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Wage Rigidity, Inflation, and Institutions*

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Abstract

We study the possible existence of downward nominal wage rigidity (DNWR) at wage growth rates different from zero in aggregate data. Even if DNWR prevails at zero for individual workers, compositional effects might lead to falling aggregate wages, while changes in relative wages combined with DNWR might lead to positive aggregate wage growth. We explore industry data for 19 OECD countries, over the 1971–2006 period. We find evidence for a floor on nominal wage growth at 6 percent in the 1970s and 1980s, at 1 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.

Keywords: Downward nominal wage rigidity; OECD; wage inflation; wage setting

JEL classification: *C*14; *C*15; *E*31; *J*3; *J*5

I. Introduction

Following the 2008–2009 financial crisis, several countries in the European Monetary Union (EMU) are faced with soaring unemployment and severe problems with high wage costs relative to the productivity level. Will they be able to cut wage costs in order to alleviate this situation? The strong evidence for downward nominal wage rigidity (DNWR) for job stayers in many OECD countries suggests that cutting wages might be difficult in a low inflation environment (e.g., Bewley, 1999; Dickens *et al.*, 2007;

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Knoppik and Beissinger, 2008). However, in some countries, we do observe extensive wage cuts – does this mean that DNWR is not a problem?

A growing body of literature shows that DNWR might have important implications for macroeconomic outcomes under low inflation; these studies include Akerlof *et al.* (1996), Carlsson and Westermark (2008), Kim and Ruge-Murcia (2009), Fahr and Smets (2010), and Benigno and Ricci (2011). In this paper, we consider the interaction between inflation and DNWR. We present a theoretical model of DNWR, which we then use to derive several novel empirical predictions for how DNWR affects the wage change distribution. We show that the effect of DNWR depends on the size of the counterfactual wage cuts, with a larger than proportional reduction for the large wage cuts. One implication of this finding is that the fraction of counterfactual wage cuts that is prevented by DNWR is likely to be increasing as a function of the inflation rate and decreasing as a function of the unemployment rate. Importantly, we show that even if the wage is cut, the resulting wage will be higher than if the wage-setting process had been completely flexible. This result contradicts the common assumption in the literature that if the wage is cut, in spite of wage rigidity, then the wage after the cut will be equal to the flexible wage, as if there had been no rigidity (e.g., Knoppik and Beissinger, 2003; Fehr and Gotte, 2005). Thus, the presence of wage cuts is not sufficient evidence to conclude that DNWR has no effect on the aggregate outcome.

In the main part of the paper, we consider the prevalence of DNWR using industry data over the 1973–2006 period for 19 OECD countries, consisting of more than 13,000 observations. This section extends our previous work (Holden and Wulfsberg, 2008) in several directions. First, we explore empirically the extent of DNWR at wage growth rates both below and above zero. The motivation is partly empirical. For example, individual DNWR at zero might be transformed to a floor below zero for aggregate wages (e.g., because older high-wage workers are replaced by younger workers 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. However, the issue is also important from a policy point of view, because it is crucial to determine how high inflation must be to avoid binding DNWR.

Second, 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 of the Great Depression, they both claimed, in Gordon's words, that "nominal wage reductions would no longer be seen as unusual if the average nominal wage was not growing" (Gordon, 1996, p. 62). Most OECD countries have now experienced more than a decade of fairly low and stable inflation, and it is important to explore whether wages are still prone to downward rigidity.

Our study does not explore whether DNWR has any effect of on output and employment. Yet, we argue that our study is also of interest when viewed from this angle. If there is no sign of DNWR in industry-level wage data, we can also question whether the DNWR found in micro data has any noticeable effect on industry output or employment. However, if there is DNWR in industry-level wage data, rigidity prevails in spite of varying compositional effects. In this case, it also seems more likely that DNWR will affect industry output and employment. Exploring the existence of such effects seems to be 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 the 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 to about zero, with a 20 percent fraction of notional wage cuts prevented by DNWR. In the 2000s, we find evidence of a floor in the Nordic countries at 0–1 percent wage growth, and also some evidence of a floor for Southern European countries, as well as for the OECD economies as a whole. We find that DNWR is strongly correlated with inflation and labor-market institutions. The evidence for DNWR in Southern European countries is important in light of their troubled economies and weak competitive positions (e.g., European Commission, 2010), because these countries are part of the euro area and thus are unable to devalue to restore competitiveness.

The rest of the paper is organized as follows. In Section II, we provide a simple bargaining model predicting wage floors that can differ from zero percent, and we discuss the link between DNWR and aggregate wage growth. In Section III, we present the data and empirical approach. In Section IV, we provide the main results as to the extent of DNWR, while in Section V, we explore whether the variation in DNWR across countries and time can be explained by institutional and economic variables. We conclude in Section VI. The online appendices contain supplementary material on data and results.

II. DNWR and Floors for Aggregate Wage Growth

In this section, we discuss the theoretical implications of DNWR that we explore in the empirical analysis. We consider a simple model of DNWR under firm-level bargaining, drawing upon Holden (2004). The use of a bargaining model reflects the fact that in most OECD countries, collective wage agreements cover the large majority of the workforce (see OECD, 2004, Table 3.3).[1] Let $\pi(W/P) = (W/P)^{1-\eta}$ denote the real revenue flow per worker to the firm where the real wage, W/P, denotes the flow pay-off to the workers; for simplicity, this model neglects any disutility of labor.[2] P is the aggregate price level, which is exogenous to the parties taking part in the bargaining process, while η is the elasticity of demand, where $\eta > 2$. There is a prevailing wage contract, W_0, which is given in nominal terms. Consistent with the institutional settings in most OECD countries, we assume that the contractual terms can only be changed by mutual consent, even if the contract has a fixed duration (see MacLeod and Malcomson, 1993; Holden, 1994). This assumption implies that the terms of the old contract will prevail after it expires, until the players have agreed to a new contract, or unless one of the players unilaterally stops production.

A crucial feature of this model is that the type of bargaining dispute is endogenous (see Haller and Holden, 1990; Cramton and Tracy, 1992). The workers can initiate a strike or the firm can initiate a lockout. In both cases, the firm obtains zero pay-off, while the workers receive R/P, the alternative real income during a work stoppage. If neither of the two players stop production, production continues under the prevailing wage contract, W_0, while the parties are bargaining (i.e., a holdout). Holdout has received scant attention in the literature, yet holdouts are more frequent than strikes, in both the US (Cramton and Tracy, 1992) and the Netherlands (Van Ours and van de Wijngaert, 1996). Following Cramton and Tracy (1992), Moene (1988), and Holden (1997), we assume that the workers can inflict a cost τ 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 ε on workers (e.g., by reducing bonus schemes). Formally, this situation can be captured by assuming that the pay-offs 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 $0 \le \tau < 1$, $0 \le \varepsilon < 1$. A holdout is costly to both players if $\tau > 0$ and $\varepsilon > 0$.

