0032	-.1334
	(.0991)	(.1051)	(.0987)	(.0306)
Log unemployment (λ)	-.0918	-.1123	-.0382	.0329
	(.0487)	(.0577)	(.0374)	(.0305)
Long-run elasticity (λ/(1- ρ))	-.0869	-.1073	-.0381	.0290
Difference	.0204		-.0671	

Note: Methodology parallels that of Table 3. There are 867 observations in the U.S. and 114 in the Norwegian second-step samples; for Great Britain, observation counts are 584 (col. 1), 376 (col. 2), 704 (col. 3) and 632 (col. 4).

Table 7: Dynamic Wage-Unemployment Relationships by Age and Union Status—Females

	Young (16-29)		Old (30-64)	
	Nonunion	**Union**	**Nonunion**	**Union**
USA (CPS)				
Static model				
Log unemployment (λ)	-.0489	.0253	-.0051	.0832
	(.0067)	(.0177)	(.0061)	(.0108)
Dynamic model				
Lagged wage (ρ)	.6381	.1593	.6537	.1763
	(.0248)	(.0352)	(.0255)	(.0355)
Log unemployment (λ)	-.0410	.0163	-.0212	.0645
	(.0050)	(.0183)	(.0045)	(.0115)
Long-run elasticity (λ/(1- ρ))	-.1133	.0194	-.0612	.0784
Difference	-.1327		-.1396	
Great Britain				
Static model				
Log unemployment (λ)	.0793	.1957	-.0320	.0399
	(.0766)	(.1224)	(.0594)	(.0891)
Dynamic model				
Lagged wage (ρ)	.0263	-.1951	.1046	.0195
	(.0449)	(.0674)	(.0460)	(.0550)
Log unemployment (λ)	.0115	.0595	-.0181	-.0293
	(.0829)	(.1688)	(.0615)	(.1003)
Long-run elasticity (λ/(1- ρ))	.0119	.0498	-.0202	-.0299
Difference	-.0379		.0097	
Norway				
Static model				
Log unemployment (λ)	-.1192	-.0173	-.0137	.0171
	(.0848)	(-.0775)	(.0573)	(.0399)
Dynamic model				
Lagged wage (ρ)	-.0395	-.1304	-.2783	-.1655
	(.1065)	(??)	(.1039)	(.1084)
Log unemployment (λ)	-.1214	-.0084	.0039	.0208
	(.0854)	(.0789)	(.0558)	(.0596)
Long-run elasticity (λ/(1- ρ))	-.1168	-.0074	.0031	.0178
Difference	-.1094		-.0147	

Note: Methodology parallels that of Table 3. There are 867 observations in the U.S. and 114 in the Norwegian second-step samples; for Great Britain, observation counts are 531 (col. 1), 324 (col. 2), 678 (col. 3) and 495 (col. 4).

Table A1: Log Wage Regressions, United States (CPS Samples)

	Males			Females		
	All	Nonunion	Union	All	Nonunion	Union
ln(UR)	-.0277	-.0403	.0104	-.0090	-.0157	.0755
	(.0031)	(.0033)	(.0054)	(.0030)	(.0031)	(.0085)
Union	.1230			.1306		
	(.0013)			(.0018)		
Schooling	.0782	.0803	.0446	.0791	.0799	.0615
	(.0002)	(.0002)	(.0005)	(.0003)	(.0003)	(.0008)
Experience	.0287	.0290	.0262	.0189	.0188	.0193
	(.0001)	(.0001)	(.0003)	(.0001)	(.0001)	(.0004)
Experience^2	-.0004	-.0004	-.0004	-.0003	-.0003	-.0003
	(.0000)	(.0000)	(.0000)	(.0000)	(.0000)	(.0000)
Married	.1257	.1312	.0744	.0421	.0430	.0290
	(.0010)	(.0011)	(.0020)	(.0009)	(.0009)	(.0028)
Black	-.1684	-.1698	-.1405	-.0947	-.0898	-.1061
	(.0016)	(.0018)	(.0030)	(.0014)	(.0015)	(.0039)
Hispanic	-.0910	-.0917	-.0888	-.0557	-.0518	-.0683
	(.0025)	(.0027)	(.0059)	(.0024)	(.0025)	(.0071)
SMSA	.0923	.0972	.0467	.1063	.1077	.0718
	(.0012)	(.0013)	(.0021)	(.0012)	(.0012)	(.0037)
Part-time	-.2390	-.2419	-.1337	-.1592	-.1648	-.0498
	(.0017)	(.0017)	(.0066)	(.0010)	(.0011)	(.0048)
Construction	-.1241	-.1721	.0897	-.2178	-.2238	-.0456
	(.0034)	(.0038)	(.0060)	(.0077)	(.0079)	(.0360)
Food	-.2506	-.2451	-.2387	-.2616	-.2644	-.2970
	(.0038)	(.0045)	(.0064)	(.0075)	(.0078)	(.0306)
Tobacco	.0217	.0152	.0599	-.0057	-.0332	.0156
	(.0136)	(.0155)	(.0246)	(.0218)	(.0251)	(.0486)
Textiles	-.2415	-.2294	-.3530	-.2962	-.2873	-.4307
	(.0050)	(.0054)	(.0111)	(.0078)	(.0080)	(.0323)
Apparel	-.3542	-.3461	-.4253	-.4595	-.4468	-.5501
	(.0069)	(.0076)	(.0147)	(.0074)	(.0077)	(.0307)
Paper	-.1331	-.1141	-.1155	-.1998	-.1991	-.2269
	(.0045)	(.0058)	(.0068)	(.0087)	(.0094)	(.0317)
Printing	-.2061	-.2195	-.0869	-.2661	-.2682	-.2305
	(.0043)	(.0047)	(.0084)	(.0076)	(.0078)	(.0321)
Chemicals	-.0595	-.0463	-.1137	-.0898	-.0794	-.2194
	(.0041)	(.0046)	(.0073)	(.0080)	(.0082)	(.0324)
Petroleum	.0054	.0424	-.0027	-.0025	-.0054	.0222
	(.0067)	(.0085)	(.0095)	(.0149)	(.0157)	(.0470)
Rubber	-.2115	-.2100	-.2020	-.2918	-.2874	-.3452
	(.0046)	(.0053)	(.0084)	(.0081)	(.0084)	(.0317)
Leather	-.3123	-.2960	-.3978	-.3983	-.3843	-.5161
	(.0101)	(.0113)	(.0212)	(.0099)	(.0105)	(.0343)
Lumber	-.2561	-.2585	-.2443	-.2764	-.2756	-.3204
	(.0045)	(.0050)	(.0088)	(.0093)	(.0098)	(.0342)
Furniture	-.2881	-.2846	-.3244	-.2761	-.2705	-.3613
	(.0048)	(.0052)	(.0103)	(.0083)	(.0086)	(.0327)
Stone	-.2085	-.2061	-.1794	-.2333	-.2253	-.2782
	(.0046)	(.0055)	(.0074)	(.0092)	(.0100)	(.0319)
Prim Metal	-.1783	-.1494	-.1659	-.1827	-.1799	-.2101
	(.0042)	(.0053)	(.0063)	(.0094)	(.0100)	(.0334)
Fabr Metal	-.2117	-.2090	-.1875	-.2527	-.2526	-.2807

