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Evaluating the Economic Significance of Downward Nominal Wage Rigidity*

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Abstract

This paper formalizes and assesses empirically the implications of widely observed evidence for downward nominal wage rigidity (DNWR). First, we formulate and solve an explicit model of wage-setting in the presence of worker resistance to nominal wage cuts – something that has previously been considered intractable. In particular, we show that this resistance renders wage increases (partially) irreversible. Second, using this model, we can explain why previous estimates of the macroeconomic effects of DNWR have been so weak despite remarkably robust microeconomic evidence. In particular, we show that previous studies have neglected the possibility that DNWR can lead to a compression of wage *increases* as well as decreases. Thus, the literature may have been overstating the increase in labor costs due to DNWR.

Using micro-data for the US and Great Britain, we find robust evidence in support of the predictions of the model. In the light of this evidence, we conclude that increased wage pressure due to DNWR may not be as large as previously envisaged, but that the data is nevertheless consistent with a model in which workers resist nominal wage cuts.

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1 Introduction

The existence of nominal rigidities is a cornerstone of macroeconomic theory. Such rigidities act as the key theoretical motivation for the existence of a trade-off between inflation and unemployment in the form of the Phillips curve, and are thus of critical importance to the conduct and efficacy of macroeconomic policy.

Perhaps the most widespread evidence for nominal rigidity comes from a burgeoning literature on the dynamics of wages using micro–data. In particular, this research details some striking characteristics of the distribution of nominal wage changes at the individual level. These include the existence of a mass point at zero nominal wage change and an asymmetry in the form of a deficit of nominal wage cuts, which are taken together as evidence for downward nominal wage rigidity (henceforth DNWR). Such evidence has been found in numerous datasets spanning a vast number of developed economies (for a survey see Kramarz, 2001, and the references therein). However, a number of issues remain unresolved in the light of this research.

An important question relates to evidence for the expected macroeconomic effects of DNWR. In particular, a number of studies have shown that the above results predict the existence of a *convex, long run* Phillips curve (Akerlof, Dickens & Perry, 1996). Intuitively, low inflation implies that reductions in real labor costs can only be effected through nominal wage cuts. If firms are prevented from cutting nominal wages, then their only recourse is to layoff workers, leading to increased unemployment. Thus, when inflation is low, increased inflation can relax the constraint of DNWR on wage-setting for a significant fraction of firms, and thereby reduce unemployment. This result has been of particular interest in recent years due to the adoption of inflation targeting by many monetary authorities. In particular, the existence of a long-run Phillips curve implies that implementing a low inflation target could result in a persistent increase in unemployment.

Much of the research on DNWR addresses precisely this issue. A typical reference is the analysis of Card & Hyslop (1997) for the US. Their micro-level analysis finds strong evidence that nominal wage cuts are restricted when inflation is low, and they conclude that the existence of DNWR leads to an increase in average real wage growth of up to 1% per annum. Card & Hyslop then assess

whether the predictions of this micro-level evidence are corroborated by evidence at a higher level of aggregation. In contrast to their micro-level results, Card & Hyslop's state-level results are much weaker. In particular, whilst they find some evidence for the existence of a Phillips curve trade-off, they obtain estimates that are too imprecise to conclude that this trade-off is stronger in periods of low inflation[1]. Moreover, informal observation of the recent incidence of low inflation together with low unemployment in the US and UK confirms the weakness of this prediction at the most basic level. Thus, there exists a puzzle: if the micro-level evidence for DNWR is so robust, why is the analogous macro-level evidence so fragile?

