Wage Rigidity, Inflation, and Institutions*

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Abstract

We study whether downward nominal wage rigidity (DNWR) for job stayers is transformed to DNWR at rates different from zero in aggregate data. Compositional effects may lead to falling aggregate wages, while changes in relative wages combined with DNWR may lead to positive aggregate wage growth. We explore industry data for 19 OECD countries, over the period 1971–2006. We find evidence for floors on nominal wage growth at 6 percent in the 1970s and 1980s, at one percent in the 1990s, and at 0.5 percent in the 2000s. Furthermore, we find that DNWR is stronger in country-years with strict employment protection legislation, high union density, centralized wage setting and high inflation.

JEL: J3, J5, C14, C15, E31
Keywords: Wage inflation, downward nominal wage rigidity, OECD, wage setting

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A number of recent studies have documented that nominal wages for job stayers are rigid downwards in many OECD countries, see e.g. Dickens et al. (2007) and Knoppik and Beissinger (2008). Akerlof, Dickens, and Perry (1996) and Bewley (1999) report strong evidence of downward nominal wage rigidity (DNWR) based on surveys of employees and employers. However, the aggregate effects of DNWR are less clear. The recent downturn with soaring unemployment and low inflation in many countries makes the existence and implications of DNWR of particular interest.

In this paper we explore the effects of DNWR on aggregate wage growth. In particular, we study whether the rigidity found in micro data may give rise to DNWR at rates different from zero in aggregate data. This could be caused by compositional changes in the labor market, e.g. when old high-wage workers are replaced by younger workers with lower wages, or if jobs are shifted over from high-wage firms with rigid wages to other firms with lower wages. There could also be mechanisms working in the opposite direction. Tobin (1972) emphasized that sector specific shocks would induce growth in aggregate wages if DNWR prevents wage reductions in labor markets with excess supply. Another mechanism, which would affect positive wage growth even at the micro level, is if workers with secure jobs and nominal wage contracts use threats of reduced work effort to enforce a rise in nominal wages.

The existence of DNWR, and the rate at which it binds, is of great importance for monetary policy. Tobin (1972) argues that if inflation is so low that DNWR binds, the result will be excess wage pressure and higher unemployment. If there is a floor for wage growth above zero, the recent downturn involves a risk of pervasive binding DNWR, which exacerbates the rise in unemployment. Introducing DNWR in a DSGE model, Carlsson and Westermark (2008) also show that DNWR has important implications for the cyclical response of the economy, and thus also for the optimal monetary policy.

To explore the prevalence of DNWR, we use industry data for 19 OECD countries, over the period 1973–2006, consisting of more than 13,000 observations from 604 country-year samples, i.e. an average of 23 industries per sample. Following Holden and Wulfsberg (2008) we construct an assumed distribution of wage changes under flexibility (i.e. without DNWR, referred to as the "notional" wage change distribution), on the basis of observations from country-years when the wage growth is high, and thus DNWR is not likely to bind.

By comparing empirical and notional country-year specific wage change distributions at different levels of wage growth, we derive country-year specific estimates of the extent of DNWR, given by the deficit of wage changes below varying levels of wage growth in the empirical samples.

Compared to the existing literature on DNWR, we make three main contributions. First, we present a theoretical model of DNWR, and point out novel implications on how DNWR affects the wage change distribution. We show that the fraction of notional wage cuts that is prevented by DNWR is likely to be decreasing in the rate of inflation and increasing in the rate of unemployment. However, even if the wage is cut, the resulting wage will be higher than it would have been with entirely flexible wage setting. Second, we explore empirically the extent of DNWR at non-zero growth rates, which has received scant attention in spite of the clear policy relevance. As we obtain country-year specific estimates of DNWR for some 600 country-year samples, we are in a good position to explore how DNWR is related to economic and institutional variables like unemployment, inflation, employment protection, and union density. Third, we include data from the years 2000–2006, providing new evidence from the Great Moderation period. More than ten years ago, Gordon (1996) and Mankiw (1996) argued that the extent of DNWR reflected a recent history of high inflation. Referring to the experience from the Great Depression, they claimed that "nominal wage reductions would no longer be seen as unusual if the average nominal wage was not growing" (Gordon, 1996, p.62). In most OECD countries we have now experienced more than a decade of fairly low and stable inflation, and it is important to explore whether wages are still rigid downwards.

Our study does not explore whether there is any effect of DNWR on output and employment. Yet we argue that our study is of interest also from this angle. If there is no sign of DNWR in industry-level wage data, one may also question whether the DNWR found in micro data has any noticeable effect on industry output or employment. On the other hand, if there is DNWR in industry-level wage data, rigidity prevails in spite of varying compositional effects. In this case effects on industry output and employment also seem more likely. To explore the existence of such effects would then seem an important issue for future research.

To preview our results, we find evidence of floors on aggregate wage growth as high as 5–6 percent in 1970s. According to our point estimates, around 40 percent of the notional industry wage changes in the 1970s below 5 percent were pushed up above 5 percent. In the 1980s, the floors were at somewhat lower levels, yet 20 percent of all wage changes below 4 percent were pushed up above 4 percent. In the 1990s, the floors fell down to about zero, with a fraction of notional wage cuts prevented by DNWR of about 20 percent. In the 2000s, we find evidence of a floor in the Nordic countries at 0–1 percent wage growth, and also some evidence for Southern European countries as well as for the OECD as a whole. We find that DNWR is strongly correlated with inflation and labor market institutions. The evidence for DNWR in the Southern European countries is important in light of their troubled economies and weak competitive positions (see e.g. European Commision, 2010), as these countries are part of the Euro area and thus unable to devalue to restore competitiveness.

The paper is organized as follows. In section 1, we provide a simple bargaining model predicting wage floors which may differ from zero percent, and discuss the link between DNWR and aggregate wage growth. Section 2 presents the data and empirical approach. Section 3 contains the main results as to the extent of DNWR, while in section 4, we explore whether the variation in DNWR across countries and time can be explained by institutional and economic variables. Section 5 concludes. The appendices contain supplementary material on data and results.

1 DNWR and floors for aggregate wage growth

In this section we discuss the theoretical implications of DNWR that we shall explore in the empirical analysis. While DNWR is often motivated from fairness considerations (see e.g. Akerlof, Dickens, and Perry (1996)), we consider a bargaining model. This choice reflects that in most OECD countries, collective agreements cover the large majority of the workforce.

We consider a simple model of DNWR under firm-level bargaining, drawing upon Holden (2004). Let $\pi(W/P) = (W/P)^{1-\eta}$ denote the real flow revenue per worker to the firm where the real wage, W/P, denotes the flow payoff to the workers, for simplicity neglecting any disutility of labor.[1] P is the aggregate price level, which is exogenous to the parties in the bairgaining, while η is the elasticity of demand, where $\eta > 2$. Consistent with the institutional setting in most OECD countries, we assume that there is a prevailing wage contract, W_0, which is given in nominal terms. The terms of the contract can only be changed by mutual consent, even if the contract has a fixed duration (see MacLeod and Malcomson, 1993, and Holden, 1994). This implies that the terms of the old contract prevail also after its expiration, until the players have agreed to a new contract, or one of the players unilaterally stop production.

A crucial feature of the model is that the type of dispute in the bargaining is endogenous, see Haller and Holden (1990) and Cramton and Tracy (1992). The workers may initiate a strike, or the firm a lock-out. In both cases the firm obtains zero payoff, while the workers get R/P, the alternative real income during a work stoppage. If neither of the players stop production, it continues under the prevailing wage contract, W_0, while the parties are bargaining ("holdout"). Following Cramton and Tracy (1992), Moene (1988), and Holden (1997) we assume that the workers may inflict a cost on the firm during a holdout, e.g. by strictly adhering to the rules of the employment contract (work-to-rule). Correspondingly, we allow for the possibility that the firm imposes a cost on workers, e.g. by reducing bonus schemes or by other measures. Formally, this can be captured by assuming that the payoffs during a holdout are $(1-\tau)(W_0/P)^{1-\eta}$ to the firm and $(1-\varepsilon)W_0/P$ to the workers, where τ and ε are parameters satisfying $0 \leq \tau < 1,\ 0 \leq \varepsilon < 1$. A holdout is costly to both players if $\tau > 0$ and $\varepsilon > 0$.

A second important feature of the model is the assumption that if a work stoppage takes place, it always involves non-negligible fixed costs to the parties. These fixed costs can be motivated as the costs associated with a minimum time before work can be resumed after a work stoppage. Alternatively, if the model is extended to allow for uncertainty as to the bargaining outcome and risk aversion, the fixed cost would be the expected payoff players are willing to give up to avoid risk (Holden, 1999). The exact way these costs enter does not affect the qualitative results, which is the basis for the subsequent empirical analysis. We assume that when production is resumed after a work

[1] This profit function follows from a model of monopolistic competition, in which firms set the output price facing a downward sloping demand curve, and there is constant return with labor the only factor of production (irrelevant constants omitted).

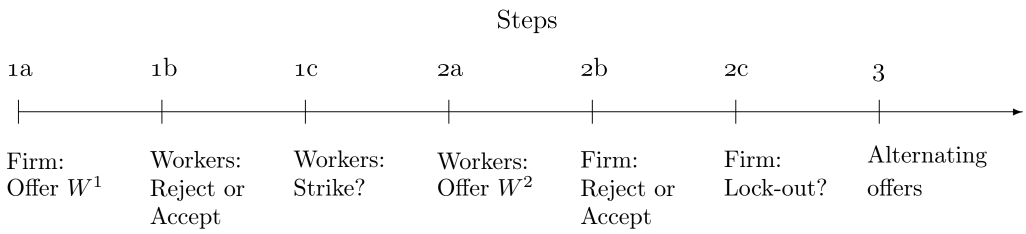

Figure 1: The wage bargaining game.

stoppage, the payoffs of the players are $\gamma\pi(W/P)^{1-\eta}$ and $\gamma W/P$, where $\gamma < 1$, so that $1-\gamma$ is the reduction in payoff caused by a work stoppage.

To determine which type of threats that prevail, we use the extension of the Rubinstein (1982) alternating offers bargaining game employed by Holden (2004). The game is illustrated in Figure 1. The first two steps of the game, which take place in negligible time, determine which type of dispute (strike, lockout or holdout) prevails in the bargaining. At the third step, a standard Rubinstein bargaining games starts, whereby players alternate in making offers, one offer per time span. In each of the first two steps, one of the players makes an offer, which the opponent may accept (in which case the bargaining ends) or reject. Upon rejection, the rejecting player may decide whether to initiate a work stoppage. A player is assumed only to initiate a work stoppage if this gives him/her strictly higher payoff than the alternative. If no work stoppage is initiated, there will be a holdout from step 3 on.

In equilibrium an agreement is reached in step 1 or 2, with no costly dispute (cf. Appendix A):

Proposition: *The unique* SPE *outcome to the wage bargaining is W given by*

$$W = \begin{cases} \ell R, & \text{if } (1+\kappa)W_0 > \ell R \qquad (\textit{Lockout}) \\ (1+\kappa)W_0, & \text{if } (1+\kappa)W_0 \in [sR, \ell R] \qquad (\textit{Holdout}) \\ sR, & \text{if } (1+\kappa)W_0 < sR \qquad (\textit{Strike}) \end{cases} \tag{1}$$

where $\ell = \left(\frac{\eta-1}{\eta-2}\right)\gamma^{\frac{1}{1-\eta}} > s = \left(\frac{\eta-1}{\eta-2}\right)\gamma > 1$.[2] κ *is defined implicitly by* $(1-\tau)(1+\kappa)^\eta + (\eta-2)(1+\kappa) - (\eta-1)(1-\varepsilon) = 0$. *Furthermore,* κ *is strictly increasing in* τ*, strictly decreasing in* ε *and* η*, and* $\kappa > 0$ *if and only if* $\varepsilon < \tau/(\eta-1)$.

[2]This assumes that $\gamma(\eta-1)/(\eta-2) > 1$. If this condition is not fulfilled, then the workers' outside alternative binds, implying that $s = 1$ and $W = R$. The remaining analysis is however unchanged.

The intuition for the proposition is as follows. If holdout threats prevail in the bargaining, the outcome depends on the old contract wage because this determines the players' payoffs during the bargaining. If $\varepsilon < \tau/(\eta - 1)$, which roughly means that a holdout is more costly to the firm than to the workers, a holdout will lead to an increase in nominal wages. Alternatively, if the holdout is more costly to the workers than the firm ($\varepsilon < \tau/(\eta - 1)$), the nominal wage will decrease when holdout threats prevail.