	(.0039)	(.0044)	(.0069)	(.0079)	(.0083)	(.0312)
Indust Metal	-.1173	-.1087	-.1507	-.1450	-.1412	-.2162
	(.0036)	(.0040)	(.0064)	(.0075)	(.0078)	(.0316)
Elect Equip	-.1091	-.0947	-.1801	-.2227	-.2239	-.2418
	(.0039)	(.0044)	(.0071)	(.0074)	(.0077)	(.0305)
Transp Equip	-.0783	-.0609	-.0445	-.0919	-.1051	-.0724
	(.0037)	(.0044)	(.0060)	(.0079)	(.0083)	(.0307)
Instruments	-.0969	-.0968	-.1824	-.2115	-.2049	-.3161
	(.0048)	(.0052)	(.0132)	(.0081)	(.0083)	(.0328)
Misc Manuf	-.2619	-.2629	-.2721	-.3636	-.3571	-.4503
	(.0058)	(.0062)	(.0137)	(.0083)	(.0086)	(.0332)
Transportation	-.1975	-.2517	-.0408	-.2264	-.2542	-.1032
	(.0036)	(.0042)	(.0061)	(.0074)	(.0077)	(.0313)
Comm	-.0591	-.0511	.0013	-.0685	-.0572	-.0923
	(.0041)	(.0050)	(.0066)	(.0075)	(.0080)	(.0303)
Utilities	-.0044	.0168	.0301	-.0458	-.0371	-.0729
	(.0039)	(.0047)	(.0064)	(.0082)	(.0087)	(.0314)
Wholesale	-.2238	-.2285	-.2020	-.2533	-.2506	-.3495
	(.0035)	(.0039)	(.0069)	(.0072)	(.0074)	(.0325)
Retail Trade	-.3947	-.4041	-.3099	-.5150	-.5166	-.4888
	(.0034)	(.0037)	(.0066)	(.0070)	(.0072)	(.0303)
Finance	-.0487	-.0632	-.0993	-.2121	-.2116	-.3026
	(.0042)	(.0045)	(.0232)	(.0071)	(.0073)	(.0324)
Insur, RE	-.1750	-.1817	-.1928	-.1958	-.1944	-.2720
	(.0042)	(.0045)	(.0119)	(.0071)	(.0074)	(.0322)
Busin Service	-.2472	-.2534	-.2846	-.2773	-.2753	-.3707
	(.0039)	(.0042)	(.0097)	(.0071)	(.0074)	(.0319)
Auto Repair	-.2845	-.2965	-.1138	-.3338	-.3340	-.2508
	(.0041)	(.0044)	(.0116)	(.0100)	(.0102)	(.0534)
House Service	-.4251	-.4213	-.2613	-.6440	-.6395	-.6418
	(.0111)	(.0113)	(.1342)	(.0080)	(.0081)	(.0529)
Personal Serv	-.4684	-.4677	-.4661	-.4840	-.4791	-.5438
	(.0045)	(.0049)	(.0097)	(.0072)	(.0074)	(.0311)
Entertainment	-.3442	-.3714	-.0863	-.3965	-.4010	-.3001
	(.0050)	(.0052)	(.0141)	(.0080)	(.0081)	(.0357)
Health Offices	-.1325	-.1435	-.1502	-.2165	-.2139	-.2800
	(.0061)	(.0064)	(.0283)	(.0071)	(.0073)	(.0317)
Hospitals	-.2708	-.2731	-.2512	-.1817	-.1765	-.2068
	(.0042)	(.0045)	(.0095)	(.0071)	(.0073)	(.0304)
Education	-.3379	-.3512	-.1672	-.3684	-.3753	-.2592
	(.0047)	(.0051)	(.0102)	(.0073)	(.0076)	(.0309)
Care	-.4209	-.4309	-.3307	-.4696	-.4697	-.4432
	(.0076)	(.0079)	(.0256)	(.0073)	(.0076)	(.0331)
Other Service	-.1934	-.2112	-.0383	-.2177	-.2181	-.2078
	(.0043)	(.0046)	(.0118)	(.0072)	(.0074)	(.0323)
Feb	.0005	.0004	.0000	.0024	.0033	-.0042
	(.0025)	(.0026)	(.0042)	(.0023)	(.0024)	(.0070)
Mar	.0075	.0061	.0141	.0054	.0065	-.0060
	(.0025)	(.0026)	(.0042)	(.0024)	(.0025)	(.0070)
Apr	.0063	.0064	.0066	.0090	.0100	-.0020
	(.0025)	(.0027)	(.0043)	(.0023)	(.0024)	(.0072)
May	.0077	.0079	.0071	.0100	.0106	.0031
	(.0025)	(.0026)	(.0046)	(.0023)	(.0024)	(.0066)
Jun	.0131	.0132	.0105	.0184	.0188	.0149
	(.0025)	(.0027)	(.0043)	(.0025)	(.0026)	(.0070)
Jul	.0124	.0119	.0133	.0166	.0174	.0091

	(.0026)	(.0027)	(.0042)	(.0024)	(.0024)	(.0069)
Aug	.0089	.0076	.0162	.0173	.0184	.0058
	(.0025)	(.0027)	(.0044)	(.0024)	(.0025)	(.0068)
Sep	.0122	.0116	.0154	.0206	.0217	.0071
	(.0025)	(.0027)	(.0044)	(.0024)	(.0025)	(.0069)
Oct	.0194	.0194	.0186	.0254	.0256	.0255
	(.0025)	(.0026)	(.0043)	(.0024)	(.0025)	(.0070)
Nov	.0210	.0213	.0178	.0240	.0244	.0223
	(.0025)	(.0027)	(.0043)	(.0023)	(.0024)	(.0068)
Dec	.0211	.0203	.0249	.0250	.0261	.0133
	(.0024)	(.0026)	(.0042)	(.0023)	(.0024)	(.0069)
1984	.0243	.0230	.0342	.0370	.0362	.0513
	(.0035)	(.0037)	(.0049)	(.0035)	(.0036)	(.0077)
1985	.0459	.0434	.0649	.0606	.0587	.0938
	(.0036)	(.0037)	(.0050)	(.0036)	(.0037)	(.0082)
1986	.0654	.0622	.0934	.0844	.0822	.1224
	(.0034)	(.0037)	(.0050)	(.0035)	(.0035)	(.0082)
1987	.0884	.0878	.1105	.1215	.1193	.1669
	(.0034)	(.0037)	(.0053)	(.0035)	(.0036)	(.0084)
1988	.1123	.1112	.1338	.1587	.1565	.2048
	(.0036)	(.0038)	(.0058)	(.0037)	(.0038)	(.0091)
1989	.1522	.1531	.1645	.1981	.1953	.2511
	(.0039)	(.0041)	(.0061)	(.0038)	(.0039)	(.0094)
1990	.1914	.1957	.1874	.2505	.2499	.2761
	(.0039)	(.0041)	(.0060)	(.0039)	(.0040)	(.0095)
1991	.2196	.2262	.2087	.2941	.2944	.3045
	(.0037)	(.0039)	(.0058)	(.0036)	(.0037)	(.0091)
1992	.2437	.2514	.2315	.3293	.3315	.3109
	(.0035)	(.0037)	(.0053)	(.0035)	(.0035)	(.0083)
1993	.2608	.2680	.2520	.3568	.3585	.3454
	(.0034)	(.0036)	(.0052)	(.0034)	(.0035)	(.0085)
1994	.2728	.2804	.2583	.3775	.3796	.3615
	(.0036)	(.0038)	(.0059)	(.0037)	(.0038)	(.0090)
1995	.3158	.3246	.2875	.4178	.4191	.4086
	(.0042)	(.0043)	(.0068)	(.0043)	(.0044)	(.0099)
1996	.3176	.3264	.2906	.4235	.4244	.4200
	(.0040)	(.0042)	(.0070)	(.0039)	(.0040)	(.0102)
1997	.3510	.3590	.3276	.4591	.4603	.4466
	(.0041)	(.0042)	(.0071)	(.0040)	(.0041)	(.0105)
1998	.3984	.4093	.3517	.5065	.5080	.4893
	(.0042)	(.0044)	(.0075)	(.0041)	(.0042)	(.0115)
1999	.4380	.4509	.3719	.5447	.5468	.5159
	(.0044)	(.0046)	(.0074)	(.0044)	(.0045)	(.0116)
2000	.4736	.4858	.4040	.5838	.5861	.5493
	(.0045)	(.0047)	(.0076)	(.0046)	(.0047)	(.0120)
AK	.3374	.3299	.3418	.3619	.3682	.2611
	(.0086)	(.0086)	(.0136)	(.0079)	(.0080)	(.0179)
AZ	.0531	.0482	.0760	.0831	.0825	.1457
	(.0052)	(.0057)	(.0142)	(.0054)	(.0056)	(.0244)
AR	-.0344	-.0373	-.0106	-.0030	-.0010	-.0162
	(.0048)	(.0054)	(.0096)	(.0049)	(.0051)	(.0193)
CA	.1527	.1454	.1607	.2160	.2170	.1757
	(.0042)	(.0046)	(.0073)	(.0041)	(.0043)	(.0140)
CO	.0792	.0741	.0692	.1079	.1054	.1515
	(.0055)	(.0059)	(.0116)	(.0055)	(.0056)	(.0204)
CT	.1653	.1690	.0939	.2063	.2143	.0951