We argue that we can make progress in resolving this issue via a more careful consideration of the theoretical underpinnings of DNWR. In particular, we present a model of DNWR informed by recent evidence that wage-setters and negotiators are reluctant to cut the nominal wages of workers (see Bewley, 1999, and the survey in Howitt, 2002). In particular, by interviewing over 300 managers, pay professionals, labor leaders etc., Bewley finds that the most common explanation provided for this reluctance is the belief that nominal wage cuts damage worker morale. Moreover, there is additional evidence that agents are subject to money illusion (Shafir, Diamond & Tversky, 1997). In particular, these studies show that agents in different economic settings exhibit significant aversion to nominal losses. A typical finding is that respondents believe it much more acceptable to receive a 5% nominal wage increase when inflation is 12%, than a 7% wage cut when there is no inflation (Kahneman, Knetsch & Thaler, 1986). This is corroborated by Genesove & Mayer (2001) who find evidence from real-estate data that condominium owners were reluctant to sell at a price below that they originally paid, even though they were typically moving locally, and hence were buying in the same market. Thus, an aversion to nominal loss applied to wage cuts can provide a key to explaining the existence of DNWR.

The need for an explicit model of wage-setting in the presence of worker resistance to wage cuts has been noted in the previous literature on money illusion, as well as by labor economists studying the distribution of wage changes:

[1] Weak macroeconomic effects have also been found by Lebow, Saks & Wilson (1999) for the US, and by Nickell & Quintini (2003) and Smith (2004) for the UK. Indeed, Lebow, Saks & Wilson coined the term "micro-macro puzzle" for the observed tension between micro- and macro-level estimates.

"Plausibly, the relationship [between wages and effort] is not continuous: there is a discontinuity coming from nominal wage cuts.... A central issue is how to model such a discontinuity." Shafir, Diamond & Tversky (1997), p.371.

"[I]t is surprising to us that there is no rigorous treatment in the literature of how forward looking firms should set wages when it is costly to cut nominal wages." Altonji & Devereux (2000), p.423 note 7.

We address these issues and show that a key insight into the implications of these behavioral models is that nominal wage increases become partially *irreversible*. In particular, consider a firm that raises the wage today, but reverses the wage increase by cutting the wage by an equal amount tomorrow. When workers resist wage cuts, the net effect on productivity will be negative: today's wage increase will raise productivity, but tomorrow's wage cut will reduce productivity by a greater amount. Thus, reversals of wage increases are costly to firms. In this sense we can think of there being an asymmetric adjustment cost to changing nominal wages.[2]

The solution to this "behavioral" model equips us with a number of predictions that can potentially reconcile the two strands of evidence mentioned above. We show that a key limitation in the previous empirical literature is that it assumes (implicitly or otherwise) that the existence of DNWR has no effect on the upper tail of the wage change distribution. In particular, this is a key identifying assumption in Card & Hyslop (1997), which leads them to use the observed upper tail of the distribution of wage changes to infer the properties of the lower tail in the absence of DNWR. The predictions of our model show that this is misguided[3]. In particular, the upper tail of wage changes will be compressed for two related reasons. First, we show that firms may actively reduce the nominal wage paid when they increase the wage relative to a "counterfactual" world without DNWR – what we will term "*active compression*". In the behavioral model, this results because

[2] Models of adjustment costs have been widely studied in the investment and labor demand literatures, typically in the form of continuous time Brownian models. In contrast, we formulate and solve our model of partial irreversibility in discrete time (Bentolila & Bertola, 1990, and Abel & Eberly, 1996, are closest in spirit to the model analyzed here). This is helpful for two reasons. First, since data are reported in discrete intervals, this method allows us to align theoretical and empirical concepts more naturally. Moreover, many wage contracts are renegotiated on an annual basis, which is more consistent with a discrete-time setup.

[3] This is not to say that Card & Hyslop (1997) is any more subject to this criticism than other previous empirical work on DNWR. Rather, it is the clarity of the statement of identifying assumptions in that paper that allows a particularly clean point of contrast with the implications of the model and results of this paper.

raising the wage today increases the likelihood of having to cut the wage, at a cost, in the future. Second, the model shows that, even if firms do not actively compress nominal wage increases, the upper tail of the wage change distribution will still be compressed relative to the counterfactual with no DNWR. This is because DNWR raises the general level of wages in the economy, and thus firms do not have to raise wages as often or as much to obtain their desired wage level. In particular, we show that this process occurs because of a steady state result that average wages and productivity grow at the same constant rate in the long run. We refer to this process as "*latent compression*".