However, the possibility of strike or lock-out implies that holdout threats will not always prevail. If the workers initiate a strike, they receive a nominal payoff sR. Thus, if a holdout gives them a lower payoff, i.e. $(1 + \kappa)W_0 < sR$, the workers will threaten to strike. As strike threats are credible in this case, the firm must in step 1 offer a wage that is sufficiently high that the workers do not strike. The firm then offers a wage $W = sR$. This gives the workers the same payoff as they could obtain by initiating a strike, hence they accept. Alternatively, if the holdout outcome were higher, $(1+\kappa)W_0 > \ell R$, it would be the firm who would use lock-out threats to push wages down. The workers would then offer $W = \ell R$ in step 2, and the firm would accept this. Finally, if $(1 + \kappa)W_0 \in [sR, \ell R]$, threats of initiating a work stoppage are not credible, because initiating a work stoppage gives lower (or the same) payoff to the initiating player than s/he would have obtained under holdout threats. Hence in this case holdout threats prevail in equilibrium.

The solid line in Figure 2 illustrates the bargaining outcome. We see that lock-out threats prevail for low R, holdout threats for intermediate R, and strike threats for high R. Note that the standard wage bargaining model, see e.g. Layard, Nickell, and Jackman (1991), is a special case of the present model. If there is no fixed costs of a work stoppage, $\gamma = 0$, strike and lockout threats would result in the same payoffs, and there would always be one player who could benefit from threatening to stop work. The holdout interval would collapse, implying that $s = \ell$, and the bargaining outcome would be $W = sR$. The wage setting would be entirely static, and the relative wage growth would be $\Delta w = \Delta r$ (lower case letter denote logs). However, with fixed costs of a work stoppage, $\gamma > 0$, the old contract may affect the wage outcome, and wages may be rigid downwards.[3]

[3]Holden (1997) considers a similar model with the extension that players take the possible effect on future wage negotiations into consideration during the bargaining. It is shown that the qualitative properties of the bargaining outcome are unchanged, even if the magnitudes are dampened.

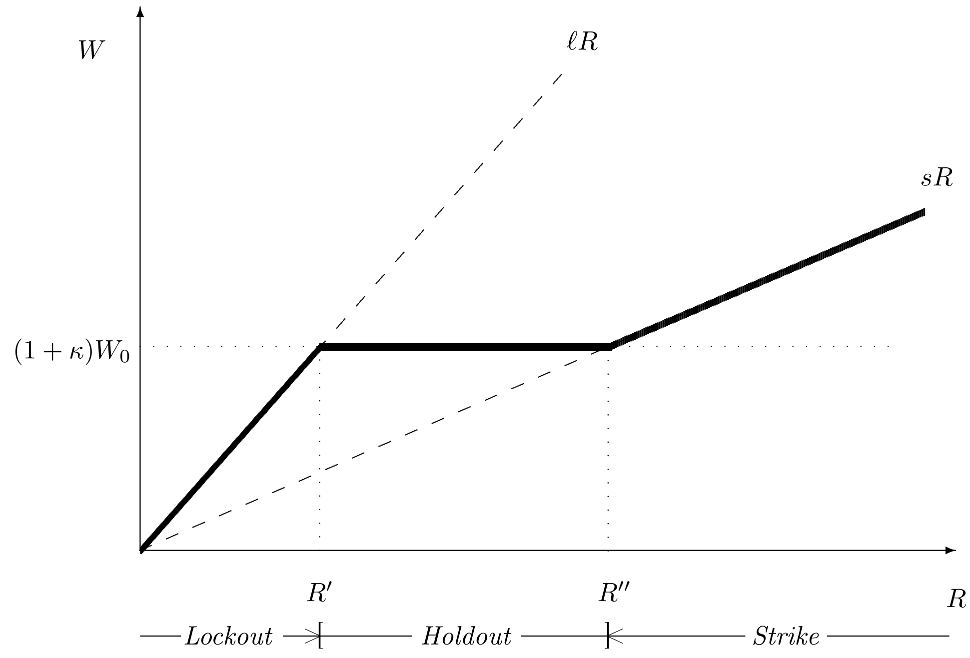

Figure 2: The wage bargaining outcome (solid line) as a function of the alternative income R. For $R < R' = (1+\kappa)W_0/\ell$, lockout threats are credible and the firm pushes wages down to ℓR by use of lock-out threats. For $R > R'' = (1+\kappa)W_0/s$, strike threats are credible and the union pushes wages up to sR by use of strike threats. For $R \in [R', R'']$, neither strike nor lock-out threats are credible, so that hold-out threats prevail.

A second aspect to note is that rigidity is in fact symmetric in the model, in the sense that the fixed costs of a work stoppage may prevent wage rises as well as wage cuts. However, due to inflation and productivity growth, the alternative income R will increase over time, implying that it is the downward rigidity that will be most noticeable.

A third aspect is that in discussions and analysis of DNWR, it is usually assumed that if the wage is cut, then wage rigidity has no impact on the actual wage outcome (see e.g. Knoppik and Beissinger, 2003 and Fehr and Gotte, 2005). In the present model, however, if DNWR exists, then it pushes the wage upwards even when the wage is cut. To see this, note that in Figure 2, the line indicating the wage outcome when lock-out threats prevail, ℓR, lies above the wage outcome when strike threats prevail, sR. This implies that in deflationary settings, where the firm uses lock-out threats to push wages down, real wages will nevertheless be higher than they would have been if strike threats had prevailed in the wage setting.

The parameter values of the model, including ε, τ, and γ, may differ across industries, countries and time, and between union and non-union settings, reflecting differences in legal and institutional settings, production technology, monetary policy and remuneration systems. For example, strict employment protection legislation and high union density is likely to strengthen the workers' side in the negotiations, consistent with the analysis in Holden (2004). In the short run, these parameters can be taken as given. However, in the longer run, these factors are clearly endogenous. For example, with lower inflation, firms will benefit from imposing a remuneration system with more extensive use of bonus schemes and less fixed pay, so as to make wages more flexible. This will increase firms' scope for reducing pay during a holdout, i.e. increase ε. This will lower κ, implying a lower nominal wage change during a holdout.

The implications of the bargaining model for the distribution of wage changes, which is the object of our empirical investigation, can be illustrated by a simple numerical example. Consider a hypothetical distribution of wage changes from the bargaining model above, where Δr is drawn from a normal distribution with mean 0.030 and standard deviation 0.025, and parameter values $\eta = 3$ and $\kappa = 0$. In the notional (rigidity free) distribution derived with $\gamma = 1$, about 12 percent of the observations are below zero, see the dashed line in the left panel of Figure 3. Adding DNWR by setting $\gamma = 0.995$, indicated by the solid line, leads to a deficit of wage changes below zero. The right panel measures the fraction of wage changes prevented (FWCP) by DNWR, which is the missing probability mass relative to the probability mass of the notional distribution, as a function of the level of wage growth. Intuitively, the FWCP is the missing number of wage cuts divided by the number of wage cuts that should have taken place without rigidity; a precise definition is given in equation (6) below.

We observe that the FWCP is downward sloping in a wage change diagram: For example, below −4 percent wage growth, DNWR leads to 60 percent fewer wage changes than the notional distribution, while below zero percent there are 20 percent fewer observations in the distribution affected by DNWR. In other words, the density deficit caused by DNWR is greater for large negative wage changes than for small. The intuition here is that while DNWR prevents some small wage cuts, it also attenuates larger notional wage cuts, as the wage outcome follows the upper line rather than the lower, see Figure 2. This reduces the deficits of smaller cuts.

Figure 4 illustrates the theoretical predictions on how the FWCP curve depends on

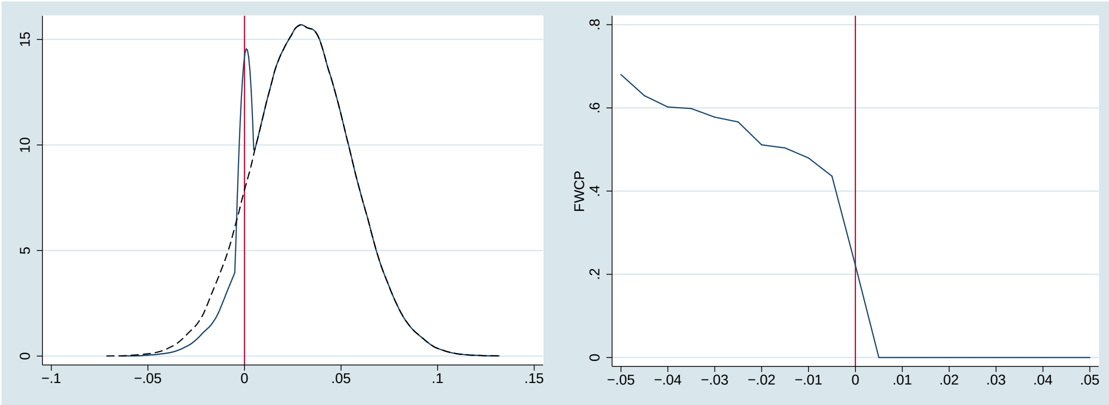

Figure 3: Left: Density of a notional distribution of nominal wage changes (dotted line) and a distribution of nominal wage changes subject to a wage floor at 0 percent (solid line). Right: The FWCP *curve at different levels of wage growth derived from the distributions in left panel. Parameter values are* $\eta = 3,\ \gamma = .995,\ \kappa = 0.$

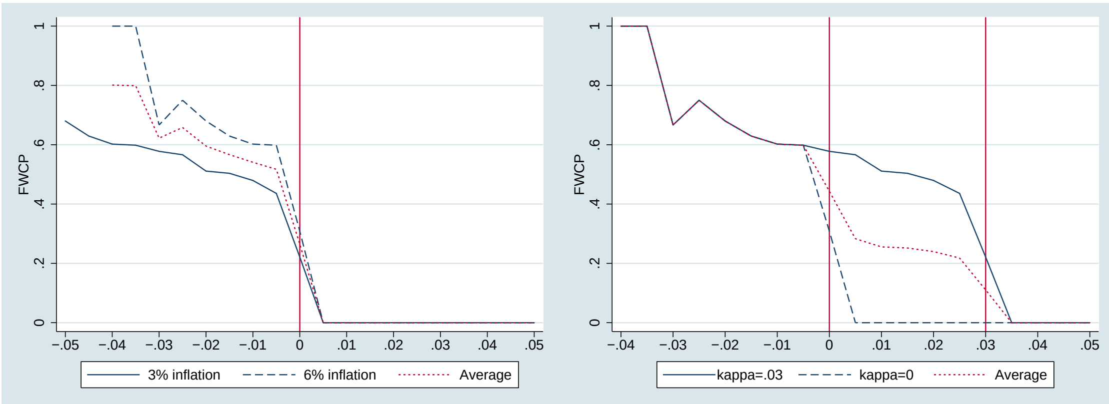

Figure 4: The FWCP*. Left: The* FWCP *curves with a wage floor at zero percent when mean wage growth is 3 percent (solid line) and 6 percent (dashed line). Right: The* FWCP *curves with wage floors at zero percent (dashed line) and three percent (solid line) when mean wage growth is 6 percent.*

the average wage change and the wage floor. First, the left panel of Figure 4 shows that a higher mean wage growth leads to a higher FWCP, because when the wage change distribution is moved to the right, a larger part of the notional wage cuts are pushed above the floor with no wage cut being realized. This implies that any factor that leads to higher notional wage growth, like higher productivity growth, lower unemployment or higher inflation, also leads to a lower FWCP. The right panel illustrates the effect of the level of the wage floor on the FWCP, for a given notional distribution of wage growth. Both panels of Figure 4 also illustrate how the aggregate FWCP may look (dotted line) when averaged over two different situations. The right panel shows aggregation over different wage floors (different κs), where half the observations have a floor at zero, the other half a floor at 3 percent. In the empirical analysis we estimate FWCP-curves comparable to those displayed in Figure 4, and investigate how these are affected by inflation, unemployment, and institutional factors.