A second important assumption in the model is that if a work stoppage (strike or lockout) 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 period of time before work can resume after a work

[1] DNWR is often motivated from fairness considerations (e.g., Akerlof *et al.*, 1996). Such effects could have been added to the model, giving an additional source of DNWR.

[2] This profit function follows from a model of monopolistic competition, in which each firm i sets the output price facing a downward sloping demand curve $Y_i = (P_i/P)^{-\eta}$, and there is constant returns, with labor the only factor of production $Y_i = N_i$. The profit-maximizing price is $P_i = \upsilon W_i$, where $\upsilon = \eta/(\eta-1)$. Substituting out in the real profit function, we obtain $\pi_i/P = (P_iY_i - W_iN_i)/P = (\upsilon-1)\upsilon^{-\eta}(W_i/P)^{1-\eta}$. In the main text, we have omitted irrelevant constants and omitted the index for the wage in firm i, because all firms are identical and will set the same wage.

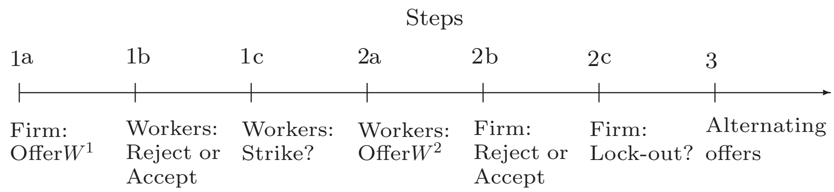

Fig. 1. The wage bargaining game

stoppage. Alternatively, in an extended model with risk aversion and an uncertain bargaining outcome, the fixed costs would be the expected pay-off that 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 resumes after a work stoppage, the player's pay-offs are $\gamma\pi(W/P)^{1-\eta}$ and $\gamma W/P$, where $\gamma < 1$, so that $1-\gamma$ is the reduction in pay-off caused by a work stoppage.

To determine the type of threats that might prevail, we use the extension of the Rubinstein (1982) alternating offers bargaining game employed by Holden (2004). The game is illustrated in Figure 1. Steps 1a-2c take place in negligible time, and determine which type of dispute (strike, lockout, or holdout) prevails in the bargaining. In Step 3, a standard Rubinstein bargaining games starts, whereby players alternate in making offers, one offer per time period. In each of the first two steps, one of the players makes an offer, which the opponent can reject or accept (in which case the bargaining ends). Upon rejection, the player who declined the offer can decide whether to initiate a work stoppage. A player is assumed to initiate a work stoppage only if this gives him/her strictly higher pay-off than the alternative. If no work stoppage is initiated, a holdout will ensue from step 3 onward. In equilibrium, an agreement is reached in step 1 or 2, with no costly dispute (see Appendix A).

Proposition 1. *The unique SPE outcome to the wage bargaining is W given by*

$$W = \begin{cases} \ell R, & \textit{if } (1+\kappa)W_0 > \ell R \quad (\textit{lockout}) \\ (1+\kappa)W_0, & \textit{if } \ (1+\kappa)W_0 \in [sR, \ell R] \quad (\textit{holdout}) \\ sR, & \textit{if } \ (1+\kappa)W_0 < sR \quad (\textit{strike}), \end{cases} \tag{1}$$

where $\ell = [(\eta-1)/(\eta-2)]\gamma^{1/(1-\eta)} > s = [(\eta-1)/(\eta-2)]\gamma > 1$.[3] *Here,* κ*, the wage floor, is defined implicitly by* $(1-\tau)(1+\kappa)^{\eta} + (\eta-2)$

[3] 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, however, does not change.

$(1-\varepsilon)=0$. *Furthermore, κ is strictly increasing in τ, strictly decreasing in ε and η, and $\kappa > 0$ if and only if $\varepsilon < \tau/(\eta-1)$.*

The intuition for Proposition 1 is as follows. If holdout threats prevail in the bargaining, the outcome depends on the prevailing contract wage W_0 because it determines the players' pay-offs 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 if the holdout threats prevail.

However, holdout threats will not always prevail. If the workers would have obtained a higher pay-off from a strike than from a holdout (i.e., if $sR > (1+\kappa)W_0$), the workers can credibly threaten to strike. To avoid a strike, the firm offers a wage $W = sR$, which gives the workers the same pay-off as they could obtain by initiating a strike; hence, the workers accept the firm's offer. Alternatively, if the holdout pay-off is higher, $(1+\kappa)W_0 > \ell R$, it would be the firm that would use lockout threats to push wages down. The workers would then offer $W = \ell R$ in step 2, and the firm would accept it. Finally, if $(1+\kappa)W_0 \in [sR, \ell R]$, threats to initiate a work stoppage are not credible, because initiating a work stoppage gives a lower (or the same) pay-off to the initiating player than he or she 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, as a function of the alternative income R. We see that lockout threats prevail for low R, holdout threats prevail for intermediate R, and strike threats prevail for high R. Thus, for a given R, wages are higher if lockout threats prevail than if strike threats prevail. This difference is key for the existence of DNWR, because the old wage affects which type of threats will prevail, and therefore also will affect the new wage. If there are no fixed costs incurred from a work stoppage, $\gamma = 1$, we would be back to using the standard wage bargaining model without DNWR (e.g., Layard *et al.*, 1991) because strike and lockout threats would result in the same pay-offs, and there would always be one player who could benefit from threatening to stop work. Then, $s = \ell$, implying that the holdout interval would vanish and the bargaining outcome would always be $W = sR$. Wages would be completely flexible, and the relative wage growth would be equal to the relative growth in the alternative income (i.e., $\Delta w = \Delta r$, where the lowercase letter denote logs and Δ is the difference operator).[4]

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[4] However, if we assume that the pay-offs are relatively more favorable to the workers during a lockout than during a strike, this would push the ℓR line in Figure 2 further in a