These properties of the model culminate in the prediction that worker resistance to wage cuts should in fact have no effect on aggregate wage growth. Contrary to conventional wisdom, we show that the costs of DNWR are not derived from increases in the cost of labor, but rather from any reduction in worker productivity due to wage cuts. Moreover, we in turn show that any erroneous suggestion that DNWR indeed raises wage growth will imply a substantial overstatement of the true costs of DNWR to firms.

In the light of this, we seek evidence for these predictions using micro-data for the US and Great Britain. We find significant evidence that the upper tail of the wage change distribution exhibits a compression of wage increases that is related to DNWR. In particular, we find that this limits the estimated increase in real wage growth due to DNWR from around 1–1.5% to no more than 0.3%. We show that this is because firms can "save" at least 75% of the increase in wage growth due to restricted wage cuts by reducing nominal wage increases.

As an additional test of the implications of the model of DNWR presented, we show that the model also implies that increased rates of turnover should mitigate the necessity for firms to restrict wage increases. This occurs because higher turnover reduces the probability that a given worker will stay in the firm an additional period, and thus renders the firm more myopic when it sets wages. Thus firms do not need to compress wage increases as a precaution against future costly wage cuts to the same extent. We again find robust evidence for this hypothesis using data from Great Britain. This reinforces the claim that a model of DNWR based on worker resistance to nominal wage cuts is a useful way of understanding the empirical properties of wage setting.

In the light of this evidence, we conclude that the macro effects of DNWR may not be as large as previously envisaged, and thus do not provide such a strong argument against the adoption of a low inflation target. However, the data is nevertheless consistent with a model in which workers resist nominal wage cuts and may thereby imply something fundamental about the nature of human behavior.

The rest of the paper is organized as follows. Section 2 presents an explicit behavioral model of wage-setting in the presence of worker resistance to nominal wage cuts; section 3 fleshes out some of the predictions of these models that we can take to the data; section 4 presents our empirical methodology and the results obtained; section 5 discusses some remaining issues for future work; and section 6 concludes. Where possible, we omit technical details from the main text, and relegate them to the appendices[4].

2 A Model of DNWR based on Worker Resistance to Wage Cuts

In this section we present an explicit model of downward nominal wage rigidity based on the observations detailed in the empirical literatures mentioned above. In particular, we study the optimal nominal wage policies of worker-firm pairs for whom the productivity of the worker (denoted e) depends upon the wage as follows:

$$e = \ln\left(\frac{\omega}{b}\right) + c\ln\left(\frac{W}{W_{-1}}\right)\mathbf{1}^{-} \tag{1}$$

where W is the nominal wage, W_{-1} the lagged nominal wage, $\mathbf{1}^{-}$ an indicator for a nominal wage cut, $\omega \equiv W/P$ the real wage, and b a measure of real unemployment benefits (which we assume to be constant over time). The parameter $c > 0$ varies the productivity cost to the firm of a nominal wage cut.

The motivation for the qualitative features of this effort function is as follows[5]. We assume that

[4] In addition, we omit some of the more straightforward proofs to save space – these are available from the author on request.

[5] The precise parametric form of (1) is chosen primarily for analytical convenience. None of the qualitative results emphasized in what follows depends on the specific parametric form of (1) – the key is that effort is increasing in the wage and kinked around the lagged nominal wage.

worker effort depends positively on the difference between the level of the real wage, ω, and real unemployment benefits, b. This captures the idea that, the higher the worker's real standard of living from being in work relative to unemployment, the harder that worker will work. In addition, we model the productivity loss due to nominal wage cuts by assuming that effort is falling in the *geometric* nominal wage cut. Our reasoning for this is that the most obvious alternative – that it is the absolute value of the cut in the nominal wage that reduces effort – is implausible in the following sense. It implies that a wage cut of a cent will cause the same loss in effort whether last period's nominal wage is \$1 or \$1,000,000. This is clearly extreme, so we employ the more sensible concept that it is the *percentage* cut in the nominal wage that affects effort[6].