2 Data and empirical approach

We use an unbalanced panel of industry level data for the annual percentage growth of gross hourly earnings for manual workers from the manufacturing, mining and quarrying, electricity, gas and water supply, and construction sectors of 19 OECD countries in the period 1973–2006. The countries included in the sample are Austria, Belgium, Canada, Germany, Denmark, Spain, Finland, France, Greece, Ireland, Italy, Luxembourg, Netherlands, Norway, New Zealand, Portugal, Sweden, the UK and the United States. The main data source are wages in manufacturing from the ILO and harmonized hourly earnings in manufacturing from Eurostat, cf. appendix B. One observation is denoted Δw_{jit} where j is index for industry, i is index for country and t is index for year. There are all together 13,694 observations distributed across 604 country-year samples, on average 23 industries per country-year.

As the observational unit is the change in the average hourly earnings in an industry, it will be affected both by the average wage change for the existing workforce, and by compositional effects due to differences in wages between new hires and the workers that leave the industry. Compositional changes and turnover may smoothen out the effect of DNWR at the individual level. Likewise, if firms respond to rigid base wages by reducing bonus schemes and fringe benefits, consistent with the evidence in Babecký et al. (2009), the effect of DNWR may vanish. However, it is also possible that these mechanisms only transform DNWR at zero (or some other rate) for individual workers to a floor for nominal wage growth at more aggregate levels that is different from zero, implying a "missing probability mass" below a rate different from zero. This is the empirical question we shall explore.

A deficit of wage changes below certain floors might also be caused by other mechanisms than DNWR. For example, if there are systematic cyclical compositional changes in the workforce, so that the share of low-skilled workers decreases in recessions, this will dampen the downward pressure on wages in the recession (Solon, Barsky, and Parker, 1994). However, these mechanisms would be in real terms, leading to a deficit of real wage changes, and not to a floor for nominal wage growth as we are looking for.

The common approach in the DNWR literature (see e.g. Card and Hyslop, 1997, Knoppik and Beissinger, 2003, Lebow, Saks, and Wilson, 2003, Nickell and Quintini, 2003 and Holden and Wulfsberg, 2008), is to construct a distribution of "notional" or rigidity-free wage changes and then explore whether the empirical wage change distribution is compressed from the left as compared to the notional distribution (cf. illustration in in Figure 3). From the theoretical model in section 1, the notional wage change in percent is simply equal to Δr. We would expect the notional wage change distribution to vary over time. The location of the distribution is likely to depend on variables like inflation, productivity growth and unemployment, while the dispersion is affected by among other things the size and dispersion of industry specific shocks in that country-yea, and possibly also on inflation, productivity growth and unemployment.

Specifically, we follow Holden and Wulfsberg (2008) and assume that absent any DNWR, the notional nominal wage growth in industry j in country i in year t, Δw_{jit}^N, is stochastic with an unknown distribution G, which is parameterized by the median nominal wage growth, μ_{it}^N, and the dispersion, σ_{it}^N; $G(\mu_{it}^N, \sigma_{it}^N)$. Thus, we allow the location and dispersion of the notional wage growth to vary across countries and years, to capture the large variation that exists across countries and across time with respect to monetary policy, wage setting, and industry structure. Note that as we control for the location of the country-year distribution, we also capture that the notional wage change distribution is likely to be persistent over time. However, we impose the same structural form (or shape) of G in all country-years. The theory model does not imply any specific distributional form so we follow the common approach and derive the shape of the notional distribution from high inflation years, where DNWR is assumed not to bind. Concretely, the structural form of G is constructed on the basis of a subset of 1,605 observations from 66 high wage growth country-year samples, selected on the basis that both the median nominal and the median real wage growth in the country year are in their respective upper quartiles over all country-years.[4]

[4]Using the country-year samples with median wage growth in the upper quartiles is clearly arbi-

Imposing the same structural form across countries and years is a strong assumption. Yet almost all studies in the literature use an assumption like this, because it would otherwise not be possible to test for the existence of DNWR. However, as the assumption at the very best can hold as a crude approximation, it is crucial to undertake extensive robustness checks. Below, we report several alternative specifications of G which document that our findings are robust.

The first step in the test is to construct an underlying distribution of wage changes, G, where the 1,605 empirical observations from the high wage growth samples are normalized with respect to the observed country-year specific median, μ_{it}, and inter percentile range, $(P75_{it} - P35_{it})$,

$$x_s \equiv \left(\frac{\Delta w_{jit} - \mu_{it}}{P75_{it} - P35_{it}} \right), \qquad s = 1, \ldots, 1605 \tag{2}$$

where subscript s runs over all j, i and t in the 66 country-year samples. We use the inter percentile range between the 75th and the 35th percentiles as our measure of dispersion, following Nickell and Quintini (2003), to avoid that DNWR has an effect on the measure of dispersion. The calculated x_s should thus be thought of as observations from the standardized underlying two-parametric distribution $X \sim G(0, 1)$.

Second, for each of the 604 country-years in the full sample, we compute the country-year specific distribution of notional wage changes by adjusting the underlying wage change distribution for the country-specific observed median and inter percentile range

$$Z_{it} \equiv X\Big(P75_{it} - P35_{it}\Big) + \mu_{it}, \qquad \forall\ i, t. \tag{3}$$

Thus, we have constructed 604 notional country-year distributions $Z_{it} \sim G(\mu_{it}, P75_{it} - P35_{it})$, each defined by $S = 1605$ wage-change observations $z_s^{it} \equiv x_s\Big(P75_{it} - P35_{it}\Big) + \mu_{it}$. In effect, we have constructed a two-parametric distribution G, where the two parameters of the distribution, $(\mu_{it}^N, \sigma_{it}^N)$, take the value of their empirical median and inter-percentile range, $(\mu_{it}, P75_{it} - P35_{it})$, while the shape or structural form is the same across all country-years, based on the wage changes in the 66 country-year samples with high wage growth. The left panel of Figure 5 displays the underlying distribution of wage

trary. However, as shown in Holden and Wulfsberg (2008), the results are robust to variations in this assumption.

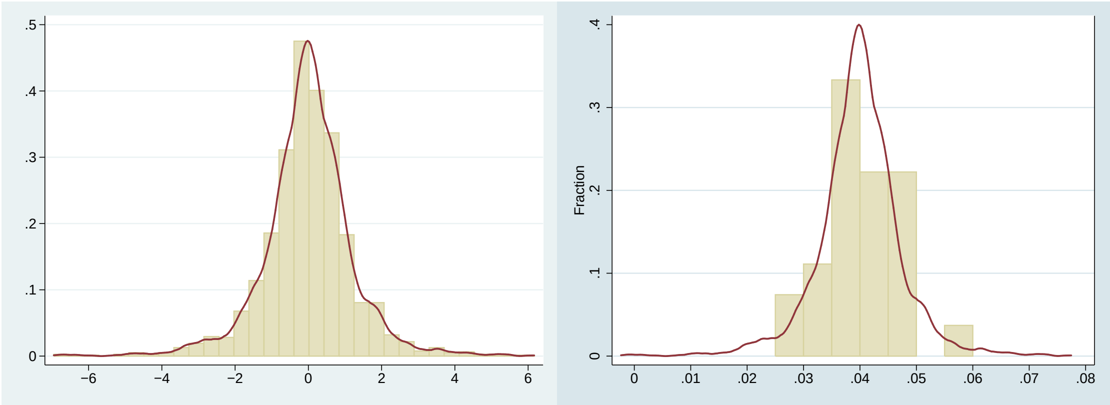

Figure 5: Left: Histogram and kernel density of the normalized underlying distribution of wage changes. Right: Histogram of observed wage changes and the notional wage change distribution in Germany, 1987.

changes, with zero median and inter-percentile range of unity. Compared with the normal distribution (not specified in the diagram), our underlying distribution has greater peak and fatter tails. Furthermore, it is skewed with the mean at −2.3 percent.

The right panel of Figure 5 compares the empirical distribution (histogram) for Germany in 1987 with the corresponding notional country-year distribution. By construction the two distributions have identical median and inter percentile range, but the shapes differ, as the notional distribution is based on the shape of the underlying distribution.

As we are looking for the existence of possible positive or negative floors for the wage growth, we look for a deficit of wage changes below floors in the range from -5 to 7 percent. For each floor $\kappa \in \{-5, -4.5, -4, \ldots, 7\}$ percent, we estimate the extent of DNWR by comparing the incidence rate of notional wage changes below the floor with the corresponding empirical incidence rate. For each floor, κ, the incidence rate of notional wage cuts is given by

$$\tilde{q}(\kappa)_{it} \equiv \frac{\# z_s^{it} < \kappa}{S}, \tag{4}$$

and is illustrated by the shaded area at the 3 percent floor in the left panel of Figure 6 for Germany 1987. Likewise, the empirical incidence rate is

$$q(\kappa)_{it} \equiv \frac{\# \Delta w^{it} < \kappa}{S_{it}}, \tag{5}$$

where S_{it} is the number of observed industries in country-year it. The empirical incidence

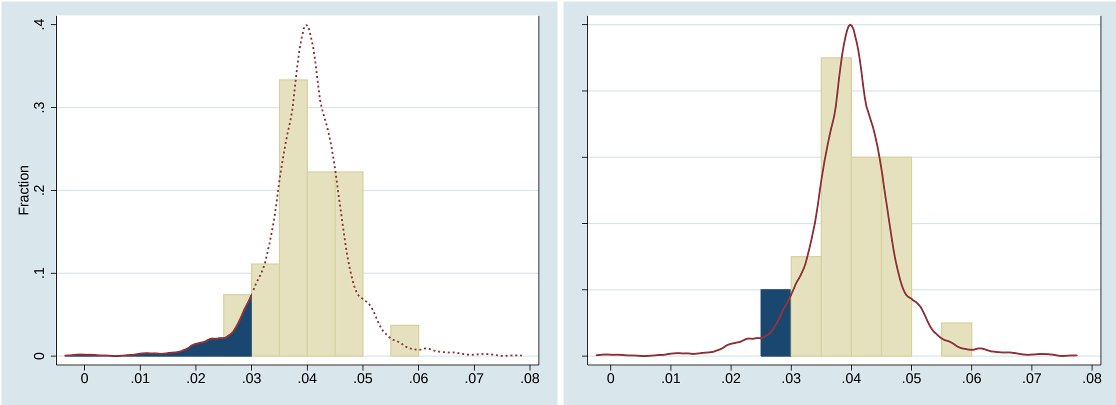

Figure 6: Illustration of the notional incidence rate below 3 percent (left) and the empirical incidence rate below 3 percent (right).

rate at 3 percent for Germany 1987 is illustrated by the dark bin in the right panel of Figure 6. The deficit of observed wage changes below floor κ relative to the notional distribution, i.e. the fraction of wage changes prevented, FWCP(κ), is calculated as

$$\text{FWCP}(\kappa)_{it} = 1 - \frac{q(\kappa)_{it}}{\tilde{q}(\kappa)_{it}}. \tag{6}$$

For Germany 1987, the empirical incidence rate below 3 percent is 0.074 and the notional incidence rate is 0.071, yielding a negative FWCP (i.e. no DNWR) at 3 percent.

As there are only on average 23 industries in each country-year sample, there may be stochastic disturbances to the country-year specific variables $\mu_{it}, P75_{it} - P35_{it}$, and q_{it}, which may induce considerable noise in $\tilde{q}_{it}$ and FWCP$_{it}$. Thus, estimates of FWCP in single country-years will be imprecise. However, averages of FWCP$_{it}$ for groups of country-years will be much more precise. Thus, we will present estimates of the average FWCP for regions and periods, similar to the average FWCP curves in Figure 4.

Under the null hypothesis of no DNWR, the notional country-year specific incidence rate $\tilde{q}(\kappa)_{it}$ is a measure of the probability that a wage change observation in that country-year is below the floor κ. Thus, in principle we could compute the probability distribution for the number of wage cuts in each country-year under the null hypothesis of no DNWR directly by employing the formulae for binomial distributions. However, this is computationally infeasible for the full sample. Hence we use the simulation method of Holden and Wulfsberg (2008). Specifically, for each country-year it, we draw S_{it} times from a binomial distribution with the country-specific notional probability $\tilde{q}(\kappa)_{it}$. We then count the number of simulated notional wage changes below κ, $\hat{Y}(\kappa)$, and compare these with the total number of observed wage changes below κ in the corresponding empirical distribution, $Y(\kappa)$. For example, to test for a floor of 1 percent in Germany, we compare the simulated number of notional wage changes below 1 percent for all years in Germany, with the corresponding empirical number. We then repeat this procedure 5000 times, and count the number of times where we simulate more notional wage changes below κ than we observe for each floor, denoted $\#(\hat{Y}(\kappa) > Y(\kappa))$. The null hypothesis is rejected with a level of significance at 5 percent if we in less than 5 percent of the simulations simulate more wage changes below the floor than the corresponding empirical number, i.e. if $1 - \#(\hat{Y}(\kappa) > Y(\kappa))/5000 \leq 0.05$.