	(.0056)	(.0062)	(.0101)	(.0059)	(.0061)	(.0175)
DE	.1073	.1074	.0842	.1432	.1424	.1204
	(.0054)	(.0061)	(.0096)	(.0054)	(.0056)	(.0203)
DC	.1828	.1922	.0808	.2665	.2724	.1656
	(.0064)	(.0072)	(.0155)	(.0059)	(.0062)	(.0192)
FL	-.0008	-.0006	-.0014	.0494	.0525	-.0182
	(.0042)	(.0047)	(.0092)	(.0042)	(.0043)	(.0182)
GA	.0787	.0779	.0496	.1060	.1059	.0839
	(.0051)	(.0055)	(.0112)	(.0050)	(.0053)	(.0196)
HI	.1712	.1547	.1807	.1994	.1979	.1975
	(.0064)	(.0071)	(.0114)	(.0062)	(.0066)	(.0184)
ID	.0494	.0445	.0695	.0655	.0666	.0828
	(.0051)	(.0056)	(.0102)	(.0050)	(.0052)	(.0196)
IL	.1211	.1224	.0852	.1249	.1299	.0456
	(.0040)	(.0046)	(.0069)	(.0041)	(.0043)	(.0145)
IN	.0305	.0242	.0400	.0404	.0399	.0240
	(.0046)	(.0054)	(.0075)	(.0050)	(.0053)	(.0157)
IA	.0101	.0008	.0512	.0231	.0213	.0797
	(.0049)	(.0055)	(.0087)	(.0050)	(.0052)	(.0183)
KS	.0168	.0069	.0479	.0379	.0380	.0598
	(.0051)	(.0057)	(.0093)	(.0054)	(.0055)	(.0203)
KY	.0101	.0073	.0175	.0363	.0384	-.0006
	(.0050)	(.0056)	(.0089)	(.0052)	(.0055)	(.0173)
LA	.0351	.0393	.0157	.0233	.0266	-.0120
	(.0058)	(.0063)	(.0120)	(.0059)	(.0062)	(.0237)
ME	.0118	.0122	.0151	.0878	.0906	.0533
	(.0052)	(.0058)	(.0094)	(.0050)	(.0051)	(.0202)
MD	.1327	.1331	.1041	.1755	.1747	.1667
	(.0055)	(.0060)	(.0102)	(.0058)	(.0060)	(.0179)
MA	.1445	.1503	.0695	.2002	.2035	.1431
	(.0050)	(.0053)	(.0085)	(.0051)	(.0053)	(.0156)
MI	.1010	.1025	.0904	.0934	.0955	.0447
	(.0039)	(.0045)	(.0069)	(.0042)	(.0045)	(.0143)
MN	.0978	.0864	.1086	.1341	.1323	.1277
	(.0051)	(.0057)	(.0084)	(.0051)	(.0053)	(.0160)
MS	-.0102	-.0050	-.0377	.0095	.0134	-.0573
	(.0051)	(.0056)	(.0111)	(.0049)	(.0051)	(.0189)
MO	.0182	-.0028	.0622	.0429	.0442	.0097
	(.0051)	(.0058)	(.0082)	(.0050)	(.0052)	(.0166)
MT	.0216	.0041	.0703	.0244	.0234	.0192
	(.0059)	(.0063)	(.0096)	(.0057)	(.0058)	(.0181)
NE	-.0250	-.0395	.0428	.0090	.0055	.1082
	(.0052)	(.0057)	(.0104)	(.0053)	(.0055)	(.0192)
NV	.1229	.1151	.1350	.1551	.1585	.1167
	(.0058)	(.0062)	(.0104)	(.0061)	(.0063)	(.0170)
NH	.1304	.1249	.1072	.1494	.1508	.1234
	(.0057)	(.0063)	(.0129)	(.0058)	(.0060)	(.0224)
NJ	.1733	.1882	.0918	.2033	.2174	.0738
	(.0046)	(.0051)	(.0076)	(.0048)	(.0051)	(.0148)
NM	.0031	.0030	.0245	.0365	.0384	.0303
	(.0054)	(.0060)	(.0120)	(.0056)	(.0058)	(.0210)
NY	.1126	.1240	.0592	.1706	.1858	.0579
	(.0043)	(.0048)	(.0073)	(.0043)	(.0045)	(.0141)
NC	.0414	.0361	.0607	.0657	.0649	.0701
	(.0044)	(.0048)	(.0099)	(.0044)	(.0046)	(.0190)
ND	-.0342	-.0467	.0372	-.0145	-.0140	.0149

	(.0060)	(.0065)	(.0108)	(.0059)	(.0061)	(.0231)
OH	.0469	.0474	.0376	.0604	.0620	.0217
	(.0040)	(.0046)	(.0068)	(.0042)	(.0043)	(.0142)
OK	-.0005	-.0049	.0271	.0163	.0164	.0362
	(.0054)	(.0060)	(.0105)	(.0060)	(.0062)	(.0187)
OR	.0840	.0768	.0998	.1175	.1202	.0700
	(.0051)	(.0058)	(.0089)	(.0053)	(.0054)	(.0172)
PA	.0464	.0549	.0099	.0812	.0883	-.0038
	(.0041)	(.0046)	(.0070)	(.0042)	(.0044)	(.0140)
RI	.0751	.0767	.0310	.1251	.1278	.0712
	(.0055)	(.0061)	(.0103)	(.0054)	(.0057)	(.0171)
SC	.0386	.0356	.0615	.0412	.0423	.0195
	(.0048)	(.0052)	(.0146)	(.0050)	(.0051)	(.0250)
SD	-.0620	-.0699	-.0074	-.0050	-.0039	-.0101
	(.0055)	(.0061)	(.0111)	(.0054)	(.0055)	(.0218)
TN	-.0015	-.0005	-.0160	.0295	.0305	.0138
	(.0049)	(.0054)	(.0096)	(.0048)	(.0051)	(.0178)
TX	.0350	.0334	.0576	.0780	.0804	.0615
	(.0044)	(.0049)	(.0090)	(.0045)	(.0047)	(.0170)
UT	.0377	.0292	.0702	.0565	.0575	.0615
	(.0053)	(.0056)	(.0114)	(.0053)	(.0055)	(.0243)
VT	.0497	.0543	-.0229	.1264	.1301	.0566
	(.0056)	(.0060)	(.0122)	(.0054)	(.0056)	(.0237)
VA	.0897	.0873	.0808	.1140	.1116	.1503
	(.0054)	(.0058)	(.0116)	(.0052)	(.0054)	(.0188)
WA	.1311	.1266	.1438	.1628	.1593	.1514
	(.0050)	(.0058)	(.0084)	(.0051)	(.0055)	(.0162)
WV	-.0213	-.0332	.0089	-.0290	-.0250	-.0885
	(.0057)	(.0062)	(.0088)	(.0055)	(.0057)	(.0187)
WI	.0504	.0450	.0605	.0591	.0609	.0394
	(.0046)	(.0053)	(.0076)	(.0048)	(.0050)	(.0155)
WY	.0827	.0676	.1711	.0385	.0379	.1133
	(.0065)	(.0071)	(.0106)	(.0070)	(.0072)	(.0268)
Constant	.8692	.8625	1.377	.7127	.7096	.9806
	(.0101)	(.0109)	(.0183)	(.0120)	(.0124)	(.0412)
R^2	.5041	.5203	.3707	.4954	.4993	.4174
Observations	1240202	1036713	203489	1113000	1031429	81571

Note: Standard errors are calculated allowing for state-by-period clustering. Dependent variable is the log hourly wage. Omitted state is Alabama, industry is mining, and year is 1983.