The qualitative features of this effort function are illustrated in Figure 1. Clearly, there is a kink at $W = W_{-1}$ reflecting the existence of DNWR. In particular, the marginal productivity loss of a nominal wage cut exceeds the marginal productivity gain of a nominal wage increase:

$$\frac{\partial e/\partial W|_{W\uparrow W_{-1}}}{\partial e/\partial W|_{W\downarrow W_{-1}}} = 1 + c > 1 \tag{2}$$

This characteristic is what makes nominal wage increases (partially) irreversible – a nominal wage increase can only be reversed at an additional marginal cost of c. Clearly, this is the key driving force in the model that we seek to analyze, and the parameter c is what drives this feature of the model.

The effort function, (1), can be interpreted as a very simple way of capturing the basic essence of the motivations for DNWR mentioned in the literature. It is essentially a parametric form of effort functions in the spirit of the fair-wage effort hypothesis expounded by Solow (1979) and Akerlof & Yellen (1988), with an additional term reflecting the impact of nominal wage cuts on effort – as envisaged in the quote from Shafir, Diamond & Tversky (1997) in the introduction. Bewley (1999) also advocates such a characterization[7]:

[6] One may be interested in a specification with a fixed effort cost due to wage cuts, or a more general convexity of effort in wage cuts informed by the literature on loss aversion (Kahneman & Tversky, 1979). Both such specifications would lead employers to cut the wage dramatically if they cut the wage at all. This differs from the model analyzed here in that we would expect to see a "hole" in the density of wage changes to the left of zero. Whilst previous studies have not found strong support for this (Card & Hyslop, 1997), further work may be worthwhile to assess this more formally.

[7] However, such is the intricacy of Bewley's study, he would probably consider (1) a simplification, not least for its

"The only one of the many theories of wage rigidity that seems reasonable is the morale theory of Solow..." Bewley (1999), p.423.

"The [Solow] theory...errs to the extent that it attaches importance to wage levels rather than to the negative impact of wage cuts." Bewley (1999), p.415.

In addition, an effort function with these properties can be derived from a compensating differentials model where worker utility exhibits nominal loss aversion. The basic intuition for this is that, if workers dislike nominal losses and the firm wishes to cut the nominal wage, then the firm must compensate the worker in the form of lower on-the-job effort in order to prevent the worker from quitting. Thus, in this sense, (1) can be considered a reduced form of a model in which workers dislike nominal loss. The goal of this paper is not to highlight the nuances of emphasis – which do indeed exist – between these behavioral foundations, but rather to show that they share a common, theoretically important, qualitative implication as to the nature of a firm's wage-setting choice. This is intended as a *start* towards richer models of these phenomena, and to this end aims to unify rather than to differentiate.

The most comparable previous attempt at explicitly modelling the behavioral foundation to DNWR is that of Akerlof, Dickens & Perry (1996). However, Akerlof et al. present a model in which firms have no operational discretion over wage-setting – wages are given by a wage-setting relationship which firms take as exogenous, and which dictates that nominal wages can never fall. Thus, the implicit assumption in their model is that firms do not cut wages because, if they did, all of their workers would quit. The model presented in this paper differs critically in that firms have a non-trivial wage-setting decision: firms can cut nominal wages if they wish, but it will have a strong adverse effect on productivity at the margin. We argue that this is a more desirable setup. In the first instance, it accords better with the evidence that firms restrict wage cuts due to concerns over *morale within the firm*, rather than because the external labor market dictates it (Bewley, 1999). Secondly, we will see that a model with wage discretion captures an important characteristic of the available data: that wage increases are also compressed when DNWR binds.

neglect of emphasis on morale as distinct from productivity, and of the internal wage structure of firms as a source of wage rigidity. We argue that it is a *useful* simplification as it provides key qualitative insights into the implied dynamics of wage-setting under more nuanced theories of morale.