The method is extremely powerful in detecting a possible difference between the empirical and notional wage change distribution. However, we cannot avoid that the construction of the notional wage change distribution involves a potentially substantial downward bias in the estimated DNWR. In a country-year with so extensive DNWR that the 35th percentile is pushed upwards, the notional distribution, which is constructed using the 35th percentile, will be compressed from below. This will reduce the estimated notional probability of observing a wage change below the floor, inducing a downward bias in the estimated DNWR. To mitigate this downward bias we exclude country-year samples where the distribution of wage changes is so far to the left that the 35th percentile would have been affected by the floor that we look for. Specifically, when we look for the existence of a floor at, say, 4 percent wage growth, we include only country-year samples in which the 35th percentile is above 4 percent. This procedure implies that our estimates of the FWCP are conditional on the wage distribution being sufficiently far to the right relative to the floor. Table B2 in the appendix shows the 35th percentile wage growth in each country-year sample.

This approach removes some of the downward bias in our estimates, but not all, as it may still be the case that DNWR at higher growth rates pushes up the 35 percentile. For example, if there is a floor at four percent wage growth, this would affect the notional wage change distribution in country-years where the 35 percentile is at or below 4 percent. As these country-year samples are included when we look for floors below 4 percent, say zero percent, there will be a downward bias in our estimate of the extent

DNWR at zero. However, removing even more country-years would also reduce the number of observations, which would reduce the precision of the estimates. We prefer to have a known bias, working against finding DNWR, rather than reducing the number of observations leading to more imprecise estimates.

Our method does not explicitly take into account the possible dynamic effect of DNWR in subsequent years. If binding DNWR in one years leads to wage changes below a wage floor in the following year, it will reduce the FWCP in this year. If it leads to lower wage growth, but still above the floors, it will not be captured by our method. Thus, we only detect DNWR when it happens, without exploring the possible persistence.

Note also that the downward bias in our method is likely to be more severe in country-years with low inflation, as DNWR may affect individuals in *all* industries when inflation is low, also in industries with high average wage increase. In that case DNWR may push the whole wage change distribution to the right. We would not detect this with our method, as it is based on the shape of the empirical distribution, not the location. We shall return to this issue in the interpretation of our results.

3 Estimates of DNWR

In Figure 7 we present estimates of the FWCP at floor levels of every half percentage point between –5 and 7 percent for each decade in the sample. A significant estimate at the 5 percent level is marked by a "×" in the figure. Some of these estimates in the tails of the distributions relate to very few observations. Our discussion below relates to the estimates where the number of notional wage changes is at least one percent of the relevant population, which are visualized in the figure by connecting lines. Figure 7 shows extensive DNWR at a wide range of wage growth rates for the first three decades. Observe that we find downward sloping FWCP-curves, as predicted by the bargaining model in section 1 and illustrated in Figures 3 and 4. Intuitively, some of the wage changes far below the floor are pushed up below but closer to the floor, reducing the deficit of wage changes just below the floor. We also observe a leftward shift in the FWCP-curve over time, implying that the DNWR applies to lower rates of wage growth for each decade.

For the 1970s, we find that DNWR affected wage changes in the range between 1 and

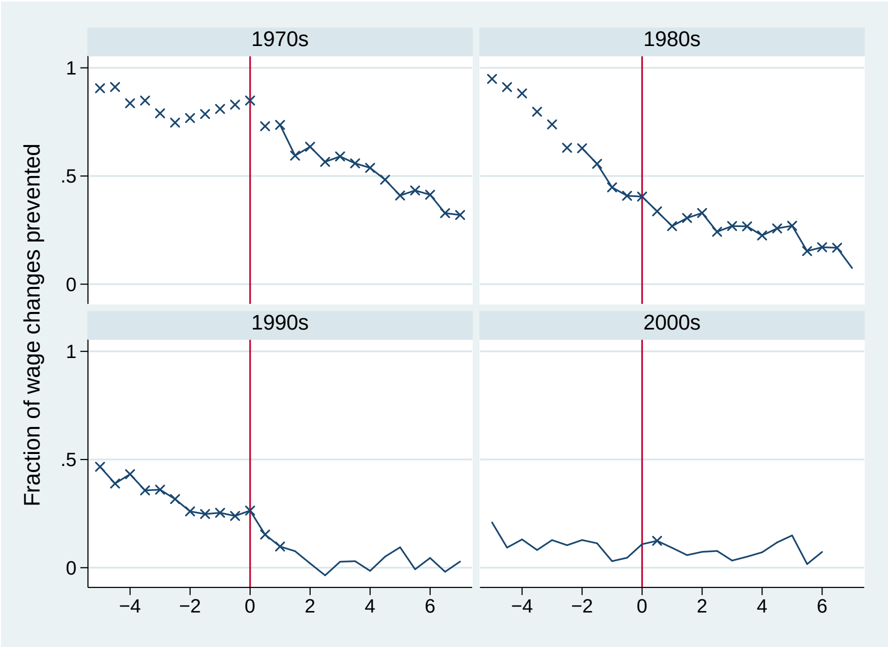

Figure 7: Fraction of wage changes prevented at different floor levels by periods. Significant estimates at the 5 percent level are marked by a "×". Estimates where the number of notional wage changes are at least one percent of the relevant sample are connected with lines.

7 percent wage growth. The FWCP at 1 percent is 0.74, implying that three out of four notional wage changes below 1 percent are pushed up above 1 percent. Even at the 7 percent floor the FWCP is 0.32. In the 1980s, the estimates indicate that DNWR has been binding at least between −2 and 6 percent, with a FWCP of 0.63 and 0.17 at the two bounds. Similarly, we see that in the 1990s, DNWR binds between −5 and 1 percent, with a FWCP between 0.47 and 0.10. In the 2000s we only find statistically significant DNWR at the 0.5 percent floor, with a FWCP equal to 0.12.

While the left/down movement of the FWCP's indicates less DNWR, there is also an opposing effect which increases the economic importance of DNWR. As inflation falls over time, the wage change distribution moves to the left, and more wage change observations are potentially affected by DNWR at any given floor. However, the former effect dominates, so that DNWR affects fewer industry-year wage change observations in the last part of the sample. In the 1970s and 1980s, around 1.5–2 percent of all industry-year wage changes are pushed up by the floors around 4–5 percent wage growth, while in the 1990s and 2000s, about 1–1.5 of the industry-year wage changes are pushed up above the significant floors around zero, see Appendix C.

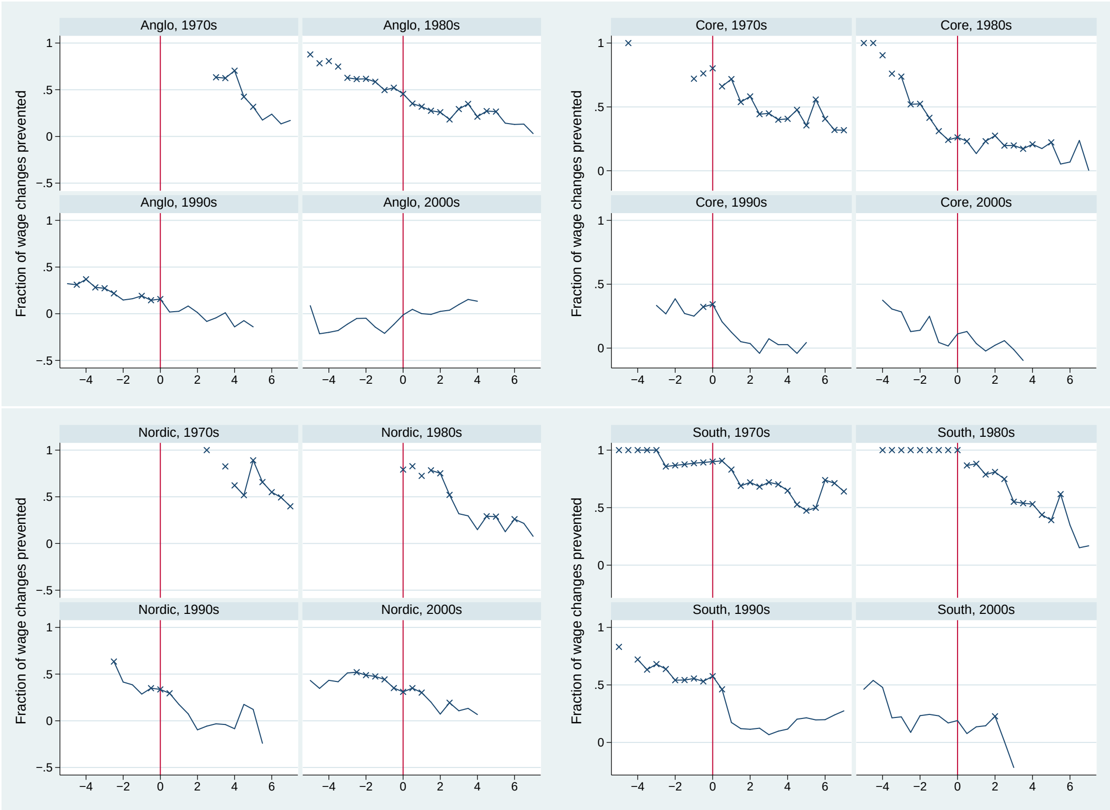

Figure 8: Fraction of wage changes prevented at different floor levels for the regions by period. Significant estimates at the 5 percent level are marked by a "×". Estimates where the number of notional wage changes is at least one percent of the relevant population are connected with lines.

To investigate the cross sectional variation in DNWR, we group the countries into four regions; Anglo (Canada, Ireland, New Zealand, the UK and the US), Core (Austria, Belgium, France, Germany, Luxembourg and the Netherlands), Nordic (Denmark, Finland, Norway and Sweden) and South (Italy, Greece, Portugal and Spain). These regions by and large consist of countries with rather similar labor market institutions (see discussion in Holden and Wulfsberg, 2008). As Figure 7 shows large variation over time, we also distinguish between decades for the regions in Figure 8.

The big picture is fairly similar across regions. In the 1970s, we find significant DNWR in the form of a floor for nominal wage growth around 4 percent in the Anglo region, and at higher levels in the other regions, with associated FWCP of 0.5 or more. In the 1980s, there are significant floors from zero to 5–6 percent wage growth in all regions, with FWCPs around 0.3 in the Anglo and Core regions, and around 0.5 in the Nordic and South regions. In the 1990s, there are significant floors around zero in all regions, with FWCPs ranging from 0.2 percent in the Anglo region to 0.5 percent in the South. In the

2000s, there are significant floors in two regions only; at 2 percent in the South, and at 1 percent in the Nordic, with FWCPs from 0.2 (South) to 0.3–0.5 (Nordic).

The weaker evidence of DNWR in the 2000s is consistent with Gordon's conjecture that the DNWR will weaken over time in periods with low inflation. However, caution is warranted as the downward bias in our estimates is likely to be stronger in country-years with low inflation, and as previous studies have shown that DNWR prevails also in recessions, see e.g. Bewley (1999) and Fehr and Gotte (2005). Note that there is evidence of DNWR in the Nordic countries and in the South, even in the 2000s, in spite of a long period with low inflation in these countries.

3.1 Robustness

In this section we consider the robustness of the results for the common shape assumption, by undertaking three alternative ways of constructing the underlying distribution. In the first, we construct country-specific underlying distributions, G_i, based on all observations for each country. In the second, we construct period-specific underlying distributions, G_τ, one for each decade, based on all observations within the decade. In both cases we then proceed with the method as before. Note that the construction of these alternative underlying distributions includes country-year samples with low median wage growth, where DNWR may bind, increasing the risk for a downward bias in the estimated DNWR.