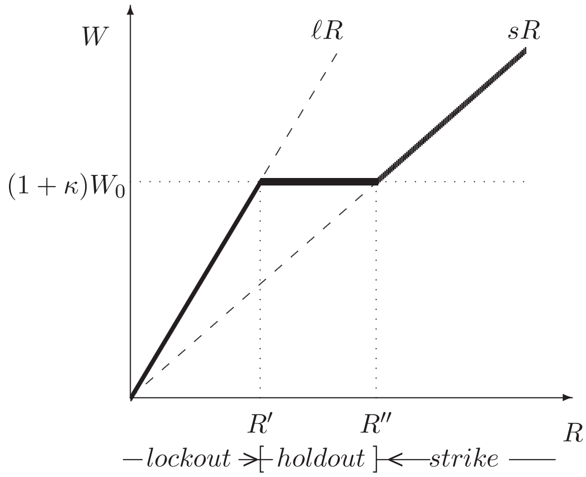

Fig. 2. The wage bargaining outcome (solid line) as a function of the alternative income R. For $R < R' = (1+\kappa)W_0/\ell$, the firm pushes wages down to ℓR by use of lockout threats. For $R > R'' = (1+\kappa)W_0/s$, the union pushes wages up to sR by use of strike threats. For $R \in [R', R'']$,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

To consider the implications for wage growth, now assume that there are annual wage negotiations driven by changes in R.[5] Most years, R will increase, as a result of productivity growth and inflation. Thus, strike threats will prevail. The left panel of Figure 3 shows the wage growth as a function of the growth in the alternative income, assuming that strike threats prevailed in the previous year, $W_0 = sR_0$. (A formal treatment is given in Appendix A.) If the alternative income increases more than the wage growth during a holdout, $\Delta r \geq \kappa$, strike threats continue to prevail in the bargaining over wages and $\Delta w = \Delta r$. However, if the alternative income grows less, so that $\Delta r^L(\gamma) < \Delta r < \kappa$, the workers use holdout threats to obtain a higher wage increase than the growth in the alternative income, as $\Delta w = \kappa > \Delta r$. If the alternative income grows even less, the firm will prevent the holdout by use of lockout threats, thus pushing wage

counterclockwise direction, increasing the size of the holdout interval, so that this interval would exist also for $\gamma = 1$. In most countries, this would be empirically plausible, for example, because of legal restrictions on firms' use of a lockout to push down wages (as in some European countries), because union workers are more likely to be able to collect unemployment benefits during a lockout (as in the US), or because the union would save on costs associated with enforcing the strike.

[5] Holden (1997) considers a similar model extended to an multi-annual setting, where 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 reduced.

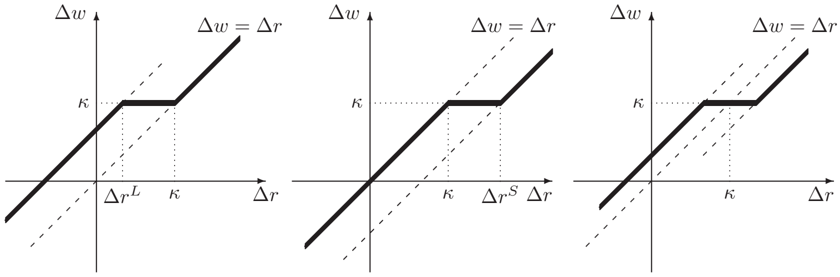

Fig. 3. Wage growth as a function of growth in the alternative income when strike threats prevailed in the previous year (left), when lockout threats prevailed (middle), and when holdout prevailed (right).

growth down. The threshold $\Delta r^L = (1+\kappa)\gamma^{\eta/(\eta-1)} - 1$ is an increasing function of γ. The smaller γ is, the weaker is the potency of lockout threats, and the wider is the interval where holdout threats prevail.

An important property apparent from the left panel of Figure 3 is that if $\Delta r < \kappa$, wages will grow at a higher rate than the alternative income, because holdout or lockout threats then prevail. This implies that the DNWR will have an upward effect on the wage even when the wage falls, in the sense that the wage would have fallen more without DNWR. This claim contrasts with the standard assumption that if the wage is cut, then wage rigidity has no effect on the actual wage outcome (e.g., Knoppik and Beissinger, 2003; Fehr and Gotte, 2005).

The middle panel of Figure 3 shows the wage growth if lockout threats prevailed during the previous year, $W_0 = \ell R_0$ (empirically, this is rare but it can happen when R falls). Strike or holdout threats prevail when $\Delta r > \kappa$, leading to $\Delta w < \Delta r$, while lockout threats continue to prevail when $\Delta r \leq \kappa$, thus implying that $\Delta w = \Delta r$. Thus, when the alternative income has fallen in the previous negotiation, the rigidity might, in fact, lower a subsequent wage increase.

Finally, the right panel of Figure 3 shows the wage growth if holdout threats prevailed in the previous year. If the alternative income grows at a rate above κ, wages will grow at a lower rate than Δr – the intuition being that wages were already slightly higher than the strike outcome the previous year as a result of the holdout threats. Correspondingly, if the alternative income grows less than κ, wages will grow more than Δr.

The model's parameter values can 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 are likely to strengthen the workers' side in the

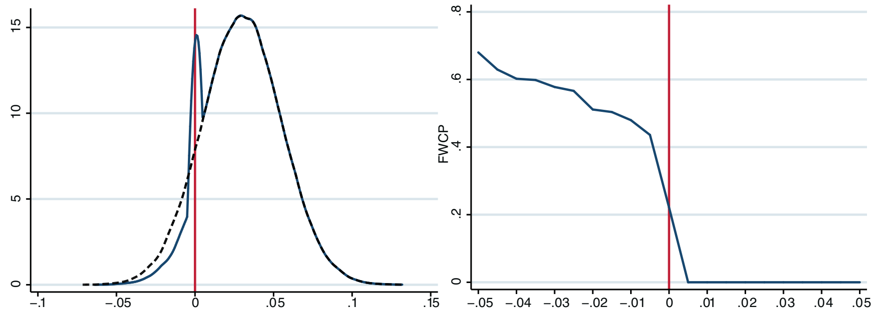

Fig. 4. Left: density of a notional distribution of nominal wage changes (dotted line) and a distribution of nominal wage changes subject to a wage floor at 0 percent (solid line). Right: FWCP curve at different levels of wage growth derived from the distributions in left panel. Parameter values are $\eta = 3$, $\gamma = 0.995$, and $\kappa = 0$.

negotiations, making it more difficult for the firm to enforce a wage cut. In the short run, these parameters can be taken as given. However, in the longer run, they are clearly endogenous. For example, in the 1970s and 1980s, most remuneration was in the form of fixed pay, with less scope for the firm to reduce wages during a holdout; that is, ε was close to zero, implying that the nominal wage change during a holdout, κ, was likely to be considerably above zero. With lower inflation, firms 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 (i.e., lowering the nominal wage change during a holdout).