The Wage Setting Problem

We consider a discrete-time, infinite-horizon model in which price-taking worker-firm pairs choose the nominal wage W_t at each date t to maximize the expected discounted value of profits. For simplicity, we assume that each worker-firm's production function is given by $a \cdot e$, where a is a real technology shock that is idiosyncratic to the worker-firm match, is observed contemporaneously, and acts as the source of uncertainty in the model. Thus, defining $\beta \in [0, 1)$ as the discount factor of the firm, the typical firm's decision problem is given by:

$$\begin{aligned} \max_{\{W_t\}} \quad & E_t \left[\sum_{s=t}^{\infty} \beta^{s-t} \left\{ a_s e_s - \omega_s \right\} \right] \\ \text{where} \quad & e_s = \ln\left(\frac{\omega_s}{b}\right) + c \ln\left(\frac{W_s}{W_{s-1}}\right) \mathbf{1}_s^- \end{aligned} \tag{3}$$

It turns out in what follows that it is convenient to re-express the firm's profit stream in constant date t prices. To this end, we multiply through by P_t, which we define as the competitive price level at date t, and assume that it evolves according to $P_t = (1 + \pi) P_{t-1}$, where π is the rate of inflation. Finally, defining the nominal counterparts, $A_t \equiv P_t a_t$ and $B_t \equiv P_t b$ and substituting for e_t, we obtain the following optimization problem for the firm:

$$\max_{\{W_t\}} E_t \left[\sum_{s=t}^{\infty} \left(\frac{\beta}{1+\pi} \right)^{s-t} \left\{ A_s \left[\ln\left(\frac{W_s}{B_s}\right) + c \ln\left(\frac{W_s}{W_{s-1}}\right) \mathbf{1}_s^- \right] - W_s \right\} \right] \tag{4}$$

We assume that the nominal shock has support $[0, \infty)$ and that its evolution can be described by the cumulative density function $F(A'|A)$. Thus, rewriting the problem in recursive form[8] we have[9]:

$$v(W_{-1}, A) = \max_{W} \left\{ A \left[\ln\left(\frac{W}{B}\right) + c \ln\left(\frac{W}{W_{-1}}\right) \mathbf{1}^- \right] - W + \frac{\beta}{1+\pi} \int v\left(W, A'\right) dF\left(A'|A\right) \right\} \tag{5}$$

[8] We adopt the convention of denoting lagged values by a subscript, $_{-1}$, and forward values by a prime, $'$.

[9] In addition, we make the assumptions that the measure $dF(A'|A)$ satisfies the Feller property and is monotone, so that the mapping (5) preserves continuity of the value function, and monotonicity of the value function in A. A sufficient condition for this is that A is governed by the stochastic difference equation, $A' = g(A, \varepsilon')$, where g is a continuous function, ε' is an i.i.d. innovation (see Stokey & Lucas, 1989, pp.237, 261–262), and $g_A > 0$. We maintain these assumptions throughout the paper.

(5) is the basic problem that we will attempt to solve in what follows[10]. Before we begin, though, we first present an intuitive outline of the type of results we might expect.

2.1 Some Intuition for the Model

As the theory presented in the forthcoming sections can seem analytically complicated, in this section we present the economic intuition for each of the predictions of the model, which we deal with in turn.

First, the model predicts that there will be a spike at zero in the distribution of nominal wage changes across firms. This occurs because of the kink in the objective function at $W = W_{-1}$. In particular, this implies that for each firm there will be a range of values ("region of inaction") for the nominal shock, A, for which it is optimal not to change the nominal wage. Since A is distributed across firms, there will exist a positive fraction of firms each period whose realization of A lies in their region of inaction that will in turn not change their nominal wage.

Second, in the event that a firm does decide to change the nominal wage, the wage change will be actively compressed relative to the case where there is no DNWR. That nominal wage cuts are attenuated is straightforward to explain – as wage cuts involve a discontinuous fall in productivity at the margin, the firm will be less willing to effect them. In particular, some small wage cuts that would have been implemented in the absence of DNWR will instead be implemented as wage freezes. Moreover, larger counterfactual wage cuts will be reduced in magnitude. It is slightly less obvious why nominal wage increases are also attenuated in this way. The reason is that, in an uncertain world, increasing the wage today increases the likelihood that you will have to cut the wage, at a cost, in the future.