Finally, we undertake the analysis with a symmetry assumption inspired by Card and Hyslop (1997). Here, the notional distributions are constructed from the empirical ones by replacing observations below the median with observations from the upper half of the distribution in the same country-year. Thus, all country-year notional samples are symmetric, but the shape of the distributions differs across country-years. Note that the symmetry approach involves a test of robustness along two dimensions. First, as the symmetry test involves no assumptions of equal shape across country-year samples, while our main approach makes no assumptions regarding symmetry, these two approaches are based on orthogonal assumptions. Hence, a finding of significant DNWR by both approaches cannot be caused by the same auxiliary assumption. Second, the symmetry test also involves a test for a possible difference between upward and downward nominal wage rigidity. One should be aware, however, that if wages are rigid downwards, and

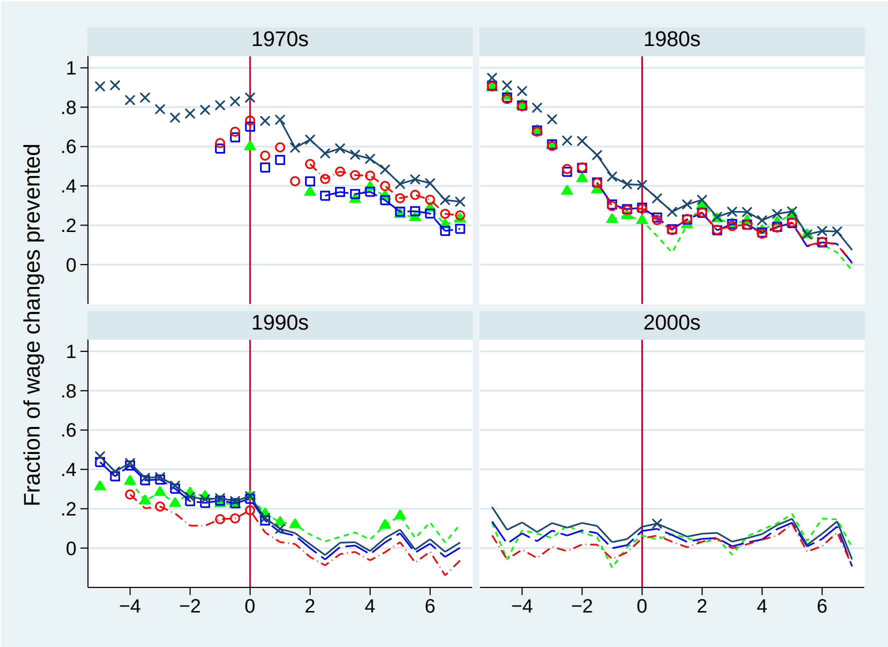

Figure 9: Estimates of the fraction of wage changes prevented at different floor levels by periods. Dark blue lines and x-markers represent estimates based on the default method presented in the paper; red lines and circle markers represent estimates based on country-specific notional distributions; lime green lines and triangle markers represent estimates based on symmetric notional distributions; and blue lines and square markers represent estimates based on period specific underlying distributions.

firms respond to this by attenuating high wage increases, so as to reduce the risk that DNWR is binding in the future (Elsby, 2009), the symmetry approach would involve a downward bias in the estimated DNWR.

Figure 9 compares the main results with estimates from the country-specific, period-specific and symmetric alternatives.[5] We observe that the results are very similar for all specifications. The estimated FWCPs are somewhat lower in the alternative specifications, as expected due to the possible downward bias. Yet the results clearly indicate that our findings are robust. Furthermore, the finding of significant DNWR by use of the symmetry approach implies that the rigidity is asymmetric, with stronger downward rigidity than any possible upward rigidity.

In the present paper, we only consider the possibility of nominal floors for the wage growth. In contrast, several studies have found empirical evidence for the existence of considerable downward *real* wage rigidity in a number of OECD countries, see Dickens et al. (2007), Bauer et al. (2007), and Christofides and Li (2005). Furthermore, Holden

[5]Figures D1–D4 in the appendix illustrate the robustness for the region-period estimates.

and Wulfsberg (2009) find support for some downward real wage rigidity in the same industry data. In view of the problem of distinguishing between real and nominal downward rigidity, we would expect that also some of the rigidity that we find, may in fact by caused by real rigidity. However, we only include country-years where the 35th percentile of the wage change distribution is above the floor, implying that the nominal floors we consider in the present paper are generally considerably below the rate of inflation in the associated country-year. Generally, we also find somewhat higher FWCP for downward nominal rigidity than was previously found for downward real. Thus, we view the floors that we identify as chiefly the result of nominal lower bounds on the wage change process.

3.2 The effect of unemployment, inflation and institutions

From the theory model above, we would expect the prevalence of DNWR to depend on economic variables like inflation (possibly in a non-linear way) and unemployment, as well as institutional variables like the strictness of the employment protection legislation (EPL) and union density. Also other institutional variables like centralization and coordination of wage setting may potentially affect the extent of DNWR. Thus, we regress the extent of DNWR as measured by the FWCP curve in each country-year sample on inflation, inflation squared, unemployment, and institutional variables, to test whether these variables are related to DNWR as measured by the FWCP.[6] We also control for the level of wage growth.

Technically, we undertake Poisson regressions where the number of observed wage changes below floor κ in each country-year sample, $Y(\kappa)_{it}$, depends on the average number of simulated wage cuts for each country-year sample, $\hat{Y}(\kappa)_{it}$, and the explanatory variables mentioned above, $\mathbf{x}_{it}$. A Poisson regression seems appropriate, as $Y(\kappa)_{it}$ is the number of times we observe an event (see Cameron and Trivedi, 1998). The conditional density of the number of observed wage cuts in country-year it in the Poisson model is

$$f\left(Y(\kappa)_{it} = y(\kappa)_{it} \mid \hat{Y}(\kappa)_{it}, \mathbf{x}_{it}\right) = \frac{e^{-\lambda_{it}}\lambda_{it}^{y(\kappa)_{it}}}{y(\kappa)_{it}!}. \tag{7}$$

[6] Messina et al. (2009) explore the variation in downward wage rigidity across sectors, and find that the variation across countries is clearly more important.

Furthermore, we assume that the Poisson parameter, λ_{it}, is given by

$$\lambda_{it} = \hat{Y}(\kappa)_{it} e^{\mathbf{x}'_{it}\boldsymbol{\beta}}, \qquad \text{if } \hat{Y}(\kappa)_{it} > 0. \tag{8}$$

where $\boldsymbol{\beta}$ is the parameter vector we want to estimate. Using the definition of the FWCP and that $\lambda_{it} = \mathrm{E}\left(Y(\kappa)_{it} \mid \hat{Y}(\kappa)_{it}, \mathbf{x}_{it}\right)$ we get

$$1 - \text{FWCP}(\kappa) = \frac{Y(\kappa)_{it}}{\hat{Y}(\kappa)_{it}} = e^{\mathbf{x}'_{it}\boldsymbol{\beta}+\varepsilon_{it}}, \qquad \text{if } \widehat{Y}(\kappa)_{it} > 0 \tag{9}$$

where ε_{it} is an error term.

In the first three columns of Table 1 we present pooled estimates of (9) reporting robust standard errors where the observations are clustered by country. We include country dummies in the second column, and time dummies in the third column. The restriction that $\mathrm{E}(Y_{it} \mid \widehat{Y}_{it}, \mathbf{x}_{it}) = \mathrm{Var}(Y_{it} \mid \widehat{Y}_{it}, \mathbf{x}_{it}) = \lambda_{it}$, an implicit assumption by the Poisson distribution, is accepted easily. We allow the slope of the FWCP-curves to vary between periods as seen from Figure 7 by including the interaction of the floor level with a period specific dummy (denoted as $\kappa \times$ *period*). The positive coefficient on the interaction variables corresponds to the downward slope of the FWCP-curves in Figure 7 (note that the dependent variable is $1 - \text{FWCP}$, i.e. the fraction of wage cuts realized). For example, using time dummies the slope estimate for the 1970s is .153, which implies that the fraction of wage changes realized decreases by a factor of .217 when the floor falls 10 percentage points from from 6 to –4 percent ($\exp(.153(-10)) = .217$). As $1 - \text{FWCP}(6) = .59$ this estimate predicts that $1 - \text{FWCP}(-4) = .13$ or $\text{FWCP}(-4) = .87$, which is close to the "observed" $\text{FWCP}(-4) = .84$. Note, however, that Figure 7 displays a bivariate relationship while the estimates are obtained from a multivariate model.

We find a negative effect of unemployment on the FWCP when country dummies are included, but not in the other specifications. There is a strong positive correlation between inflation and the FWCP. These correlations are consistent with the bargaining model presented above, and illustrated in the left panel of Figure 4: Under high inflation nominal wage cuts are usually small, and a high proportion are prevented by DNWR. In a recession when inflation is low nominal wage cuts are larger, and while DNWR works to

Table 1: Pooled regressions.

Dependent variable	$(1 - \text{FWCP}(\phi))$			Empirical incidence rate		
κ ×1970s	0.054***	0.057***	0.153***	0.332***	0.330***	0.505***
	(0.013)	(0.012)	(0.036)	(0.025)	(0.021)	(0.040)
κ ×1980s	0.083***	0.088***	0.103***	0.494***	0.441***	0.402***
	(0.013)	(0.014)	(0.015)	(0.024)	(0.021)	(0.020)
κ ×1990s	0.069***	0.071***	0.049***	0.477***	0.427***	0.407***
	(0.015)	(0.014)	(0.018)	(0.027)	(0.023)	(0.025)
κ × 2000s	0.048***	0.047***	0.020	0.405***	0.405***	0.398***
	(0.017)	(0.018)	(0.020)	(0.030)	(0.032)	(0.029)
Employment protection	−0.066**	−0.015	−0.074**	−0.400***	−0.203	−0.251***
	(0.033)	(0.098)	(0.033)	(0.069)	(0.184)	(0.057)
Union density	−0.194	−0.847	−0.332**	−1.539***	−2.116**	−0.345
	(0.146)	(0.538)	(0.159)	(0.302)	(0.993)	(0.241)
Centralisation	−0.091*	−0.095	−0.078	−0.196**	−0.254**	−0.283***
	(0.046)	(0.065)	(0.050)	(0.093)	(0.128)	(0.079)
Coordination	0.089**	0.168***	0.072*	0.132	0.311**	0.219***
	(0.040)	(0.063)	(0.041)	(0.088)	(0.125)	(0.073)
Inflation	−0.071***	−0.078***	−0.062***	−0.585***	−0.439***	−0.295***
	(0.024)	(0.026)	(0.027)	(0.034)	(0.034)	(0.039)
Inflation squared	0.000	0.001	0.000	0.016***	0.010***	0.006***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
Unemployment	0.008	0.021*	−0.004	−0.022*	0.079***	0.045***
	(0.006)	(0.011)	(0.007)	(0.012)	(0.022)	(0.013)
Dummies	No	Country	Time	No	Country	Time

Note: Robust standard errors clustered by country in parentheses. The estimates are marked with * if $p < 0.1$, with ** if $p < 0.05$, and with *** if $p < 0.01$. κ is the floor to the wage growth.

reduce the size of the cuts, a larger fraction of them are nevertheless realized.

We also find strong correlation with institutional variables. More strict EPL, higher union density and more centralization are positively correlated with DNWR, and are thus associated with an upward shift in the FWCP-curves. The effects are however only significant in some of the specifications. For example, the effect of EPL is not significant when we include country dummies, presumably due to the limited time variation in the EPL-variable. The effects of EPL and union density are consistent with the results in Holden and Wulfsberg (2008), and also consistent with the theoretical arguments in Holden (2004). Dickens et al. (2007) do not find significant effects of EPL on the extent of DNWR, while the effect of union density is negative, i.e. the opposite of what we find.

That centralization leads to more DNWR is consistent with previous findings that centralization of wage setting leads to wage compression (see Wallerstein, 1999). In contrast, coordination of wage setting seems to induce less DNWR, shifting the FWCP-curve downwards. This is in line with the idea that coordination of wage setting is about ensuring overall wage moderation, without necessarily affecting relative wages. Countries with both centralization and coordination of wage setting thus have about the same DNWR as countries with neither centralization nor coordination. [7]

As a further test of the effect of institutions on DNWR, we exploit the idea that in the absence of DNWR, the institutional variables should not be able to explain the country-year variation in the empirical incidence rate of wage cuts. The last three columns of Table 1 present regressions of the empirical incidence rate of wage changes (below each floor) instead of the fraction of wage changes realized, using the same explanatory variables. Thus, this model does not rely on the notional incidence rates, making it complementary to the former regressions. Specifically, the expected number of wage cuts, λ_{it}, depends on the number of observations (industries) in the country-year sample, S_{it} (rather than $\widehat{Y}_{it}$ as in (8)):

$$\lambda_{it} = S_{it} e^{\mathbf{x}'_{it}\gamma}. \tag{10}$$

Since $\lambda_{it} = \mathrm{E}\left(Y(\kappa)_{it} \mid S_{it}, \mathbf{x}_{it}\right)$ we get

$$q(\kappa)_{it} = \frac{Y(\kappa)_{it}}{S_{it}} = e^{\mathbf{x}'_{it}\gamma + \varepsilon_{it}}, \tag{11}$$

where ε_{it} is an error term. With this specification the poisson restriction $\mathrm{E}(Y_{it} \mid S_{it}, \mathbf{x}_{it}) = \mathrm{Var}(Y_{it} \mid S_{it}, \mathbf{x}_{it}) = \lambda_{it}$ is rejected, hence we use the negative binomial regression model, which allows for "overdispersion."[8]

In accordance with the theoretical predictions, we find a negative effect of EPL, union density, centralization and inflation on the incidence of nominal wage changes below each floor, while coordination and unemployment have a positive effect, even if some of the coefficients are not significant in all specifications.