Table A2: Log Wage Regressions, United States (NLSY Samples)

	Males			Females		
	All	Nonunion	Union	All	Nonunion	Union
ln(UR)	-.0714	-.0832	-.0354	-.0259	-.0328	.0219
	(.0139)	(.0153)	(.0272)	(.0132)	(.0141)	(.0324)
Union	.1638			.0902		
	(.0047)			(.0056)		
Schooling	.0654	.0669	.0448	.0686	.0690	.0605
	(.0011)	(.0011)	(.0029)	(.0012)	(.0012)	(.0033)
Experience	.0464	.0427	.0643	.0204	.0192	.0304
	(.0020)	(.0022)	(.0045)	(.0020)	(.0021)	(.0055)
Experience^2	-.0005	-.0003	-.0018	.0005	.0006	-.0005
	(.0001)	(.0001)	(.0002)	(.0001)	(.0001)	(.0003)
Tenure	.0543	.0536	.0519	.0554	.0538	.0602
	(.0016)	(.0017)	(.0035)	(.0017)	(.0018)	(.0046)
Tenure^2	-.0026	-.0025	-.0025	-.0028	-.0028	-.0027
	(.0001)	(.0001)	(.0002)	(.0001)	(.0001)	(.0003)
Black	-.1470	-.1511	-.1161	-.0904	-.0914	-.0627
	(.0054)	(.0062)	(.0104)	(.0049)	(.0053)	(.0133)
Hispanic	-.0730	-.0726	-.0748	-.0051	-.0074	.0027
	(.0067)	(.0076)	(.0128)	(.0063)	(.0066)	(.0158)
SMSA	.0957	.0906	.1043	.0997	.1043	.0715
	(.0051)	(.0057)	(.0110)	(.0054)	(.0057)	(.0155)
Part-time	-.0397	-.0451	-.0183	-.0156	-.0206	.0185
	(.0069)	(.0078)	(.0145)	(.0051)	(.0055)	(.0131)
Construction	.0923	.1008	.0299	.1586	.1073	.2789
	(.0170)	(.0181)	(.0445)	(.0338)	(.0359)	(.0782)
Food	-.1319	-.0974	-.2759	-.0460	-.0607	-.1064
	(.0108)	(.0132)	(.0189)	(.0186)	(.0193)	(.0585)
Tobacco	-.1109	-.0729	-.1977	-.1583	-.2836	.2179
	(.0446)	(.0562)	(.0300)	(.0855)	(.0775)	(.2001)
Textiles	-.1547	-.1151	-.3489	-.1132	-.0857	-.3656
	(.0141)	(.0152)	(.0356)	(.0201)	(.0207)	(.0659)
Apparel	-.2702	-.2209	-.5425	-.1816	-.1460	-.3789
	(.0177)	(.0201)	(.0357)	(.0184)	(.0199)	(.0565)
Paper	-.0913	-.0549	-.2123	.0493	-.0028	.0601
	(.0156)	(.0204)	(.0257)	(.0308)	(.0347)	(.0650)
Printing	-.1319	-.1095	-.2161	-.0662	-.0704	-.0188
	(.0141)	(.0157)	(.0295)	(.0206)	(.0213)	(.0720)
Chemicals	-.0393	.0034	-.2	.0293	.0403	-.0888
	(.0150)	(.0167)	(.0319)	(.0217)	(.0235)	(.0652)
Petroleum	.0051	-.0060	.0527	.1503	.1480	.1693
	(.0341)	(.0385)	(.0734)	(.0554)	(.0597)	(.1609)
Rubber	-.1434	-.1203	-.2411	-.0423	-.0497	-.0770
	(.0160)	(.0193)	(.0297)	(.0255)	(.0262)	(.0858)
Leather	-.1752	-.1139	-.4357	-.1286	-.1338	-.1484
	(.0265)	(.0290)	(.0603)	(.0282)	(.0303)	(.0850)
Lumber	-.2451	-.2037	-.4391	-.0666	-.0648	-.0155
	(.0118)	(.0128)	(.0302)	(.0299)	(.0314)	(.0995)
Furniture	-.2238	-.1847	-.3871	-.0810	-.0404	-.3622
	(.0165)	(.0189)	(.0318)	(.0224)	(.0243)	(.0683)
Stone	-.1539	-.1164	-.2580	-.0168	-.0270	-.0972
	(.0143)	(.0155)	(.0321)	(.0275)	(.0308)	(.0713)
Prim Metal	-.0134	-.0001	-.1295	-.0125	-.0590	-.0284

	(.0135)	(.0195)	(.0209)	(.0322)	(.0339)	(.0757)
Fabr Metal	-.1383	-.1024	-.2859	-.0115	.0005	-.1568
	(.0131)	(.0145)	(.0277)	(.0249)	(.0271)	(.0737)
Indust Metal	-.0297	.0164	-.2212	.0492	.0639	-.0936
	(.0103)	(.0117)	(.0209)	(.0198)	(.0209)	(.0599)
Elect Equip	-.0423	.0127	-.2889	-.0149	-.0075	-.0988
	(.0130)	(.0142)	(.0281)	(.0189)	(.0201)	(.0601)
Transp Equip	.0041	.0313	-.0968	.0905	.0598	.1101
	(.0099)	(.0121)	(.0181)	(.0218)	(.0241)	(.0622)
Instruments	.0163	.0526	-.1586	.1013	.1329	-.1643
	(.0217)	(.0238)	(.0510)	(.0268)	(.0287)	(.0688)
Misc Manuf	-.1223	-.0893	-.2513	-.0733	-.0700	-.1293
	(.0141)	(.0155)	(.0332)	(.0223)	(.0231)	(.0809)
Transportation	-.0659	-.0555	-.1092	-.0039	-.0324	.0090
	(.0105)	(.0117)	(.0214)	(.0190)	(.0198)	(.0618)
Comm	-.0750	-.0144	-.2535	.0762	.0733	.0267
	(.0157)	(.0183)	(.0251)	(.0201)	(.0244)	(.0570)
Utilities	.0175	.0206	-.0434	.1135	.1282	.0345
	(.0127)	(.0194)	(.0180)	(.0229)	(.0274)	(.0605)
Wholesale	-.1848	-.1609	-.2793	-.0825	-.0718	-.2084
	(.0091)	(.0103)	(.0211)	(.0180)	(.0186)	(.0608)
Retail Trade	-.3261	-.2983	-.4265	-.2391	-.2301	-.3469
	(.0064)	(.0069)	(.0163)	(.0155)	(.0160)	(.0547)
Finance	-.0176	.0177	-.2216	-.0691	-.0515	-.3029
	(.0167)	(.0178)	(.0442)	(.0164)	(.0171)	(.0659)
Insur, RE	-.0757	-.0395	-.2886	.0058	.0177	-.1632
	(.0141)	(.0147)	(.0400)	(.0168)	(.0174)	(.0567)
Busin Service	-.1921	-.1502	-.3983	-.0500	-.0361	-.2306
	(.0095)	(.0102)	(.0249)	(.0169)	(.0175)	(.0584)
Auto Repair	-.2032	-.1694	-.3643	-.1518	-.1338	-.2961
	(.0103)	(.0111)	(.0307)	(.0258)	(.0275)	(.0900)
House Service	-.2880	-.2664	1.2482	-.6249	-.6133	-.4985
	(.0531)	(.0518)	(.0452)	(.0286)	(.0290)	(.1917)
Personal Serv	-.3838	-.3445	-.5521	-.2122	-.2030	-.2990
	(.0133)	(.0146)	(.0301)	(.0179)	(.0186)	(.0603)
Entertainment	-.2957	-.2899	-.2467	-.1407	-.1319	-.2069
	(.0204)	(.0219)	(.0585)	(.0243)	(.0266)	(.0659)
Health Offices	-.0267	-.0008	-.1119	-.0202	-.0058	-.1850
	(.0340)	(.0368)	(.1017)	(.0167)	(.0174)	(.0622)
Hospitals	-.2704	-.2050	-.5167	-.0497	-.0253	-.2405
	(.0135)	(.0151)	(.0259)	(.0159)	(.0166)	(.0554)
Education	-.2956	-.2530	-.4245	-.2876	-.2885	-.3128
	(.0238)	(.0257)	(.0586)	(.0179)	(.0187)	(.0603)
Care	-.2670	-.2309	-.4189	-.1229	-.1173	-.1186
	(.0399)	(.0397)	(.1458)	(.0350)	(.0365)	(.0783)
Other Service	-.0655	-.0373	-.2152	.0291	.0395	-.1510
	(.0142)	(.0151)	(.0406)	(.0177)	(.0180)	(.0705)
Feb	.0030	.0039	.0128	.0200	.0195	.0116
	(.0102)	(.0109)	(.0194)	(.0094)	(.0102)	(.0216)
Mar	.0163	.0172	.0232	.0258	.0251	.0155
	(.0104)	(.0113)	(.0196)	(.0099)	(.0107)	(.0207)
Apr	.0575	.0619	.0463	.0504	.0512	.0408
	(.0116)	(.0126)	(.0240)	(.0113)	(.0121)	(.0250)
May	.0572	.0576	.0618	.0715	.0737	.0427
	(.0126)	(.0139)	(.0248)	(.0129)	(.0140)	(.0299)
Jun	.0658	.0628	.0791	.0619	.0584	.0765