A direct implication of this last prediction is that increases in the productivity cost of cutting the nominal wage, c, will accentuate all these effects. That is, a higher productivity cost due to nominal wage cuts will widen the region of inaction, thereby increasing the mass point at zero in

[10] There is an issue that, for sufficiently low values of the wage, effort is potentially negative. However, accounting for such a non-negativity constraint significantly complicates the solution to the model without much gain in relevance. We maintain the assumption that the level of benefits is sufficiently low relative to wages as to allow almost all firms to ignore this constraint.

the distribution of nominal wage changes, and will also render the active compression of nominal wage changes more acute.

An additional, perhaps more fundamental effect that obtains from the model even in the absence of active compression is what we will refer to as "latent compression" of wage increases. This effect captures the idea that an inability on behalf of firms to cut wages will tend to raise the wages that firms inherit from the past. As a result, when raising wages, firms do not have to increase wages by as much or as often in order to achieve their optimal wage. Thus, this process of latent compression works in tandem with the active compression of wage increases outlined above[11].

The final prediction we want to emphasize at this stage is the effect of increased inflation on nominal wage increases. In particular, we find that the active compression of nominal wage increases becomes less pronounced as inflation rises. As explained earlier, this is because the only reason firms restrict wage increases in the model is the prospect of costly *nominal* wage cuts in the future. Since higher inflation reduces the probability of this occurring, firms no longer need to worry as much about increasing the nominal wage.

2.2 The Dynamic Model

In this section we solve the full dynamic optimization problem as stated in (5) above. First we will present the general structure of the solution, and then we will obtain its specific form under additional assumptions as to the distribution of shocks $F(\cdot)$.

The basic structure of the solution to the full dynamic model is as follows. We solve the problem by first taking the first-order condition with respect to W, conditional on $\Delta W \neq 0$:

$$\left(1 + c\mathbf{1}^{-}\right)\frac{A}{W} - 1 + \frac{\beta}{1+\pi}D(W, A) = 0, \quad \text{if } \Delta W \neq 0 \tag{6}$$

where $D(W, A) \equiv \int v_W(W, A')\, dF(A'|A)$ is the marginal effect of the current wage choice on the future profits of the firm. Clearly, a key step in solving for the firm's optimal wage policy involves

[11] Identifying this additional effect is an important benefit of the infinite horizon model studied here. In particular, one might imagine that active compression-type results could be obtained from a "simpler" two-period model. Latent compression will be shown to be an outcome of steady state considerations, which cannot be treated in a two-period context.

characterizing the properties of the function $D(\cdot)$. However, we leave this for the moment and note first that the general structure of the optimal nominal wage policy is as follows:

Proposition 1 *The optimal nominal wage policy in the dynamic model is of the form:*

$$\begin{array}{llll} \textit{If} & A > u(W_{-1}) \equiv A_u, & \Delta W > 0 & \textit{until} \quad W = u^{-1}(A) \\ \textit{If} & A < l(W_{-1}) \equiv A_l, & \Delta W < 0 & \textit{until} \quad W = l^{-1}(A) \\ \textit{If} & A \in [A_l, A_u], & \Delta W = 0 & \textit{or} \quad W = W_{-1} \end{array} \tag{7}$$

where the functions $u(\cdot)$ and $l(\cdot)$ satisfy:

$$\begin{aligned} \frac{u(W)}{W} - 1 + \frac{\beta}{1+\pi} D(W, u(W)) &\equiv 0 \\ (1+c)\frac{l(W)}{W} - 1 + \frac{\beta}{1+\pi} D(W, l(W)) &\equiv 0 \end{aligned} \tag{8}$$

The reasoning for this is very straightforward. In particular, Proposition 1 uses the conditional first-order condition (6) to define the functions $u(\cdot)$ and $l(\cdot)$, as in (8). These functions determine the optimal relationship between the nominal wage, W, and the nominal shock, A, in the event that wages are adjusted up or down respectively. The rest of the result follows from the fact that, by virtue of the continuity and concavity of the firm's objective, (5), the optimal value of W must be a continuous function of A[12].