The implications of the DNWR bargaining model for the distribution of wage changes can be illustrated by a simple numerical example. Assume that strike threats prevailed in all industries during the previous year, so that $W_0 = sR_0$. Furthermore, let Δr be 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 $\gamma = 1$ and $\Delta w = \Delta r$, about 12 percent of the observed wage changes are below zero, illustrated by the dashed line in the left panel of Figure 4. Adding DNWR by setting $\gamma = 0.995$, strike threats will continue to prevail for the industries when $\Delta r \geq \kappa = 0$. For industries where $\Delta r < \kappa = 0$, holdout threats will prevail and will prevent wages in some industries from falling. This leads to a distribution of wage changes with a deficit below zero, indicated by the solid line.

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

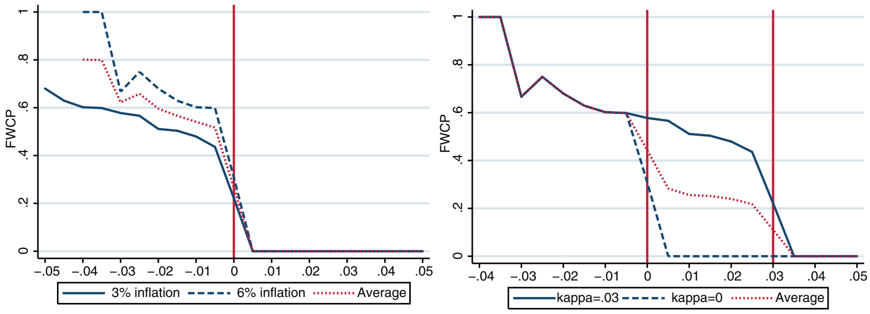

Fig. 5. Left: the FWCP curves with a wage floor at zero percent when mean wage growth is 3 percent (solid line) and 6 percent (dashed line). Right: the FWCP curves with wage floors at zero percent (dashed line) and three percent (solid line) when mean wage growth is 6 percent.

cuts divided by the number of wage cuts that would have been enacted in the absence of rigidity; a precise definition is given in equation (6). 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 (0) 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 ones. The intuition here is that while DNWR prevents some small wage cuts, it also attenuates larger notional wage cuts; see the left panel of Figure 3, where $\Delta w > \Delta r$ for $\Delta r < \kappa$. This reduces the deficits of smaller cuts.

Figure 5 illustrates the theoretical predictions for how the FWCP curve depends on the average wage change (Δr) and the wage floor (κ). First, the left panel of Figure 5 shows that a higher mean wage growth leads to a higher FWCP. This is because when the wage change distribution moves to the right, a larger part of the notional wage cuts are pushed above the floor and no wage cut is realized. This feature of the model implies that any factor that leads to higher notional wage growth, such as higher productivity growth, lower unemployment, or higher inflation, also leads to a higher 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.

Note that the downward rigidity here is conceptually very different from downward real wage rigidity (DRWR). The location of the rigidity under DRWR depends on the rate of inflation, because workers resist a reduction in the real wage. Here, the location of the rigidity is a given nominal growth rate depending on the wage growth under holdout threats (κ). Even in this model, inflation will indirectly affect the FWCP because inflation affects the location of the notional wage distribution (see Figure 5, left), but annual fluctuations in inflation will not affect the location of the rigidity at κ.

Both panels of Figure 5 also illustrate how the aggregate FWCP might look (dotted line) when averaged over two different wage-setting regimes, which could arise from heterogeneity within or across industries. The right panel shows aggregation over different wage floors (different values of κ), where half the wage change observations have a floor at zero, and the other half have a floor at 3 percent. In the empirical analysis, we estimate FWCP curves comparable to those displayed in Figure 5, and investigate how these are affected by inflation, unemployment, and institutional factors.

III. 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 1973–2006 period. The countries included in the sample are Austria, Belgium, Canada, Germany, Denmark, Finland, France, Greece, Ireland, Italy, Luxembourg, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, the UK, and the US. The main data source is the wages in manufacturing obtained from the ILO and harmonized hourly earnings in manufacturing from Eurostat (see the online appendices). One observation is denoted Δw_{jit}, where j is the index for the industry, i is the index for the country, and t is the index for the year. In total, there are 13,694 observations distributed across 604 country-year samples, on average 23 industries per country-year.

Because the observational unit is the change in the average hourly earnings in an industry, it will be affected by both the average wage change for the existing workforce, and the compositional effects as a result of wage differences between newly hired workers and workers that leave the industry. Compositional changes and turnover might smoothen out the effect of DNWR at the level of individual workers. Likewise, if firms respond to rigid base wages by reducing bonus schemes and fringe benefits, consistent with the evidence found by Babecký *et al.* (2010), the effect of DNWR might 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 growth rate different from zero. The possible existence of such floors, and the form these can take, will be key in our empirical analysis.

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 during recessions, this effect will dampen the downward pressure on wages in the recession (Solon *et al.*, 1994).

The common approach in the DNWR literature (e.g., Card and Hyslop, 1997; Knoppik and Beissinger, 2003; Lebow *et al.*, 2003; Nickell and Quintini, 2003; Holden and Wulfsberg, 2008) is to construct a distribution of notional or rigidity-free wage changes and then to explore whether the empirical wage change distribution is compressed from the left as compared to the notional distribution (see Figure 4). From the theoretical model in Section II, the notional relative wage change 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 such as 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-year, and possibly also by inflation, productivity growth, and unemployment.

Specifically, we follow Holden and Wulfsberg (2008) and assume that in the absence of any DNWR, the notional nominal wage growth in industry j in country i in year t, Δw^N_{jit}, is stochastic with an unknown distribution G, which is parametrized by the median nominal wage growth, μ^N_{it}, and the dispersion, σ^N_{it}; $G(\mu^N_{it}, \sigma^N_{it})$. Thus, we allow the location and dispersion of the notional wage growth to vary across countries and years, in order to capture the large variation that exists across countries and across time with respect to monetary policy, wage setting, and industry structure. Note that because we control for the location of the country-year distribution, we also capture the fact 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 theoretical 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 1,908 observations from 86 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.[6]

It is a strong assumption to impose the same structural form across countries and years. Yet, almost all studies in the literature use this assumption,

[6] As expected, the 86 country-years differ from the rest of the sample by having a higher mean rate of inflation (11.4 versus 5.6 percent). The use of country-year samples with median wage growth in the upper quartiles is clearly arbitrary. However, as shown by Holden and Wulfsberg (2008), the results are robust to variations in this assumption. Table B1 in the online appendices shows which country-year samples are represented in the notional distribution.

because otherwise it would not be possible to test for the existence of DNWR. However, because the assumption must be thought of as a crude approximation, it is crucial to undertake extensive robustness checks. In the online appendices, 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 1,908 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, $(P75_{it} - P35_{it})$,

$$x_s \equiv \left(\frac{\Delta w_{jit} - \mu_{it}}{P75_{it} - P35_{it}} \right), \qquad s = 1, \ldots, 1{,}908, \tag{2}$$

where subscript s runs over all j, i, and t in the 86 country-year samples. We use the inter-percentile range between the 75th and 35th percentiles as our measure of dispersion, following Nickell and Quintini (2003), to avoid DNWR affecting the measure of dispersion. The calculated x_s should thus be thought of as observations from the standardized underlying two-parametric distribution $X \sim G(0, 1)$.