[7]We have also tried to include lagged FWCP to capture possible dynamic effects, but the coefficient is close to zero and insignificant (not reported).

[8]Overdispersion means that the variance in the data is greater than the mean, in contrast to the Poisson assumption that the variance and the mean are equal. Using a goodness-of fit test from a Poisson regression of Y_{it}/S_{it}, we reject no overdispersion with $\chi^2(11, 721) = 12566.57$. Including a Gamma-distributed error term, ε_{it}, allows the variance-to-mean ratios of Y_{it} to be larger than unity.

4 Concluding remarks

We explore the existence of floors on nominal wage growth on industry data for 19 OECD countries, for the period 1973–2006. For the 1970s and 1980s, we find evidence of floors on nominal wage growth in OECD countries at all rates from minus 2 to plus 6 percent. Thus, there were significantly less nominal wage changes below these growth rates than one would have expected if wage setting had been entirely flexible, for all the four regions we consider, namely Anglo (native English-speaking countries), Southern European countries, Nordic countries, and Core European countries.

The floors on nominal wage growth in the 1970s and 1980s, even if considerably below average inflation (the unweighted average inflation in our sample was 10 percent in the 1970s and 8 percent in the 1980s), may have contributed to causing persistent inflation in this period. Recent contributions (Nelson, 2005 and Meltzer, 2005) have argued that the rise and persistence of high inflation in these years were mainly caused by several errors in the conduct of the monetary policy. Our results add to these explanations by showing that in the 1970s and 80s, an inflationary tendency was entrenched in the wage setting system in many OECD countries. Given the existence of floors on nominal wage growth that were above zero, a tighter monetary policy would have led to further compression of the wage change distribution, inducing both greater wage pressure and compression of relative wages. The upshot would have been greater short run costs in the form of higher unemployment from anti-inflation policies. Thus, our results can help explain why policymakers in most OECD countries failed to pursue a sufficiently tight monetary policy in the 1970s, and why the costs in terms of higher unemployment were so severe when policy finally was tightened in the 1980s and 1990s.

The existence of these floors to growth in nominal wages reflected institutional features in the wage setting, which must be seen in light of the persistent high inflation rates in these decades. However, it is also clear that the existence of the floors was not only a matter of persistent high inflationary expectations. High inflation expectations would affect the location of the wage change distribution, as wage setters would set high nominal wage increases to reach their target real wages, but it should not affect the shape of the distribution. Thus, high inflationary expectations would not by itself compress the lower part of the wage change distribution, which is the effect we find in the 1970s and 1980s.

In the 1990s we also find evidence of a floor on nominal wage growth at zero percent, implying that 17 percent of the industry wage changes that should have been negative, are pushed above zero by binding DNWR. In the 2000s, the evidence is weaker. However, there is some evidence for DNWR in the Southern European countries, and more robust evidence for the Nordic countries. The evidence of DNWR in Southern European countries, Greece, Italy, Portugal and Spain, is especially interesting in light of their membership in the EMU. In these countries high nominal wage growth relative to the productivity growth has led to a steady loss of competitiveness, amplifying the difficult economic situation these countries are in (European Commision, 2010). DNWR would make it difficult to escape from a position with weak international competitiveness. Note also that our bargaining model predicts that wage rigidity will push up the real wage even if the wage is cut, in the sense that the wage cut would have been greater without any rigidity. This implies that as long as there is evidence of wage rigidity, observing widespread wage cuts does not imply that wage rigidity has vanished or is without importance.

The more limited and weaker evidence of DNWR in the 2000s may to some extent reflect institutional changes like lower union density in many countries. It may also reflect a stronger pressure on wages arising from increasing globalization, or changes in wage setting systems, as more flexible pay systems, providing firms with more flexibility to reduce wages. However, our finding of less significant DNWR in the 2000s may to some extent also be due to our method being less able to detect DNWR in a low inflation era.

We also find that DNWR depends on institutional and economic variables, with more prevalent DNWR when employment protection legislation is strict, and union density is high. Coordination of wage setting leads, however, to less DNWR.

Our finding of widespread DNWR in OECD countries over the recent decades raises the question of how this feature of the wage setting has affected other important variables like output, employment and unemployment. Our analysis yields country-year specific estimates of DNWR, albeit noisy, for 19 countries over more that 30 year, hence it provides a good starting point for future work. A considerable extension of the data set in terms of variables would be required, but it seems an interesting avenue for future research.

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Appendix

A Proof of Proposition 1

The equilibrium outcome will be given by the union accepting the firm's offer in step 1, or the firm accepting the union's offer in step 2, and no work stoppage will take place. However, to find the SPE outcome, we must analyze the game backwards, to see what would happen if a player deviates from the equilibrium path. As of step 3, we have the Rubinstein (1982) bargaining game, where players' threat points are determined in the preceding steps. Binmore, Rubinstein, and Wolinsky (1986) show that, in the limit when the time delay between offers converges to zero, the outcome is given by the Nash bargaining solution (assuming for simplicity that the players have equal discount factor).

Lemma 1: If holdout threats prevail in the bargaining, the outcome is

$$W^H = (1+\kappa)W_0, \tag{A1}$$

where κ is defined implicitly by

$$(1-\tau)(1+\kappa)^\eta + (\eta-2)(1+\kappa) - (\eta-1)(1-\varepsilon) = 0, \tag{A2}$$

and where $\kappa > 0$ if and only if $\varepsilon < \frac{\tau}{\eta-1}$ and κ is strictly increasing in τ, and strictly decreasing in ε and η. (Proof, see below.)

Alternatively, if a work stoppage is initiated, the bargaining outcome as of step 3 is given by

$$\frac{W^N}{P} = \text{argmax} \left(\gamma \left(\frac{W}{P}\right)^{1-\eta} \gamma \left(\frac{W}{P} - \frac{R}{P}\right)\right) \tag{A3}$$

which solves for $W^N = \frac{\eta-1}{\eta-2}R$.

Consider now the situation if the firm has rejected the offer by the union in step 2b. The firm can ensure a payoff $\gamma(W^N/P)^{1-\eta} = \gamma\left(\frac{\eta-1}{\eta-2}R/P\right)^{1-\eta}$ by initiating a lockout, and the payoff $(W^H/P)^{1-\eta} = ((1+\kappa)W_0/P)^{1-\eta}$ by letting a holdout prevail. Clearly, the firm will choose the alternative which gives the higher payoff. Thus, it will initiate a lockout if and only if

$$\gamma \left(\frac{\eta - 1}{\eta - 2} \frac{R}{P} \right)^{1-\eta} > \left(\frac{(1+\kappa)W_0}{P} \right)^{1-\eta} \tag{A4}$$

(assuming the convention that no player will initiate a work stoppage when this gives the same payoff as a holdout) or, equivalently,

$$\ell R < (1+\kappa)W_0, \quad \text{where } \ell = \left(\frac{\eta - 1}{\eta - 2} \right) \gamma^{1/(1-\eta)} \tag{A5}$$

In step 2a, the union will offer the highest wage which the firm will accept. The firm will accept an offer if it can get at least as high payoff as it can obtain from rejecting the offer. From the analysis above, this is given by $\min[\ell R, (1+\kappa)W_0]$ (where we have omitted the price level to simplify notation).

Then consider the situation if the union has rejected the offer by the firm in step 1b. If the union initiates a strike, this will lead to a subsequent agreement on W^N (cf. above), giving the union a payoff

$$\gamma \frac{W^N}{P} = s \frac{R}{P}, \quad \text{where } s = \gamma \left(\frac{\eta - 1}{\eta - 2} \right). \tag{A6}$$

(where we have substituted out for W^N from (A3)). Alternatively, if the union does not strike, it will obtain the payoff from step 2, which is $\min[\ell R, (1+\kappa)W_0]$. The union will choose the alternative which gives the higher payoff. Thus, it will strike if and only if $sR > \min[\ell R, (1+\kappa)W_0]$.

In step 1, the firm will offer the highest lowest wage which the union will accept. The union will accept any offer giving it at least the same payoff which it can obtain from rejecting the offer. From the analysis above, this is given by $\max\{sR, \min[\ell R, (1+\kappa)W_0]\}$. (Note that $\ell > s$ so there is an interval where holdout threats prevail.)

Let us sum up and check that the strategies are optimal. If $(1+\kappa)W_0 > \ell R$, the firm will obtain a higher payoff from initiating a lockout than from a holdout. Thus, lockout threats are credible. To avoid a costly lockout, the union will in step 2 offer the firm a wage $W = \ell R$, which will be accepted by the firm – and the firm may as well make the same offer in step 1, which will be accepted by the union (both are equilibrium paths).

Clearly, the union will not threaten to strike, as this gives lower payoff.

If $\ell R \leq (1+\kappa)W_0 \leq sR$, neither player may profit from initiating a work stoppage. Thus, threats of doing so are not credible, and holdout threats prevail in equilibrium, leading to an immediate proposal and acceptance of $W = (1+\kappa)W_0$.

If $(1+\kappa)W_0 < sR$, the union will obtain higher payoff from initiating a strike than from a holdout. Thus, strike threats are credible. To avoid a costly strike, the firm will offer $W = sR$ in step 1, which will be accepted by the union. QED

Proof of Lemma 1

If holdout threats are in use in the from step 3 on, the bargaining outcome is given by

$$\frac{W^H}{P} = \text{argmax}\left(\left(\frac{W}{P}\right)^{1-\eta} - (1-\tau)\left(\frac{W_0}{P}\right)^{1-\eta}\right)\left(\frac{W}{P} - (1-\varepsilon)\frac{W_0}{P}\right) \tag{A7}$$

Note first that the Nash maximand (A7) is zero for $W = (1-\varepsilon)W_0$ and $W = (1-\tau)^{1/(1-\eta)}W_0$ (when one of the terms in (A7) is zero), while it is positive for all values of W^H in the interval between these two values. (If $\tau = \varepsilon = 0$, the interval collapses and the solution is $W^H = W_0$.) Furthermore, the Nash maximand is continuous and differentiable within this interval, so a maximum will exist and be given by the first order condition

$$\frac{(1-\eta)\left(\frac{W^H}{P}\right)^{-\eta}\frac{1}{P}}{\left(\frac{W^H}{P}\right)^{1-\eta} - (1-\tau)\left(\frac{W_0}{P}\right)^{1-\eta}} + \frac{\frac{1}{P}}{\frac{W^H}{P} - (1-\varepsilon)\frac{W_0}{P}} = 0 \tag{A8}$$

which simplifies to

$$\frac{(1-\eta)}{W^H - (1-\tau)W_0(1+k)^{\eta}} + \frac{1}{W^H - (1-\varepsilon)W_0} = 0, \quad \text{where } (1+k) = W^H/W_0 \tag{A9}$$

Multiplying (A9) with W_0 and rearranging, we obtain

$$f(k) = (1-\tau)(1+k)^{\eta} + (\eta-2)(1+k) - (\eta-1)(1-\varepsilon) = 0 \tag{A10}$$

As $f''(k) = (1-\tau)\eta(\eta-1)(1+k)^{\eta-2} > 0$ for all $k > -1$, $f(k)$ is a strictly convex function in $[-1, \infty)$. Furthermore, $f(-1) = -(\eta-1)(1-\varepsilon) < 0$ and $f(k) \to \infty$ when

$k \to \infty$, so there is only one solution to (A10) in $[-1, \infty)$. Observe that (A10) is equivalent to (A2), implying that κ is also the solution to (A9).