	(.0137)	(.0150)	(.0284)	(.0146)	(.0157)	(.0335)
Jul	.0865	.0844	.0976	.0838	.0816	.0802
	(.0145)	(.0160)	(.0291)	(.0150)	(.0161)	(.0356)
Aug	.1065	.1046	.1195	.0926	.0832	.1470
	(.0148)	(.0162)	(.0299)	(.0155)	(.0166)	(.0372)
Sep	.1031	.0977	.1263	.1041	.0954	.1554
	(.0155)	(.0171)	(.0319)	(.0166)	(.0178)	(.0442)
Oct	.1336	.1425	.0873	.1082	.0978	.1606
	(.0174)	(.0188)	(.0346)	(.0187)	(.0201)	(.0469)
Nov	.1390	.1490	.1105	.1277	.1133	.1921
	(.0195)	(.0216)	(.0413)	(.0223)	(.0242)	(.0513)
Dec	.1260	.1209	.1292	.0563	.0316	.2485
	(.0304)	(.0337)	(.0557)	(.0342)	(.0363)	(.1200)
1980	.0640	.0900	-.0118	.0672	.0639	.0831
	(.0153)	(.0171)	(.0336)	(.0171)	(.0184)	(.0349)
1981	.1246	.1501	.0494	.1458	.1454	.1403
	(.0150)	(.0161)	(.0344)	(.0156)	(.0173)	(.0337)
1982	.1592	.1800	.1000	.1846	.1870	.1569
	(.0149)	(.0158)	(.0340)	(.0154)	(.0170)	(.0340)
1983	.1466	.1704	.0880	.2060	.2106	.1809
	(.0171)	(.0181)	(.0370)	(.0171)	(.0188)	(.0371)
1984	.1580	.1824	.0964	.2279	.2310	.2184
	(.0151)	(.0158)	(.0345)	(.0161)	(.0174)	(.0383)
1985	.1775	.2039	.1050	.2649	.2658	.2659
	(.0139)	(.0148)	(.0322)	(.0147)	(.0159)	(.0361)
1986	.1923	.2226	.1012	.2807	.2816	.2902
	(.0140)	(.0150)	(.0332)	(.0147)	(.0159)	(.0360)
1987	.1894	.2293	.0669	.2935	.3005	.2599
	(.0148)	(.0162)	(.0355)	(.0162)	(.0176)	(.0382)
1988	.1890	.2199	.1004	.3286	.3413	.2456
	(.0173)	(.0187)	(.0394)	(.0184)	(.0199)	(.0459)
1989	.2084	.2434	.0961	.3700	.3774	.3195
	(.0173)	(.0189)	(.0370)	(.0185)	(.0198)	(.0461)
1990	.2254	.2610	.1066	.3737	.3868	.2781
	(.0178)	(.0192)	(.0400)	(.0192)	(.0206)	(.0487)
1991	.2518	.2891	.1320	.4388	.4512	.3427
	(.0184)	(.0200)	(.0396)	(.0192)	(.0207)	(.0462)
1992	.2593	.2936	.1647	.4236	.4335	.3511
	(.0186)	(.0197)	(.0426)	(.0192)	(.0206)	(.0493)
1993	.2884	.3273	.1671	.4502	.4608	.3712
	(.0183)	(.0202)	(.0408)	(.0197)	(.0212)	(.0478)
1994	.2898	.3140	.2336	.4604	.4661	.4141
	(.0195)	(.0212)	(.0424)	(.0199)	(.0215)	(.0500)
1996	.3133	.3411	.2524	.4857	.4811	.5273
	(.0185)	(.0198)	(.0406)	(.0188)	(.0202)	(.0484)
1998	.3775	.3936	.3700	.5341	.5336	.5388
	(.0189)	(.0203)	(.0436)	(.0186)	(.0200)	(.0462)
AK	.3703	.3325	.4663	.4455	.4420	.5162
	(.0249)	(.0293)	(.0608)	(.0345)	(.0318)	(.1748)
AZ	-.0090	-.0186	.0318	.0550	.0580	.0848
	(.0185)	(.0205)	(.0424)	(.0177)	(.0191)	(.0404)
AR	-.0170	-.0200	-.0761	-.0265	-.0033	-.1494
	(.0209)	(.0227)	(.0517)	(.0180)	(.0208)	(.0384)
CA	.1949	.1781	.2577	.1942	.1980	.1858
	(.0115)	(.0133)	(.0274)	(.0125)	(.0133)	(.0302)
CO	.0871	.0720	.1604	.0813	.0619	.2297