In the second step, 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,

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

Thus, we have constructed 604 notional country-year distributions $Z_{it} \sim G(\mu_{it}, P75_{it} - P35_{it})$, each defined by $S = 1,908$ wage-change observations $z_s^{it} \equiv x_s(P75_{it} - P35_{it}) + \mu_{it}$. 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 the 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 86 country-year samples with high wage growth. The left panel of Figure 6 displays the underlying distribution of wage changes, with a zero median and inter-percentile range of unity. Compared with the normal distribution (not specified in the diagram), our underlying distribution has a greater peak and fatter tails. Furthermore, it is skewed with the mean at −2.3 percent.

The right panel of Figure 6 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, because the notional distribution is based on the shape of the underlying distribution.

Because we are looking for the possible existence of floors for wage growth, we look for a deficit of wage changes below floors in the range

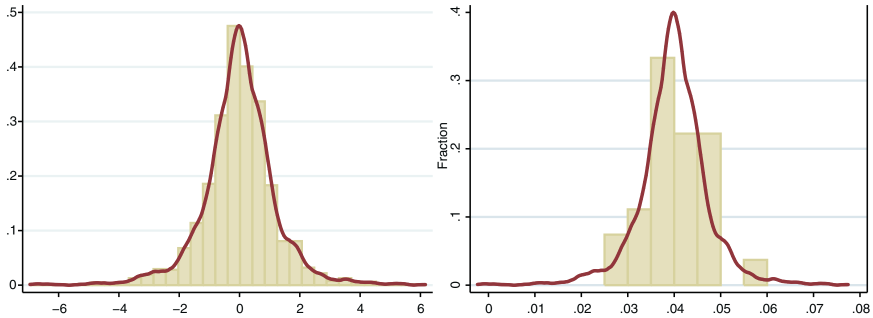

Fig. 6. Left: histogram and kernel density of the normalized underlying distribution of wage changes. Right: histogram of observed wage changes and the notional wage change distribution for Germany in 1987.

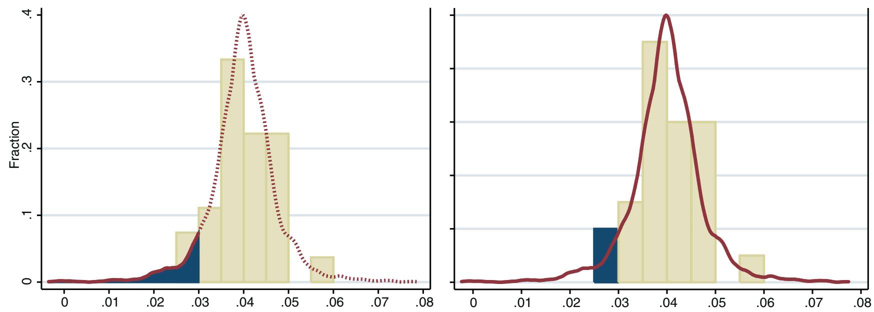

Fig. 7. Illustration of the notional incidence rate below 3 percent (left) and the empirical incidence rate below 3 percent (right) for German wage changes in 1987.

from −5 to 7 percent. For each floor $k \in \{-5, -4.5, -4, \ldots, 7\}$ 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, k, the incidence rate of notional wage cuts is given by

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

and is illustrated by the dark area at the 3 percent floor in the left panel of Figure 7 for Germany in 1987. Likewise, the empirical incidence rate is

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

where S_{it} is the number of observed industries in country-year it. The empirical incidence rate at 3 percent for Germany in 1987 is illustrated by the dark bin in the right panel of Figure 7. The deficit of observed wage changes below floor k relative to the notional distribution (i.e., the fraction of wage changes prevented, $FWCP(k)$) is calculated as

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

For Germany in 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.

The estimates of FWCP in single country-years will be imprecise, because the small number of industries in each country-year sample introduce stochastic disturbances to the country-year specific variables. Thus, we present estimates of the average FWCP for regions and periods, which are likely to be more precise.

In order to test the statistical significance of our estimates of DNWR, we use the simulation method of Holden and Wulfsberg (2008). Under the null hypothesis of no DNWR, the notional country-year specific incidence rate $\tilde{q}(k)_{it}$ is a measure of the probability that a wage change observation in that country-year is below the wage floor k. Specifically, for each country-year it, we draw S_{it} times from a binomial distribution with the country-specific notional probability $\tilde{q}(k)_{it}$. We then compare the number of simulated notional wage changes below k, $\hat{Y}(k)$, with the total number of observed wage changes below k in the corresponding empirical distribution, $Y(k)$. For example, to test for a wage 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 5,000 times, and count the number of times where we simulate more notional wage changes below k than we observe for each floor, denoted $\#[\hat{Y}(k) > Y(k)]$. The null hypothesis is rejected with a level of significance at 5 percent if fewer than 5 percent of the simulations yield more wage changes below the floor than the corresponding empirical number, that is, if $1 - \#[\hat{Y}(k) > Y(k)]/5{,}000 \leq 0.05$.

The method is very powerful for detecting a possible difference between the empirical and notional wage change distribution. However, the construction of the notional wage change distribution involves a potentially substantial downward bias in the estimated DNWR. In a country-year with such 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 are looking 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 for the wage change 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 online appendices shows the 35th percentile wage growth in each country-year sample.

This approach removes some but not all of the downward bias in our estimates, because it might still be the case that DNWR at higher growth rates pushes up the 35th percentile. However, to remove 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.[7]

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

IV. Estimates of DNWR

Figure 8 presents 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 the symbol × in the figure. However, because some of the estimates are based on very few observations, we illustrate by connecting lines the estimates where the number of notional wage changes is at least 1 percent of the relevant population. Figure 8 shows extensive DNWR at a wide range of wage growth rates for the first three decades. The FWCP curves are downward sloping, as predicted by the bargaining model in Section II and illustrated

[7] The test procedure implies that all industry wage changes are treated symmetrically, because they are drawn from the same country-year specific distribution. It would have been possible to allow for variation across industries in the probability of a wage cut. This would have involved less dispersion in the distribution for the number of notional wage cuts in each country-year. This is most easily seen in the extreme case where the probability of a wage cut is either zero or one, and the distribution for the number of wage cuts would collapse to a mass point. Thus, by neglecting variation across industries, our method allows for maximum dispersion in the notional distribution. This biases the procedure against us, making it less likely that any DNWR found will be statistically significant. However, the quantitative importance of this aspect is probably limited; Messina *et al.* (2010) have explored the variation in downward wage rigidity across sectors, and have found that the variation across countries is clearly more important.