However, to complete our characterization of the firm's optimal nominal wage policy, we need to establish the functions $u(\cdot)$, and $l(\cdot)$, to which we now turn. In particular, we can see from (8) that, in order to solve for these functions, we require knowledge of the functions $D(W, u(W))$ and $D(W, l(W))$. This is aided by Proposition 2:

Proposition 2 *The function $D(\cdot)$ satisfies:*

$$D(W, A) = \int_{l(W)}^{u(W)} \left[\frac{A'}{W} - 1\right] dF - \int_0^{l(W)} c\frac{A'}{W} dF + \frac{\beta}{1+\pi} \int_{l(W)}^{u(W)} D\left(W, A'\right) dF \tag{9}$$

[12] This follows from the Theorem of the Maximum (see e.g. Stokey & Lucas, 1989, pp. 62-63).

which is a contraction mapping in $D(\cdot)$ over the relevant range, and thus has a unique fixed point over this range.

The intuition for this result is as follows. The first term on the RHS of (9) represents tomorrow's expected within-period marginal benefit, given that W' is set equal to W. To see this, note that the firm will freeze tomorrow's wage if $A' \in [A'_l \equiv l(W), A'_u \equiv u(W)]$, and that in this event a wage level of W today will generate a within-period marginal benefit of $\left[\frac{A'}{W} - 1\right]$. Similarly, the second term on the RHS of (9) represents tomorrow's expected marginal cost, given that the firm cuts the nominal wage tomorrow. Finally, the last term on the RHS of (9) accounts for the fact that, in the event that tomorrow's wage is frozen, the marginal effects of W persist into the future in a *recursive* fashion. It is this recursive property that provides us with the key to determining the function $D(\cdot)$.

For the purposes of the present paper, we use a specific form for $F(\cdot)$. In particular, we imagine that real shocks, a, evolve according to the following geometric random walk:

$$\begin{aligned} \ln a' &= \ln a - \frac{1}{2}\sigma^2 + \varepsilon' \\ \varepsilon' &\sim \mathcal{N}\left(0, \sigma^2\right) \end{aligned} \tag{10}$$

Given that prices are assumed to evolve according to $P' = (1+\pi)P$, we obtain the following process for nominal shocks, A:

$$\ln A' = \ln(1+\pi) + \ln A - \frac{1}{2}\sigma^2 + \varepsilon' \tag{11}$$

Note that this implies that $E(A'|A) = (1+\pi)A$. We can then use this information to determine the full solution as follows. First, we solve for the functions $D(W, u(W))$ and $D(W, l(W))$ using equation (9), via the method of undetermined coefficients. Then, given these, we obtain the solutions for $u(W)$ and $l(W)$ using the equations in (8). Following this method yields Proposition 3:

Proposition 3 *If nominal shocks evolve according to the geometric random walk, (11), the func-*

tions $u(\cdot)$ *and* $l(\cdot)$ *are of the form:*

$$\begin{aligned} u(W) &= u \cdot W \\ l(W) &= l \cdot W \end{aligned} \tag{12}$$

where u *and* l *are given constants that depend upon the parameters of the model,* $\{c, \beta, \pi, \sigma\}$.

Thus, the optimal nominal wage policy takes the following piecewise linear form:

$$\begin{array}{lccccc} \text{If} & A > u \cdot W_{-1} \equiv A_u, & \Delta W > 0 & \text{until} & W = A/u \\ \text{If} & A < l \cdot W_{-1} \equiv A_l, & \Delta W < 0 & \text{until} & W = A/l \\ \text{If} & A \in [A_l, A_u], & \Delta W = 0 & \text{or} & W = W_{-1} \end{array} \tag{13}$$

2.3 Some Special Cases

In order to get a feeling for how the model works, we present solutions to special cases of the above full dynamic model. In particular, we consider two cases: where nominal wage increases are fully reversible ($c = 0$), and the case where nominal increases are partially irreversible ($c > 0$), but where firms are myopic ($\beta = 0$).