	(.0170)	(.0185)	(.0361)	(.0177)	(.0180)	(.0472)
CT	.1592	.1502	.1800	.2387	.2432	.2293
	(.0176)	(.0188)	(.0379)	(.0192)	(.0197)	(.0497)
DE	.0444	.0417	.1589	.1024	.0953	.0503
	(.0598)	(.0626)	(.0547)	(.0792)	(.0833)	(.0810)
DC	.1511	.1707	.0711	.1990	.1964	.2114
	(.0232)	(.0262)	(.0437)	(.0224)	(.0237)	(.0491)
FL	.0599	.0685	-.0313	.0634	.0716	-.0193
	(.0135)	(.0150)	(.0329)	(.0128)	(.0136)	(.0345)
GA	.0370	.0436	.0086	.0913	.1085	-.0250
	(.0130)	(.0147)	(.0277)	(.0146)	(.0156)	(.0345)
HI	.2257	.2206	.3118	.1143	.1262	.0873
	(.0698)	(.0840)	(.0856)	(.0824)	(.0969)	(.1544)
ID	.1960	.1983	.0963	.0661	.0449	.2564
	(.0535)	(.0469)	(.2362)	(.0627)	(.0722)	(.1211)
IL	.1472	.1266	.2046	.1315	.1384	.0991
	(.0129)	(.0151)	(.0279)	(.0133)	(.0152)	(.0341)
IN	.0658	.0440	.1223	-.0079	-.0001	-.0657
	(.0193)	(.0215)	(.0375)	(.0190)	(.0200)	(.0530)
IA	-.0372	-.0687	.0808	-.0126	-.0267	.1065
	(.0243)	(.0268)	(.0557)	(.0205)	(.0231)	(.0677)
KS	.0334	-.0009	.1606	-.0747	-.0934	.1256
	(.0258)	(.0308)	(.0410)	(.0259)	(.0280)	(.1101)
KY	.0367	.0419	-.0064	-.0383	-.0413	-.0050
	(.0466)	(.0505)	(.0825)	(.0368)	(.0391)	(.0809)
LA	.0793	.0855	.0510	.0514	.0622	-.0911
	(.0260)	(.0280)	(.0674)	(.0209)	(.0220)	(.0720)
ME	-.1206	-.0764	-.2691	.0114	.0079	.1464
	(.0807)	(.0919)	(.0562)	(.0488)	(.0508)	(.0404)
MD	.0767	.0851	.0124	.1386	.1578	.0223
	(.0193)	(.0208)	(.0432)	(.0195)	(.0218)	(.0424)
MA	.1411	.1151	.2589	.1430	.1509	.1133
	(.0199)	(.0217)	(.0446)	(.0188)	(.0202)	(.0400)
MI	.0516	.0397	.1097	.0296	.0253	.0599
	(.0124)	(.0142)	(.0301)	(.0132)	(.0144)	(.0298)
MN	-.0043	-.0296	.1276	.0674	.0627	.0977
	(.0150)	(.0162)	(.0357)	(.0180)	(.0191)	(.0561)
MS	-.0681	-.0542	-.1245	-.0571	-.0440	-.0629
	(.0191)	(.0213)	(.0456)	(.0182)	(.0184)	(.0629)
MO	-.0575	-.0813	.0223	.0454	.0597	-.0641
	(.0145)	(.0179)	(.0320)	(.0155)	(.0163)	(.0403)
MT	.0155	.0072	.0501	-.0652	-.1085	.1240
	(.0226)	(.0283)	(.0533)	(.0337)	(.0345)	(.0730)
NE	-.0473	-.0539	-.0474	-.1139	-.1193	-.0874
	(.0397)	(.0414)	(.1296)	(.0341)	(.0332)	(.1549)
NV	.1910	.1728	.3204	.0947	.1093	-.0479
	(.0296)	(.0329)	(.0792)	(.0282)	(.0310)	(.0606)
NH	.0349	.0312	-.0235	-.0484	-.0506	.
	(.0580)	(.0638)	(.0938)	(.0483)	(.0487)	(.)
NJ	.1808	.1747	.2063	.2078	.2180	.0382
	(.0168)	(.0200)	(.0286)	(.0171)	(.0177)	(3.650)
NM	-.0203	-.0355	.0521	-.0295	-.0513	.0557
	(.0215)	(.0221)	(.0553)	(.0216)	(.0243)	(1.960)
NY	.1391	.1366	.1570	.1441	.1566	.0316
	(.0141)	(.0157)	(.0274)	(.0146)	(.0163)	(2.960)
NC	.0178	.0062	.0878	.0528	.0527	.0439

	(.0132)	(.0151)	(.0369)	(.0149)	(.0154)	(.6000)
ND	.1611	.1854	-.3069	-.0221	-.0194	.
	(.0649)	(.0628)	(.0545)	(.0668)	(.0678)	(.)
OH	.0009	.0004	.0239	-.0019	.0015	-.5500
	(.0123)	(.0138)	(.0264)	(.0127)	(.0134)	(.5840)
OK	.0367	.0408	.0344	.0242	.0194	.7200
	(.0177)	(.0192)	(.0443)	(.0182)	(.0210)	(.4720)
OR	.1270	.1060	.2544	.0567	.0723	-2.410
	(.0250)	(.0284)	(.0561)	(.0310)	(.0303)	(.0160)
PA	.0236	.0250	.0316	.0469	.0498	1.81
	(.0124)	(.0147)	(.0275)	(.0125)	(.0136)	(.0700)
RI	.2403	.3398	-.3277	.2199	.2206	.
	(.1091)	(.1114)	(.1486)	(.0772)	(.0777)	(.)
SC	.0236	.0328	-.0530	.0203	.0301	.3010
	(.0156)	(.0176)	(.0340)	(.0149)	(.0161)	(-.1451)
SD	-.0940	-.0787	-.0879	-.0441	-.0405	.5550
	(.0315)	(.0313)	(.0750)	(.0316)	(.0343)	(-.2883)
TN	-.0364	-.0412	.0002	.0288	.0259	.1020
	(.0148)	(.0160)	(.0451)	(.0140)	(.0150)	(-.0128)
TX	.0352	.0328	.0359	.0490	.0559	.7400
	(.0127)	(.0139)	(.0295)	(.0129)	(.0136)	(-.0596)
UT	.1112	.1453	-.1065	.0900	.0678	.0020
	(.0466)	(.0444)	(.1485)	(.0390)	(.0392)	(.0827)
VT	-.0502	-.0459	-.0977	.0662	.0642	.0750
	(.0254)	(.0264)	(.1286)	(.0278)	(.0286)	(-.0251)
VA	-.0092	.0115	-.0849	-.0192	-.0109	.0520
	(.0147)	(.0171)	(.0313)	(.0156)	(.0169)	(-.1582)
WA	.1199	.0940	.1995	.0661	.0506	.0000
	(.0169)	(.0194)	(.0332)	(.0200)	(.0211)	(.0778)
WV	.0528	.0190	.2211	-.0070	.0055	.1710
	(.0203)	(.0213)	(.0513)	(.0201)	(.0218)	(-.1777)
WI	.0031	-.0094	.0681	-.0179	-.0315	.0210
	(.0143)	(.0170)	(.0297)	(.0152)	(.0158)	(.0140)
WY	.1484	.1355	.3168	.1264	.1410	.
	(.0749)	(.0821)	(.0588)	(.0811)	(.0805)	(.)
Constant	.7201	.6882	1.1693	.3318	.3263	.3672
	(.0341)	(.0370)	(.0759)	(.0380)	(.0412)	(.7588)
R^2	.5675	.5679	.5698	.5830	.5881	.5888
Observations	47330	38141	9189	41880	36775	5105

Note: Standard errors are calculated allowing for state-by-period clustering. Dependent variable is the log hourly wage. Omitted state is Alabama, industry is mining, and year is 1979.