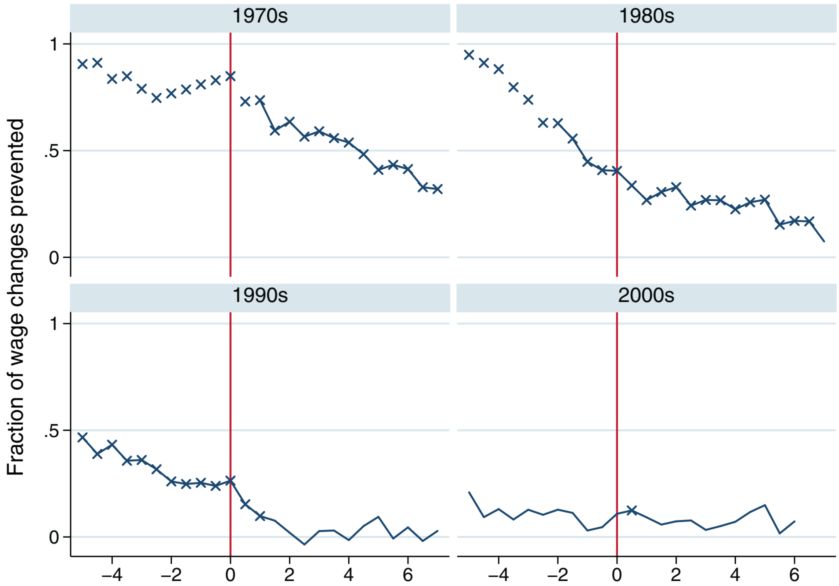

Fig. 8. Fraction of wage changes prevented at different floor levels by periods. Significant estimates at the 5 percent level are marked by a ×. Estimates where the number of notional wage changes are at least 1 percent of the relevant sample are connected with lines.

in the right panel of Figure 4. Intuitively, some of the wage changes far below the floor are pushed up just below the floor, at those points reducing the deficit of wage changes. We also observe a downward shift in the FWCP curve over time, consistent with the predicted effect of lower mean wage growth associated with lower inflation (see Figure 5).[8] The location of the rigidity tends to move to the left over time, which in the bargaining model is associated with a reduction in κ. As suggested in Section II, a lower κ could reflect a change to more flexible remuneration schemes, strengthening the bargaining position of the employer during a holdout.

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

[8] The average inflation rate in the sample was 10.6 percent in the 1970s, 8.5 percent in the 1980s, 3.2 percent in the 1990s, and 2.4 percent in the 2000s.

find statistically significant DNWR at the 0.5 percent floor, with a FWCP equal to 0.12.[9]

While the downward left movement of the FWCP over time indicates less downward wage rigidity, there is also an opposing effect that increases the economic importance of DNWR. As inflation falls over time, the wage change distribution moves to the left, and more industries 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 at 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 online Appendix C). In the 2000s, 9 percent of the wage change observations were negative.

To investigate the cross-sectional variation in DNWR, we group the 19 OECD 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); South (Italy, Greece, Portugal, and Spain). Each of these four regions largely consists of countries with similar labor-market institutions. Because Figure 8 shows large variation over time, we also distinguish between decades for the regions in Figure 9.

Overall, the scenario is fairly similar across the regions. In the 1970s, we find significant DNWR in the form of a floor for nominal wage growth at 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 region, with FWCPs from 0.2 (South) to 0.3–0.5 (Nordic). This result could reflect the fact that countries in these regions have stronger unions and/or stronger employment protection legislation compared to most other countries, particularly in the Anglo region.

The weaker evidence of DNWR in the 2000s is consistent with the conjecture by Gordon (1996) that DNWR will weaken over time in periods

[9] Note that a positive FWCP at, say, 6 percent means that there are fewer empirical than notional wage changes below 6 percent. Nevertheless, there might be more empirical than notional wage changes in the interval from 4 to 6 percent, because some of the notional wage changes below 4 percent are pushed up between 4 and 6 percent. See online Appendix D for details.

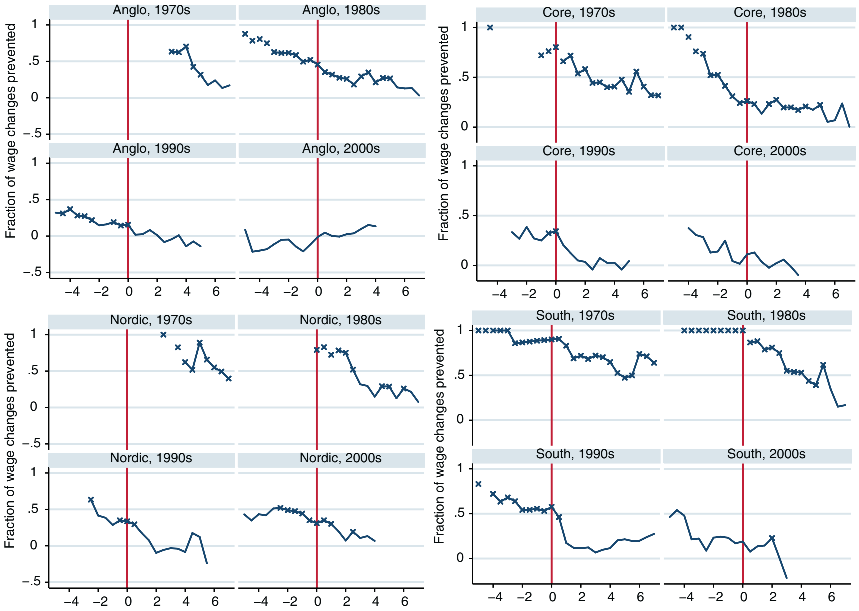

Fig. 9. Fraction of wage changes prevented at different floor levels for the regions by period. Significant estimates at the 5 percent level are marked by a ×. Estimates where the number of notional wage changes is at least 1 percent of the relevant population are connected with lines.

with low inflation. However, caution is warranted, because the downward bias in our estimates is likely to be stronger in country-years with low inflation, and previous studies have shown that DNWR also prevails in recessions (e.g., Bewley, 1999; Fehr and Gotte, 2005). Note that even in the 2000s there is evidence of DNWR in the Nordic and South countries, in spite of these countries experiencing a long period with low inflation.