For (A9) we see that if $W^H = u$ is a solution to (A9) for $W_0 = x$, then $W^H = tu$ is a solution to (A9) for $W_0 = tx$. Thus, W^H is homogenous of degree one in W_0 so W^H can be written on the form $W^H = (1+\kappa)W_0$ where κ is defined implicitly by (A2). Note also that $(1+k) = W^H/W_0 > 1$ is equivalent to $f(0) < 0$, which again is equivalent to $\varepsilon < \frac{\tau}{\eta-1}$. Thus $\kappa > 0$ if and only if $\varepsilon < \frac{\tau}{\eta-1}$.

The relationship between κ and τ, ε and η follows from implicit differentiation of (A2):

$$-(1+\kappa)^\eta + (1-\tau)\eta(1+\kappa)^{\eta-1}\frac{\partial\kappa}{\partial\tau} + (\eta-2)\frac{\partial\kappa}{\partial\tau} = 0,$$

or

$$\frac{\partial\kappa}{\partial\tau} = \frac{(1+\kappa)^\eta}{(1-\tau)\eta(1+\kappa)^{\eta-1} + (\eta-2)} > 0. \tag{A11}$$

Likewise

$$\frac{\partial\kappa}{\partial\varepsilon} = \frac{-(\eta-1)}{(1-\tau)\eta(1+\kappa)^{\eta-1} + (\eta-2)} < 0, \tag{A12}$$

and

$$\frac{\partial\kappa}{\partial\eta} = \frac{-(1-\tau)(1+\kappa)^\eta\log(1+\kappa) - \kappa - \varepsilon}{(1-\tau)\eta(1+\kappa)^{\eta-1} + (\eta-2)} < 0. \tag{A13}$$

QED

B Data

We have obtained wage data for most countries and years from the ILO. The precise source is Table 5A (Wages by economic activity) and Table 5B (Wages in manufacturing), from yearly statistics under the domain Wages in the *Laborsta* database (`laborsta.ilo.org`). Wage changes are calculated for observations within groups of the same source, worker coverage, sex, type of data, industry classification, country and year. We use observations covering both men and women. If observations for both men and women are not available, we use observations covering only men or only women. The wage variable is labeled either "Earnings per hour" or "Wage rates per hour". When they serve as alternatives, "Earnings per hour" is used. The worker coverage is either "Salaried employees", "Employees" or "Wage earners". The latter is preferred over the others, "Employees" is preferred over "Salaried employees". If there are several series with different sources available we have chosen the series with most observations.

The industry classifications used are ISIC rev. 2 and ISIC rev. 3. The data classified by ISIC rev. 2 cover the groups: Mining and Quarrying (2), Manufacturing (3), Electricity, Gas and Water (4) and Construction (5). The data classified by ISIC rev. 3 covers Mining and Quarrying (C), Manufacturing (D), Electricity, Gas and Water Supply (E) and Construction (F). We use data on various levels of aggregation from the section levels (e.g., D Manufacturing) to group levels (e.g., DA 15 Manufacture of Food Products and Beverages), however, using the most disaggregate level available in order to maximize the number of observations. If, for example, wage data are available for D and DA 15, we use the latter only to avoid counting the same observations twice. The data chosen for Germany is the series labeled "Germany, Fed. Rep. of before 3.10.1990", thus the series relates to the territory of the Fed. Rep. of Germany before 3.10.1990.

Some country-year samples are missing from the ILO database, for which we use data from Eurostat. These country-years are Austria (1974–1999), Belgium (1995), Germany (1973) Spain (1996), France (1976, 1985, 1988, 1994–1996), Ireland (1985), Italy (1973–1985), Portugal (1994, 1995, 1998) and United Kingdom (1980, 1983). The precise source is Table *hmwhour* in the Harmonized Earnings Domain under the Population and Social Conditions theme in the NEWCRONOS database. *hmwhour* is labeled "Gross hourly earnings of manual workers in industry". Gross earnings cover remuneration in cash paid directly and regularly by the employer at the time of each wage payment, before tax deductions and social security contributions payable by wage earners and retained by the employer. Payments for leave, public holidays, and other paid individual absences, are included in principle, in so far as the corresponding days or hours are also taken into account to calculate earnings per unit of time. The weekly hours of work are those in a normal week's work (i.e., not including public holidays) during the reference period (October or last quarter). These hours are calculated on the basis of the number of hours paid, including overtime hours paid. Furthermore, we use data in national currency, and males and females are both included in the data. The data for Germany does not include GDR before 1990 or new Länder. The data are recorded by classification of economic activities (NACE Rev. 1). The sections represented are Mining and quarrying (C), Manufacturing (D), Electricity, gas and water supply (E), and Construction (F). The principles guiding which observation to choose when several are available are the same as described for the ILO data. Data for wages in manufacturing for Italy 2002–2006 are from Istat (*Istituto nazionale di statistica*). Data for Norway are from Statistics Norway. The distribution of the number of observations by years and countries are reported in Table B1. We have removed 3 extreme observations from the sample. The average number of observations per country-year sample is 22.6, with a standard error of 4.9. Table B2 shows the 35th percentile wage growth in each country-year sample.

Data for inflation and unemployment are from the OECD Economic Outlook database.

The primary sources for the employment protection legislation (EPL) index, are OECD (2004) for the 1980–1999 period and Lazear (1990) for the years before 1980. We follow the same procedure as Blanchard and Wolfers (2000) to construct time-varying series, which is to use the OECD summary measure in the 'Late 1980s' for 1980–89 and the 'Late 1990s' for 1995–99. For 1990–94, we interpolate the series. For 1973–79, the percentage change in Lazear's index is used to back-cast the OECD measure. However, we are not able to reconstruct the Blanchard and Wolfers data exactly.

Data for union density before 1990 is from OECD (2004). Data for Greece for 1978 and 1979 are interpolated, while data before 1977 is extrapolated at the 1977 level. From 1990 onwards we have used data from Visser (2006). Data for centralization and coordination are from the OECD (2004, Table 3.5) extrapolated with information from Du Caju et al. (2010).

Table B1: The number of industry observations by countries and years.

	Anglo					Core						Nordic				South				
	CA	IE	NZ	UK	US	AT	BE	FR	DEW	LU	NL	DK	FI	NO	SE	GR	IT	PT	ES	Total
1971	26			24	22		24		25	14	25			21		20				201
1972	26		30	23	22				25		25			21	30	20		24		246
1973	26		30	24	21		26	22	23	17	32	19	18	22	30	20	23	26		379
1974	26	24	30	24	22	16	26	22	27	17	32	19	18	23		20	24	22		392
1975	26	24	31	24	20	16	26	22	27	16	31	19	18	23	30	13	24	17		407
1976	26	23	30	24	20	16	26	22	27	17	31	19	18	23	28	20	24	24		418
1977	26	23	30	24	20	16	26	22	27	18	31	19	18	23	28	20	24	25		420
1978	26	24	30	24	20	16	26	22	27	18	31	19	18	23	26	20	24	28	10	432
1979	26	24	30	24	20	16	26	22	27	18	31	19	18	23	30	20	24	28	10	436
1980	26	24	30	22	20	16	26	22	27	18	31	19	18	23	30	20	24	28	10	434
1981	25	24	30	24		16	26	22	27	14	31	19	18	23	30	20	23	27	10	409
1982	22	24	28	24	21	16	26	20	27	20	31	19	18	23	31	20	24	29	10	433
1983		24	30	24	21	16	26	20	27	22	31	19	18	23	31	11	24	30	10	407
1984	31	23	31	27	21	16	26	20	27	21	25	19	18	23	31	20	24	28	10	441
1985	31	19	31	27	21	16	26	23	27	16	25	19	18	23	31	18	24	28	13	436
1986	31	23	30	30	21	16	26	20	27	17	28	19	19	23	30	20		28	13	421
1987	29	23	30	25	21	16	27	28	27	17	28	19	19	23	31	20		28	13	424
1988	31	23	31	25	18	16	27	23	27	16	27	19	19	23	31	20		27	12	415
1989	31	23		25	23	16	27	18	27	20	27	18	19	23	31	20		28	13	389
1990	31	23	31	25	23	16	27	18	27	11	29	19	19	23	30	20		19	28	419
1991	27	23	31	25	23	16	27	18	27	11	29	19	19	23		20		19	28	385
1992	29	23	31	25	23	16	27	18	27	12	27	19	19	23	16	20		20	27	402
1993	29	20	31	25	23	16	27	18	27	13	25	20	19	23	17	20		20	28	401
1994	23	20	31	24	23	16	27	15	27	12			19	23	17	20		23	28	348
1995	23	19	31	24	23	16	21	10	27	12	20		18	23	17	20		23	28	355
1996	26	19	31	25	23	14	24	12	25	18	21	26		23	17	20		21	24	369
1997	26	19	31	25	21	14	23	26	26	18	22	25		23	15	20		21	26	381
1998	26	19	31	26	21	14	23	26	26	18	22	26		23	17	20		29	26	393
1999	26	15	25	27	21	14		26	26	18	22	25		23	17			16	26	327
2000	26	15		27	21	25	26	26	24	18	22	26	11	23	17				26	333
2001	26	15	24	27	21	25	26	26	24	18	22	26	11	23	22			16	26	378
2002	26	15	24	27	18	25	26	25	24	18	22	26	11	23	22		22	16	26	396
2003	26	15	24	27	15	24	25	25	24	18	22	26		23	22		22		26	364
2004	26	15	24	26	15	24	25	24		18	22	26	11	23	22		22	17	26	366
2005	26	15	23	26	15	25	25	26		18	22	26		24	22		22	17	26	358
2006	26	15		25		25				18					22		22		26	179
Total	939	677	935	904	703	581	848	709	866	585	902	657	467	801	821	542	420	752	585	13,694

Table B2: The 35th percentile wage growth by countries and years. Percent.