Table A3: Log Wage Regressions, Great Britain

	Males			Females		
	All	**Nonunion**	**Union**	**All**	**Nonunion**	**Union**
lurate	-.1010	-.1056	-.0747	-.0177	-.0033	-.0102
	(.0294)	(.0390)	(.0523)	(.0316)	(.0400)	(.0563)
d_un	.0503			.0651		
	(.0086)			(.0095)		
age	.0705	.0787	.0569	.0527	.0490	.0659
	(.0022)	(.0031)	(.0037)	(.0024)	(.0029)	(.0051)
sqrage	-.0008	-.0009	-.0006	-.0006	-.0006	-.0008
	(.0000)	(.0000)	(.0000)	(.0000)	(.0000)	(.0001)
d_mrdv	.1376	.1156	.1603	.0491	.0623	-.0032
	(.0103)	(.0129)	(.0176)	(.0116)	(.0140)	(.0219)
d_ptime	-.0470	-.0267	-.0681	-.1171	-.1007	-.1218
	(.0331)	(.0381)	(.0934)	(.0130)	(.0154)	(.0280)
othqual	.0097	.0888	-.0795	.1890	.2182	.0353
	(.0376)	(.0538)	(.0527)	(.0420)	(.0698)	(.0458)
apprent	.1312	.1250	.1052	.1346	.1817	.0487
	(.0214)	(.0327)	(.0318)	(.0657)	(.0773)	(.1564)
cse	.1567	.2039	.1100	.0995	.0583	.1643
	(.0167)	(.0231)	(.0265)	(.0222)	(.0264)	(.0378)
commql	.1967	.3051	-.1229	.1184	.1241	.0832
	(.1415)	(.2124)	(.1704)	(.0177)	(.0212)	(.0323)
o_level	.1847	.2269	.1357	.1856	.1781	.1712
	(.0128)	(.0194)	(.0188)	(.0127)	(.0159)	(.0253)
a_level	.2733	.3301	.2067	.2598	.2275	.2871
	(.0135)	(.0191)	(.0198)	(.0163)	(.0201)	(.0326)
nursql	.2873	.3210	.0556	.3357	.3336	.3136
	(.0871)	(.0940)	(.0469)	(.0357)	(.0430)	(.0642)
highql	.3274	.3629	.2756	.2454	.2535	.1981
	(.0135)	(.0194)	(.0183)	(.0170)	(.0197)	(.0316)
teachql	.2933	.2847	.3379	.3093	.3222	.2939
	(.0533)	(.0721)	(.0781)	(.0478)	(.0596)	(.1173)
firstdg	.5304	.5669	.4602	.5142	.5020	.5029
	(.0174)	(.0241)	(.0293)	(.0218)	(.0267)	(.0481)
highdeg	.6228	.6387	.5829	.7093	.7849	.4705
	(.0351)	(.0430)	(.0634)	(.0839)	(.0947)	(.1792)
d_sic2	-.0184	-.0451	-.0202	-.0034	.0261	-.0616
	(.0214)	(.0363)	(.0275)	(.0355)	(.0567)	(.0477)
d_sic3	-.0463	-.0030	-.0976	-.0111	.0564	-.1457
	(.0187)	(.0323)	(.0227)	(.0312)	(.0478)	(.0472)
d_sic4	-.0950	-.0596	-.1312	-.1069	-.0375	-.2154
	(.0195)	(.0319)	(.0245)	(.0292)	(.0463)	(.0405)
d_sic5	-.0328	.0035	-.1013	.0035	.0717	-.3197
	(.0226)	(.0330)	(.0396)	(.0522)	(.0656)	(.1054)
d_sic6	-.1719	-.1302	-.2387	-.2033	-.1302	-.3640
	(.0198)	(.0312)	(.0273)	(.0285)	(.0460)	(.0398)
d_sic7	-.1186	-.1189	-.1424	-.0079	-.0137	-.0444
	(.0232)	(.0412)	(.0252)	(.0352)	(.0542)	(.0472)
d_sic8	.0797	.1092	.0286	.0677	.1139	-.0169
	(.0200)	(.0334)	(.0261)	(.0278)	(.0455)	(.0384)
d_sic9	-.1569	-.1067	-.2403	-.2096	-.1758	-.1707
	(.0269)	(.0387)	(.0396)	(.0307)	(.0486)	(.0470)
d_size2	-.0237	-.0576	-.0374	-.0527	-.0178	-.1479

	(.0435)	(.0541)	(.0696)	(.0379)	(.0423)	(.0858)
d_size3	.0622	.0695	.0104	.0931	.1077	.0490
	(.0123)	(.0146)	(.0233)	(.0129)	(.0152)	(.0284)
d_size4	.1324	.1693	.0072	.1344	.1659	.0393
	(.0136)	(.0173)	(.0211)	(.0136)	(.0176)	(.0264)
d_size5	.1299	.1657	.0199	.1491	.1798	.0569
	(.0125)	(.0169)	(.0207)	(.0148)	(.0197)	(.0259)
d_size6	.1348	.1751	.0479	.1825	.2258	.1063
	(.0123)	(.0159)	(.0198)	(.0133)	(.0179)	(.0254)
d_size7	.2076	.2359	.1279	.1724	.2044	.1250
	(.0164)	(.0266)	(.0229)	(.0186)	(.0284)	(.0305)
d_size8	.2613	.2973	.1771	.2067	.1798	.1692
	(.0149)	(.0224)	(.0222)	(.0187)	(.0329)	(.0293)
d_sicms	.0111	-.0835	.1452	-.0658	-.0168	-.1390
	(.0851)	(.1214)	(.0878)	(.0957)	(.1241)	(.1818)
adhimis	.2944	.3041	.1761	.1127	.1298	.0269
	(.0348)	(.0400)	(.0926)	(.0591)	(.0703)	(.0924)
d_sizems	-.0569	-.2602	.2090	-.1215	-.1969	-.0108
	(.1344)	(.1446)	(.1429)	(.1104)	(.0728)	(.2608)
wave2	.0397	.0326	.0485	.0482	.0457	.0506
	(.0120)	(.0157)	(.0183)	(.0125)	(.0158)	(.0206)
wave3	.0432	.0246	.0665	.0631	.0655	.0600
	(.0115)	(.0167)	(.0179)	(.0130)	(.0167)	(.0199)
wave4	.0589	.0356	.0989	.0790	.0914	.0569
	(.0103)	(.0141)	(.0169)	(.0113)	(.0149)	(.0186)
wave5	.0653	.0439	.1048	.1182	.1255	.1202
	(.0130)	(.0151)	(.0201)	(.0112)	(.0138)	(.0201)
wave6	.0940	.0719	.1379	.1439	.1606	.1351
	(.0121)	(.0164)	(.0186)	(.0123)	(.0157)	(.0232)
wave7	.0723	.0596	.1093	.1895	.2182	.1542
	(.0208)	(.0260)	(.0376)	(.0222)	(.0281)	(.0385)
_cons	-.3551	-.5911	.5350	.2630	.2860	.0160
	(.1865)	(.1443)	(.2338)	(.1413)	(.1846)	(.2952)
R^2	.4795	.5089	.4626	.4241	.4318	.4662
Observations	11829	7248	4581	9280	6541	2739

Note: Regressions also include 216 indicator variables for travel-to-work area (TTWA). Standard errors are calculated allowing for TTWA-by-year clustering. Dependent variable is the log hourly wage.