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 Christofides and Li, 2005; Bauer *et al.*, 2007; Dickens *et al.*, 2007). Furthermore, Holden and Wulfsberg (2009) have found support for some DRWR in the same industry data. In view of the problem of distinguishing between real and nominal downward rigidity, we would also expect that some of the rigidity that we find might, in fact, be caused by real rigidity. However, the nominal floors we consider in the present paper are generally considerably below the inflation rate in the associated country-year. Generally, we also find a somewhat higher FWCP for downward nominal rigidity than has previously been found for downward real rigidity. Thus, we view the floors that we identify as chiefly the result of nominal lower bounds on the wage change process.

V. Effect of Unemployment, Inflation, and Institutions

From the above theoretical model, we would expect the prevalence of DNWR to depend on economic variables, such as inflation (possibly in a non-linear way) and unemployment, as well as on institutional variables, such as the strictness of the employment protection legislation (EPL) and union density. Other institutional variables, such as centralization and coordination of wage setting, can also 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, in order to test whether these variables are related to DNWR as measured by the FWCP. We also control for the level of wage growth.

Technically, we undertake Poisson regressions where the number of observed wage changes below floor k in each country-year sample, $Y(k)_{it}$, depends on the average number of simulated wage cuts for each country-year sample, $\hat{Y}(k)_{it}$, and the explanatory variables mentioned above, $\mathbf{x}_{it}$. A Poisson regression seems appropriate, because $Y(k)_{it}$ 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

$$f[Y(k)_{it} = y(k)_{it} \mid \hat{Y}(k)_{it}, \mathbf{x}_{it}] = \frac{\mathrm{e}^{-\lambda_{it}} \lambda_{it}^{y(k)_{it}}}{y(k)_{it}!}. \tag{7}$$

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

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

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

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

where ε_{it} is an error term.

In the first three columns of Table 1, we present pooled estimates of equation (9) by 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 $\mathrm{E}(Y_{it} \mid \hat{Y}_{it}, \mathbf{x}_{it}) = \mathrm{Var}(Y_{it} \mid \hat{Y}_{it}, \mathbf{x}_{it}) = \lambda_{it}$, an implicit assumption by the Poisson distribution, is accepted easily. As seen from Figure 8, we allow the slope of the FWCP curves to vary between periods by including the

Table 1. *Pooled regressions*

Dependent variable	$[1 - FWCP(\phi)]$			Empirical incidence rate		
$k\times$ 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)
$k\times$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)
$k\times$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)
$k\times$ 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	−0.066**	−0.015	−0.074**	−0.400***	−0.203	−0.251***
protection	(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)
Centralization	−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

Notes: Robust standard errors clustered by country are given 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.

interaction of the floor level with a period specific dummy (denoted as $k\times$ *period*). The positive coefficient on the interaction variables corresponds to the downward slope of the FWCP curves in Figure 8 (note that the dependent variable is $1 - FWCP$, i.e., the fraction of wage cuts realized). For example, with time dummies, the slope estimate for the 1970s is 0.153, which implies that the fraction of wage changes realized decreases by a factor of 0.217 when the floor falls 10 percentage points from 6 to −4 percent ($\exp[0.153(-10)] = 0.217$). Because $1 - FWCP(6) = 0.59$, this estimate predicts that $1 - FWCP(-4) = 0.13$ or $FWCP(-4) = 0.87$, which is close to the observed $FWCP(-4) = 0.84$. Note, however, that Figure 8 displays a bivariate relationship while the estimates are obtained from a multivariate model.

We find that unemployment has a negative effect on the FWCP when country dummies are included, but this does not hold for 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 5. Under high inflation, nominal wage cuts are usually small, and a high proportion are prevented by DNWR. In contrast, when inflation is low, the nominal wage change distribution is further to the left in the diagram, implying that the notional nominal wage cuts are larger. DNWR works to reduce the size of the cuts, yet nevertheless a larger fraction of them are realized. Likewise, high unemployment reduces workers' alternative income in the wage-bargaining process, thus pushing the wage change distribution to the left and reducing the FWCP.

We also find strong correlation with institutional variables. More strict EPL, higher union density, and more centralized wage setting are positively correlated with DNWR, and are thus associated with an upward shift in the FWCP curves. However, the effects are only significant in some of the specifications. For example, the effect of EPL is not significant when we include country dummies, presumably because of the limited time variation in the EPL variable. The effects of EPL and union density are consistent with the results of Holden and Wulfsberg (2008), and also consistent with the theoretical arguments of 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).

The fact that centralization leads to more DNWR is consistent with previous findings that centralized 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 coordinated wage setting is about ensuring overall wage moderation without necessarily affecting relative wages. Thus, countries with both centralized and coordinated wage setting have about the same DNWR as countries where wage setting is neither centralized nor coordinated.[10]

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 $\hat{Y}_{it}$ as in equation (8)):

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

Because $\lambda_{it} = \mathrm{E}\,[Y(k)_{it} \mid S_{it}, \mathbf{x}_{it}]$, we obtain

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

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

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.[11]

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.

VI. Concluding Remarks

We explore the existence of floors on nominal wage growth using industry data for 19 OECD countries, between 1973 and 2006. For the 1970s and 1980s, we find evidence of floors on nominal wage growth in OECD countries at all rates from −2 to 6 percent. Thus, there were significantly fewer nominal wage changes below these growth rates than one would have expected if wage setting had been entirely flexible. This result applies to all four regions we consider, that is, the Anglo (native English-speaking) countries, the Core European countries, the Nordic countries, and the Southern 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), might have contributed to the creation of persistent inflation in this period. 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 1980s, 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 leading to lower inflation 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 social and economic costs in the form of higher unemployment from anti-inflation policies. Potentially, a desire to avoid inflicting these costs might have been part of the reason why policymakers in most OECD countries failed to pursue a sufficiently tight monetary

[11] 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 $\chi^2(11{,}721) = 12{,}566.57$. Including a Gamma-distributed error term, ε_{it}, allows the variance-to-mean ratios of Y_{it} to be larger than unity.

policy in the 1970s. This reasoning could also explain why the costs in terms of higher unemployment were so severe when policy was finally tightened in the 1980s and 1990s.