	Anglo					Core						Nordic				South			
Year	CA	IE	NZ	UK	US	AT	BE	FR	DEW	LU	NL	DK	FI	NO	SE	GR	IT	PT	ES
1971	8.3			10.0	6.4		11.8		11.4	8.0	14.1			11.6		6.0			
1972	7.4		8.5	13.2	6.5				8.8		12.3			10.0	11.3	5.1		13.8	
1973	8.1		13.8	12.3	5.9		13.0	14.1	10.9	12.6	14.5	17.0	16.5	10.0	7.7	13.5	24.2	12.7	
1974	12.4	18.8	15.5	19.6	7.3	14.7	24.2	20.3	10.7	17.9	16.5	18.2	20.8	16.4		29.4	20.6	46.7	
1975	15.6	25.8	13.2	24.1	8.2	10.9	15.7	15.2	7.4	17.8	13.1	17.3	20.3	18.3	15.8	29.2	23.8	37.7	
1976	12.4	15.0	13.6	11.2	7.6	9.1	10.2	14.3	6.3	13.1	7.3	11.4	13.9	16.1	12.1	23.1	21.9	18.3	
1977	9.6	13.0	13.8	7.8	7.9	8.2	8.2	12.2	7.0	6.1	10.0	9.9	8.2	12.1	8.0	17.8	21.0	14.0	
1978	6.8	14.9	13.4	14.9	7.8	5.3	5.6	11.2	5.2	2.8	5.1	8.9	7.0	9.1	11.6	25.0	14.6	13.3	28.7
1979	8.7	16.9	14.5	16.0	8.0	4.9	6.6	11.4	5.3	5.0	4.1	10.2	10.8	4.8	8.3	20.9	16.7	18.2	23.6
1980	9.7	13.6	11.3	16.5	8.5	6.5	8.3	15.2	6.1	8.1	4.7	8.4	11.8	8.0	10.3	25.0	18.5	21.5	16.8
1981	12.0	18.8	18.0	8.6		6.4	11.2	14.1	5.5	5.6	5.2	9.2	12.2	10.5	8.7	23.5	24.0	19.2	18.0
1982	10.6	13.8	16.3	7.7	6.8	5.6	5.1	12.8	4.5	6.1	6.4	9.4	10.1	10.3	5.2	35.4	16.0	19.2	15.3
1983		8.5	3.4	8.7	4.5	4.0	4.9	12.4	3.3	7.0	1.5	6.8	9.7	8.4	5.9	16.6	14.9	17.7	14.8
1984	4.5	8.8	1.1	6.2	3.2	4.1	3.7	5.6	2.3	3.6	0.8	4.6	8.7	9.4	10.6	15.5	5.3	15.3	10.1
1985	3.0	6.9	6.2	6.8	3.2	5.4	4.0	5.6	3.0	3.3	1.7	4.3	7.5	8.7	7.8	16.1	10.1	20.4	8.0
1986	2.6	6.1	17.8	6.5	1.9	4.0	-0.0	2.8	3.2	1.3	1.1	3.3	5.7	8.7	7.0	11.5		18.7	8.1
1987	2.7	3.3	11.3	7.0	1.8	2.5	1.8	3.9	3.9	2.6	2.1	9.3	7.0	12.8	6.4	9.9		13.6	7.4
1988	3.1	4.7	8.2	6.4	2.7	3.3	1.7	1.4	3.8	3.2	1.4	6.1	7.6	5.5	7.5	16.8		11.9	7.4
1989	4.2	3.4		8.0	2.9	2.8	4.0	2.9	3.6	5.5	2.5	4.0	7.0	4.1	9.1	20.2		15.6	9.7
1990	3.9	5.0	4.6	9.1	3.2	6.1	4.7	4.3	4.9	3.9	3.4	3.8	8.9	5.8	9.4	19.0		13.5	9.1
1991	3.8	3.6	2.9	9.5	3.1	5.1	5.4	4.4	5.9	4.5	3.3	4.3	4.1	5.0		15.7		12.6	10.4
1992	2.8	3.7	1.6	6.1	2.3	5.3	4.1	3.6	5.3	7.7	5.4	3.2	1.7	2.4	7.1	13.0		12.4	6.5
1993	0.6	5.7	-0.1	4.2	2.1	3.2	3.7	1.9	4.9	4.8	3.4	1.5	1.5	2.6	-0.1	10.5		6.6	5.6
1994	0.6	1.0	0.1	0.8	2.4	3.7	3.3	-2.1	2.6	3.2			2.9	3.0	3.4	12.5		3.8	3.5
1995	1.4	1.9	2.7	3.0	2.0	3.6	-5.9	2.6	3.1	2.8	2.3		6.1	4.4	2.9	10.6		4.0	4.3
1996	1.6	1.5	2.9	2.9	2.7	3.6	2.3	3.2	1.9	-0.7	3.4	5.3		4.1	6.2	7.3		5.3	3.0
1997	0.4	2.2	3.5	2.0	2.6	1.6	1.9	1.3	1.1	3.2	2.7	2.0		4.4	3.0	8.3		6.2	3.7
1998	0.9	2.5	2.5	5.1	3.0	2.1	2.6	2.0	1.5	1.3	3.2	3.1		5.4	2.8	4.3		1.3	3.6
1999	-0.4	4.9	2.0	2.5	2.8	1.8		1.2	2.2	3.3	3.7	5.3		5.1	1.8			11.4	2.1
2000	1.0	5.0		3.0	3.5	1.9	2.9	4.4	2.0	3.4	2.9	1.3	4.7	4.7	3.6				2.3
2001	1.2	8.2	2.1	4.1	2.5	2.3	2.7	4.2	1.3	-0.3	5.3	4.2	3.3	5.7	2.4			6.6	2.9
2002	2.4	6.3	2.6	4.2	2.3	2.9	1.5	7.8	2.1	3.4	2.7	3.4	2.8	6.4	1.8		2.4	9.3	3.4
2003	0.9	4.5	-0.5	2.7	2.5	2.6	3.1	3.1	2.1	3.0	3.0	3.2		3.0	3.2		2.3		4.0
2004	0.4	2.9	2.5	2.3	2.5	1.7	2.6	2.6		3.8	1.4	0.3	2.7	2.4	2.9		1.7	2.4	3.0
2005	1.3	3.5	0.5	3.2	1.0	2.3	1.8	1.7		3.2	0.9	3.0		3.6	2.3		2.3	0.6	1.9
2006	0.3	3.1		2.7		2.5				-0.4					2.8		2.4		4.5

C The fraction of industry-years affected

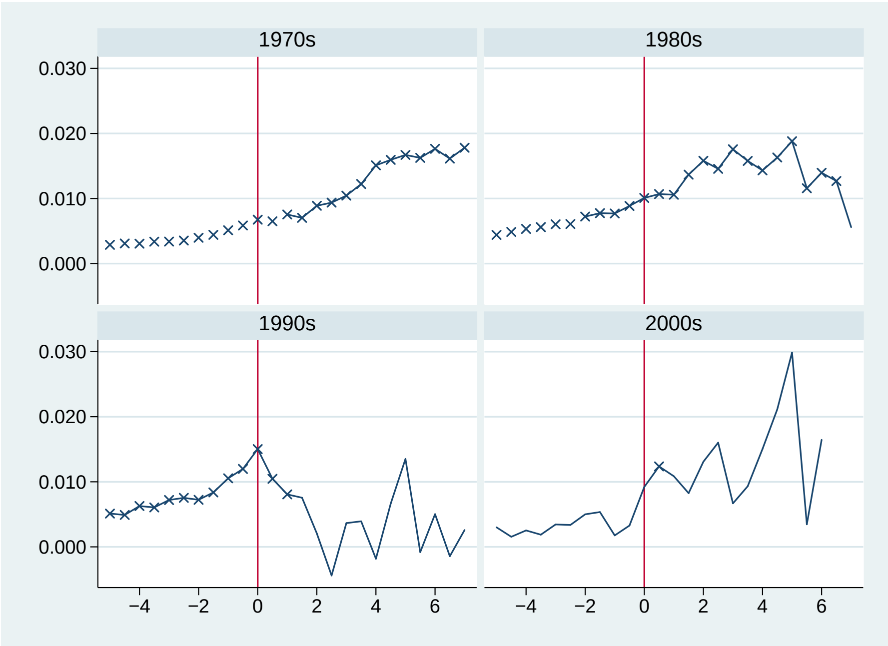

Figure C1: The fraction of industry-years affected $(q - \tilde{q})$ *by periods.*

D Robustness

Figure D1: Fraction of wage changes prevented at different floor levels by periods for the Anglo countries.

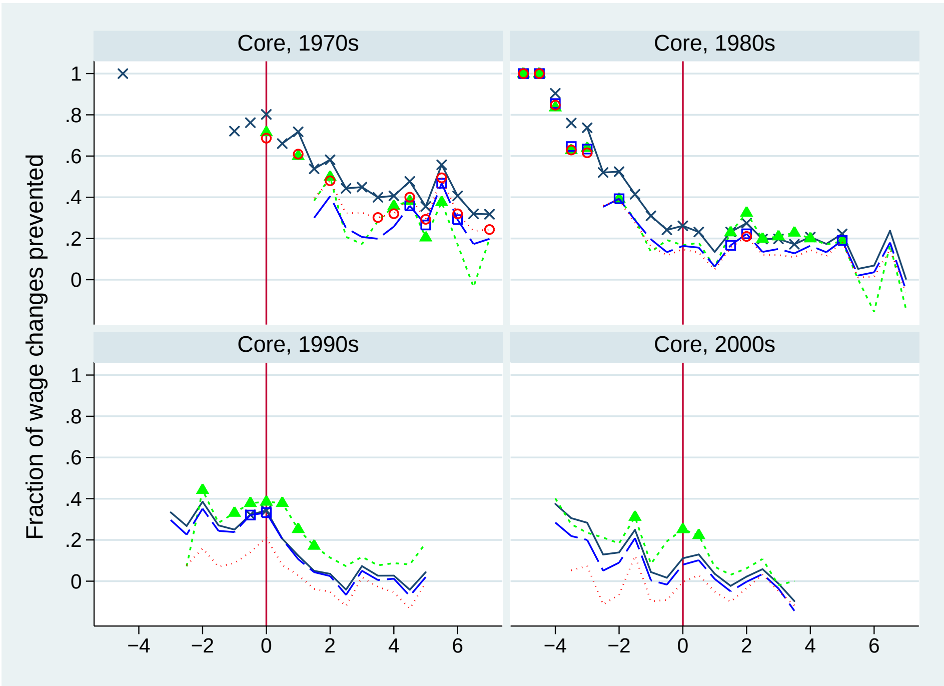

Figure D2: Fraction of wage changes prevented at different floor levels by periods for the Core region.

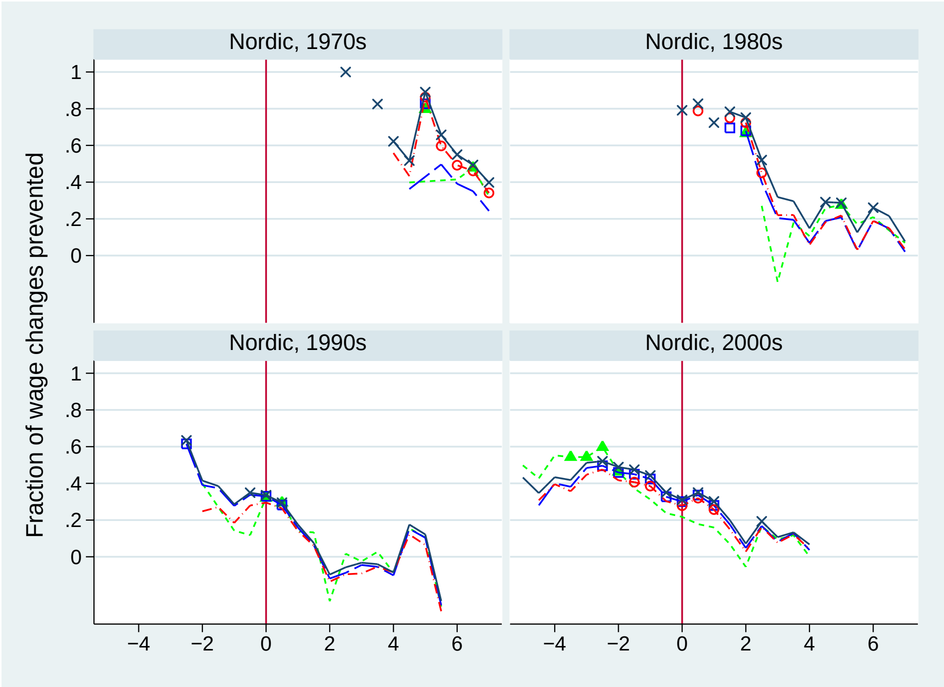

Figure D3: Fraction of wage changes prevented at different floor levels by periods for the Nordic region.

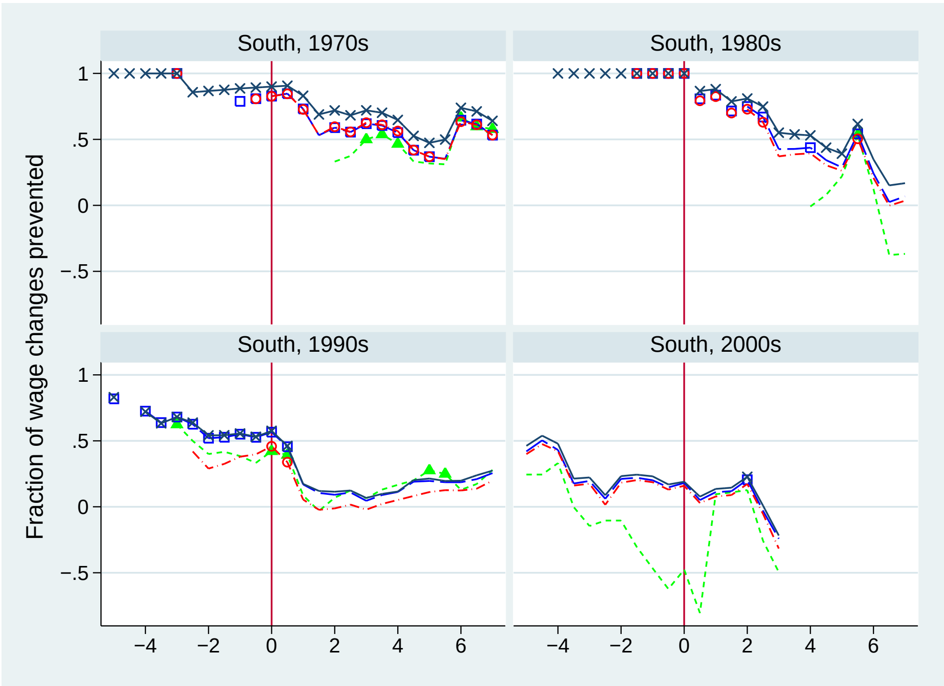

Figure D4: Fraction of wage changes prevented at different floor levels by periods for the South region.