Table A4: Log Wage Regressions, Norway

	Males			Females		
	All	Nonunion	Union	All	Nonunion	Union
logunemp	-.0466	-.0660	-.0169	-.0159	-.0453	.0097
	(.0177)	(.0242)	(.0262)	(.0229)	(.0343)	(.0268)
union	-.0162			.0398		
	(.0035)			(.0043)		
school	.0635	.0677	.0549	.0627	.0708	.0528
	(.0007)	(.0009)	(.0009)	(.0013)	(.0022)	(.0014)
exp	.0303	.0369	.0211	.0245	.0304	.0165
	(.0008)	(.0008)	(.0007)	(.0008)	(.0011)	(.0006)
expsq	-.0005	-.0006	-.0003	-.0004	-.0005	-.0002
	(.0000)	(.0000)	(.0000)	(.0000)	(.0000)	(.0000)
senior	.0041	.0076	.0019	.0076	.0114	.0049
	(.0003)	(.0004)	(.0004)	(.0004)	(.0006)	(.0004)
ind_c	.3218	.3053	.3528	.3247	.3073	.3567
	(.0253)	(.0212)	(.0330)	(.0192)	(.0220)	(.0245)
ind_da	.0929	.0903	.1043	-.0017	-.0133	.0202
	(.0072)	(.0104)	(.0102)	(.0096)	(.0165)	(.0124)
ind_db	-.0299	-.0328	-.0214	-.0781	-.0217	-.1066
	(.0174)	(.0260)	(.0198)	(.0171)	(.0273)	(.0185)
ind_dc	-.0927	-.0865	-.1090	-.0362	.0073	-.0716
	(.0435)	(.0584)	(.0556)	(.0512)	(.0680)	(.0678)
ind_dd	-.0029	.0076	-.0072	.0019	.0233	-.0119
	(.0121)	(.0160)	(.0136)	(.0286)	(.0451)	(.0355)
ind_de	.1794	.1287	.2237	.1318	.0981	.1799
	(.0074)	(.0146)	(.0099)	(.0101)	(.0142)	(.0156)
ind_df	.3575	.3194	.3789	.3323	.2676	.3670
	(.0223)	(.0326)	(.0274)	(.0476)	(.0939)	(.0568)
ind_dg	.2254	.2504	.2396	.1640	.1389	.1996
	(.0149)	(.0204)	(.0178)	(.0162)	(.0254)	(.0210)
ind_dh	.0207	.0359	.0159	.0624	.0999	.0380
	(.0148)	(.0205)	(.0200)	(.0200)	(.0265)	(.0286)
ind_di	.0607	.0747	.0644	.0438	.0443	.0595
	(.0122)	(.0181)	(.0151)	(.0282)	(.0422)	(.0280)
ind_dj	.1354	.1299	.1515	.1334	.0775	.1676
	(.0087)	(.0123)	(.0112)	(.0180)	(.0302)	(.0212)
ind_dk	.1442	.1680	.1361	.1255	.1170	.1393
	(.0108)	(.0145)	(.0130)	(.0190)	(.0247)	(.0288)
ind_dl	.1526	.1469	.1725	.0947	.0994	.0998
	(.0112)	(.0157)	(.0126)	(.0156)	(.0239)	(.0183)
ind_dm	.1305	.1422	.1291	.1641	.1544	.1746
	(.0105)	(.0144)	(.0127)	(.0182)	(.0300)	(.0213)
ind_dn	-.0152	.0082	-.0331	.0460	.0253	.0649
	(.0123)	(.0176)	(.0151)	(.0166)	(.0273)	(.0213)
ind_e	.1222	.1161	.1437	.0870	.1213	.1041
	(.0095)	(.0234)	(.0101)	(.0162)	(.0571)	(.0152)
ind_f	.0240	.0071	.0627	.0366	.0300	.0556
	(.0066)	(.0071)	(.0108)	(.0139)	(.0153)	(.0352)
ind_h	-.2138	-.2069	-.1957	-.0477	-.0382	-.0453
	(.0160)	(.0173)	(.0287)	(.0111)	(.0119)	(.0161)
ind_i	.0570	.0515	.0838	.0955	.0956	.1185
	(.0091)	(.0088)	(.0147)	(.0081)	(.0119)	(.0135)
ind_j	.2017	.3056	.1857	.1886	.2516	.2131

	(.0078)	(.0091)	(.0146)	(.0081)	(.0161)	(.0119)
ind_k	.0992	.0883	.1368	.0903	.0750	.1208
	(.0073)	(.0074)	(.0111)	(.0094)	(.0101)	(.0147)
ind_l	.0295	.0064	.0831	.0372	-.1203	.1101
	(.0457)	(.0726)	(.0364)	(.0326)	(.0726)	(.0352)
ind_m	-.0646	-.0813	-.0011	-.0253	-.0934	.0538
	(.0107)	(.0150)	(.0130)	(.0135)	(.0220)	(.0159)
ind_n	-.0360	-.0099	-.0335	.0172	-.0153	.0595
	(.0129)	(.0184)	(.0140)	(.0099)	(.0123)	(.0122)
ind_o	-.0519	-.0822	.0196	-.0593	-.0995	.0309
	(.0113)	(.0131)	(.0129)	(.0125)	(.0150)	(.0138)
ind_x	.0180	-.0182	.2664	.0480	.0386	.1070
	(.0766)	(.0805)	(.1681)	(.0518)	(.0627)	(.0666)
year1992	.0496	.0564	.0417	.0330	.0429	.0263
	(.0054)	(.0063)	(.0082)	(.0066)	(.0091)	(.0077)
year1993	.0935	.0968	.0904	.0875	.0969	.0816
	(.0056)	(.0067)	(.0083)	(.0061)	(.0089)	(.0067)
year1994	.1286	.1308	.1306	.1212	.1297	.1157
	(.0053)	(.0058)	(.0082)	(.0064)	(.0084)	(.0074)
year1995	.1729	.1676	.1823	.1553	.1530	.1609
	(.0045)	(.0051)	(.0077)	(.0064)	(.0096)	(.0058)
year1996	.2132	.2089	.2226	.2044	.1997	.2136
	(.0049)	(.0066)	(.0074)	(.0072)	(.0119)	(.0083)
year1997	.2710	.2605	.2920	.2606	.2545	.2747
	(.0093)	(.0104)	(.0129)	(.0132)	(.0177)	(.0142)
regio_1	-.1366	-.1473	-.1199	-.1581	-.1945	-.1148
	(.0045)	(.0055)	(.0064)	(.0073)	(.0085)	(.0076)
regio_2	-.0367	-.0593	.0045	-.0613	-.0634	-.0533
	(.0103)	(.0135)	(.0148)	(.0120)	(.0174)	(.0152)
regio_4	-.1928	-.2101	-.1660	-.1998	-.2474	-.1504
	(.0054)	(.0100)	(.0112)	(.0074)	(.0101)	(.0121)
regio_5	-.1940	-.2129	-.1617	-.1907	-.2278	-.1455
	(.0068)	(.0081)	(.0103)	(.0083)	(.0128)	(.0121)
regio_6	-.1102	-.1116	-.0984	-.1236	-.1332	-.1006
	(.0061)	(.0109)	(.0113)	(.0113)	(.0149)	(.0118)
regio_7	-.1254	-.1293	-.1126	-.1653	-.1799	-.1351
	(.0075)	(.0088)	(.0100)	(.0067)	(.0100)	(.0069)
regio_8	-.1211	-.1417	-.0954	-.1673	-.2010	-.1260
	(.0066)	(.0099)	(.0099)	(.0099)	(.0160)	(.0085)
regio_9	-.1776	-.1945	-.1462	-.1717	-.1894	-.1275
	(.0068)	(.0089)	(.0156)	(.0164)	(.0215)	(.0151)
regio_10	-.1194	-.1201	-.1108	-.1993	-.2218	-.1601
	(.0066)	(.0099)	(.0107)	(.0113)	(.0087)	(.0181)
regio_11	-.0229	-.0413	.0052	-.0968	-.1114	-.0705
	(.0094)	(.0078)	(.0147)	(.0086)	(.0105)	(.0116)
regio_12	-.0865	-.0998	-.0633	-.1351	-.1577	-.0957
	(.0042)	(.0061)	(.0072)	(.0062)	(.0082)	(.0088)
regio_14	-.1864	-.2254	-.1389	-.1910	-.2413	-.1300
	(.0133)	(.0196)	(.0196)	(.0167)	(.0275)	(.0234)
regio_15	-.1530	-.1806	-.1164	-.1929	-.2332	-.1425
	(.0077)	(.0085)	(.0107)	(.0084)	(.0121)	(.0099)
regio_16	-.1514	-.1553	-.1385	-.1398	-.1402	-.1219
	(.0054)	(.0073)	(.0078)	(.0068)	(.0158)	(.0085)
regio_17	-.2119	-.2507	-.1716	-.2222	-.2726	-.1691
	(.0073)	(.0114)	(.0088)	(.0082)	(.0145)	(.0075)
regio_18	-.1643	-.1834	-.1372	-.1755	-.1790	-.1543

	(.0053)	(.0094)	(.0080)	(.0106)	(.0211)	(.0074)
regio_19	-.1797	-.1970	-.1541	-.1621	-.1786	-.1326
	(.0076)	(.0098)	(.0110)	(.0086)	(.0091)	(.0124)
regio_20	-.2078	-.2161	-.1890	-.1468	-.1222	-.1488
	(.0121)	(.0092)	(.0178)	(.0103)	(.0239)	(.0078)
_cons	4.2312	4.1796	4.2694	4.0806	4.0659	4.1453
	(.0317)	(.0413)	(.0478)	(.0396)	(.0598)	(.0451)
R^2	.3685	.3743	.3884	.3436	.3349	.3600
Observations	77537	39300	38237	34126	17138	16988

Note: Standard errors are calculated allowing for region-by-year clustering. Dependent variable is the log hourly wage.

Figure 1. The separation and wage curve elasticities.