The existence of these floors to nominal wage growth reflected institutional features in the wage-setting process, 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, because wage setters would set high nominal wage increases to reach their target real wages, but high expected inflation 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 zero percent floor on nominal wage growth, 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 for a floor 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 (i.e., 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 that these countries are currently facing (European Commission, 2010). DNWR would make it difficult to escape from a position with weak international competitiveness. During the deep crisis, we have nevertheless observed extensive wage cuts in Greece, but less in the other Southern European countries. Note, however, that our bargaining model implies that DNWR will push up the real wage, even when there are nominal wage cuts, in the sense that the wage cuts would have been larger, and consequently the real wage lower, if wages had been flexible with no DNWR. 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 might reflect institutional changes such as lower union density in many countries. Stronger pressure on wages arising from increasing globalization, or changes in wage-setting systems, as more flexible pay systems are implemented, might also have provided firms with more flexibility to reduce wages. However, our finding that DNWR is less significant in the 2000s could also be because our method is less able to detect DNWR in an era of low inflation.

We find that DNWR depends on institutional and economic variables, with more prevalent DNWR when EPL is strict and union density is high. Lower inflation leads to a reduction on the fraction of counterfactual wage changes that are prevented by DNWR, consistent with our theoretical bargaining model.

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

Appendix A: Proof of Proposition 1

The equilibrium outcome is 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 takes 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. In step 3, we have the Rubinstein (1982) bargaining game, where the players' threat points are determined in the preceding steps. Binmore *et al.* (1986) have shown 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*

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

where $\kappa > 0$ *if and only if* $\varepsilon < \tau/(\eta-1)$*,* 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

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

which solves for $W^N = (\eta-1)/(\eta-2)R$.

Now consider the situation if the firm has rejected the union's offer in step 2b. The firm can ensure a pay-off

$$\gamma (W^N/P)^{1-\eta} = \gamma \left(\frac{\eta - 1}{\eta - 2} R/P \right)^{1-\eta}$$

by initiating a lockout, and the pay-off $(W^H/P)^{1-\eta} = [(1+\kappa)W_0/P]^{1-\eta}$ by letting a holdout prevail. Clearly, the firm will choose the alternative that gives the higher pay-off. Thus, it will initiate a lockout if and only if

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

Equation (A4) is based on the conventional assumption that no player will initiate a work stoppage if this will result in the same pay-off as a holdout. Equivalently,

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

In step 2a, the union will offer the highest wage that the firm will accept. The firm will accept the union's offer if it can obtain at least as high a pay-off as it can obtain from rejecting the offer. From the analysis above, this is given by $\min[\ell R, (1+\kappa)W_0]$ (given in nominal terms to simplify the notation).

Next, consider the situation if the union has rejected the firm's offer in step 1b. If the union initiates a strike, this will lead to a subsequent agreement on W^N (see above), giving the union a pay-off of

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

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

In step 1, the firm will offer the lowest wage that the union will accept. The union will accept any offer that gives it at least the same pay-off that 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 exists 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 pay-off from initiating a lockout than from initiating a holdout. Thus, lockout threats are credible. To avoid a costly lockout, in step 2, the union will offer the firm a wage $W = \ell R$, which the firm will accept – and the firm might 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, because this yields it a lower pay-off.

If $\ell R \leq (1+\kappa)W_0 \leq sR$, neither player can 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 a higher pay-off 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 the union will accept. ■

Proof of Lemma 1

If holdout threats are in use from step 3 onwards, 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 equation (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 will be given by the first-order condition

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

which simplifies to

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

where $(1+k) = W^H/W_0$. Multiplying equation (A9) with W_0 and rearranging, we obtain

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

Because $f''(k) = (1-\tau)\eta(\eta-1)(1+k)^{\eta-2} > 0$ 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 equation (A10) in $[-1, \infty)$. Observe that equation (A10) is equivalent to equation (A2), implying that κ is also the solution to equation (A9).

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

The relationship between κ and τ, ε and η follows from the implicit differentiation of equation (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

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

■

Implications for wage growth

(i) If strike threats prevailed in the previous year, $W_0 = sR_0$. If $\Delta r > \kappa$, it follows that $sR \equiv (1+\Delta r)sR_0 > (1+\kappa)W_0$, implying from equation (1) that $W = sR$ and thus $\Delta w = \Delta r$. Holdout threats prevail if $\Delta r^L(\gamma) \le \Delta r \le \kappa$, where Δr^L is given by $(1+\kappa)W_0 = \ell R \equiv \ell R_0(1+r^L)$. Substituting out for $W_0 = sR_0$, s and ℓ, we obtain $\Delta r^L(\gamma) = (1+\kappa)\gamma^{\eta/(\eta-1)} - 1$. Thus, $\Delta w = \kappa$ for $\Delta r^L(\gamma) \le \Delta r \le \kappa$. Lockout threats prevail if $\Delta r < \Delta r^L(\gamma)$, in which case $\Delta w = \Delta r + \kappa - \Delta r^L$.

(ii) If lockout threats prevailed in the previous year, $W_0 = \ell R_0$. From equation (1), it follows that holdout threats prevail if $sR \le (1+$

$\kappa)W_0 \leq \ell R$, and then $\Delta w = \kappa$. This condition is equivalent to $\kappa \leq \Delta r \leq \Delta r^S(\gamma)$, where $\Delta r^S(\gamma)$ is given by $(1+\kappa)W_0 = sR \equiv [1+\Delta r^S(\gamma)]sR_0$. Substituting out for $W_0 = \ell R_0$, s, and ℓ, we obtain $\Delta r^S(\gamma) = (1+\kappa)\gamma^{(-\eta)/(\eta-1)} - 1$. Thus, lockout threats continue to prevail if $\Delta r < \kappa$, in which case $\Delta w = \Delta r$. If $\Delta r > \Delta r^S(\gamma)$, strike threats prevail and $\Delta w = \Delta r + \kappa - \Delta r^S$.

(iii) If holdout threats prevailed in the previous year, $W_0 = [sR_0, \ell R_0]$. If wages were at the lower bound in the previous year, $W_0 = sR_0$, the situation would be exactly the same as if strike threats had prevailed during the previous year. Thus, holdout threats would prevail if $\Delta r^L(\gamma) \leq \Delta r \leq \kappa$, and strike threats would prevail if $\Delta r > \kappa$. If W_0 were slightly higher than sR_0, holdout threats would prevail also for Δr slightly above κ, which would mean that the horizontal line in the left panel of Figure 3, indicating a holdout, would move to the right in the figure. If W_0 were higher, the thick lines indicating the outcome would move even further to the right, and for $W_0 = \ell R_0$, the solution would correspond to the lockout case.

Supporting Information

The following supporting information can be found in the online version of this article at the publisher's web site.

Online Appendices.

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