The Case where $c = 0$

Note that the assumption that $c = 0$ removes any dynamic considerations from the firm's wage-setting choice by removing the dependence of effort on last period's wage. In this case, it is straightforward to show that the solution for this problem is:

$$W = A \quad \Longrightarrow \quad \Delta W = \Delta A \tag{14}$$

In this case, wage changes fully reflect changes in productivity, and the distribution of nominal wage changes across firms will be exactly the same as the distribution of changes in the nominal shock. We term this result the *counterfactual solution.*

The Case where $\beta = 0$

This is another static special case of (5), but retains the productivity cost of cutting the nominal wage, c. In this case it is straightforward to confirm that the optimal wage policy is the special case of (13) where $u = 1$ and $l = \frac{1}{1+c}$. By comparing this wage policy to the case where $c = 0$, we can see that the firm is taking counterfactual nominal wage cuts in the interval $\left[\frac{W_{-1}}{1+c}, W_{-1}\right]$ and is instead implementing them as wage freezes. Moreover, for all counterfactual wages below $\frac{W_{-1}}{1+c}$, the firm is reducing the magnitude of wage cuts by a factor $\frac{1}{1+c}$. Thus nominal wage cuts are being actively compressed as a result of DNWR.

However, the same is *not* true for nominal wage increases. All counterfactual wage increases are being implemented without alteration. The reason for this is that $\beta = 0$ implies that the firm doesn't care about the future consequences of raising the nominal wage in the current period. We shall see that this is in stark contrast to the general case where we allow $\beta > 0$, to which we turn in the following section. However, it should be noted at this point that even in this simple case the methods of Card & Hyslop (1997) and other previous studies will be potentially biased. Whilst this special case yields no *active* compression of wage increases by firms, there will still be some *latent* compression: since DNWR places upward pressure on the level of wages in the past, the firm does not have to raise wages as frequently to achieve their target wage today.

3 Predictions

3.1 Active Compression

Returning to the more general solution in (13), it can be seen that active compression of wage changes can be related to the parameters u and l. Numerical simulations of the model establish that $u > 1 > l$ and that $1/l > u$.[13] This is precisely in accordance with our original intuition (section 2.1). Since $u > l$ there exists a region of inaction for the nominal shock variable in which it is optimal not to change the nominal wage. Moreover, because $l < 1$ there will be an active compression of nominal wage cuts. This follows directly from the discontinuous fall in effort following a wage cut at the margin. In addition, $u > 1$ means that nominal wage increases will also be actively compressed relative to the counterfactual solution. Recall that the intuition for this is

[13] Unfortunately, due to the analytical complexity of the solution, a formal proof of this result has proved elusive.

that raising the nominal wage today raises the likelihood that the firm will wish to cut the wage, at a cost, in the future. Finally, the fact that $1/l > u$ implies that the active compression of wage increases will not be as strong as that for wage cuts. The reason for this is that the potential costs associated with wage increases are discounted in two ways. First, some discounting derives from the fact that raising the wage may only increase the costs of wage cuts *in the future*. But, in addition to this, the probability that these additional future costs will be realized is less than one, leading to further discounting.

Recall that our main concern is with the characteristics of the nominal wage change distribution. Using (13) it is straightforward to establish the following proposition for the form of the log nominal wage change distribution, *conditional on the lagged wage*:

Proposition 4 *The log nominal wage change density, conditional on the lagged wage, implied by the model of section 2 is given by:*

$$f\left(\Delta \ln W | W_{-1}\right) = \begin{cases} \tilde{f}\left(\Delta \ln W + \ln u | W_{-1}\right) & \text{if } \Delta \ln W > 0 \\ \tilde{F}\left(\ln u | W_{-1}\right) - \tilde{F}\left(\ln l | W_{-1}\right) & \text{if } \Delta \ln W = 0 \\ \tilde{f}\left(\Delta \ln W + \ln l | W_{-1}\right) & \text{if } \Delta \ln W < 0 \end{cases} \tag{15}$$

where $\tilde{F}\left(\cdot | W_{-1}\right)$ *and* $\tilde{f}\left(\cdot | W_{-1}\right)$ *are the c.d.f. and p.d.f. of the counterfactual (no DNWR) conditional log nominal wage change distribution.*

Figure 2 illustrates this result. In particular, it shows that the distribution of log wage cuts is exactly the same as the counterfactual distribution below $\ln l < 0$, just shifted horizontally by an amount $-\ln l > 0$. A similar result obtains for wage increases. The residual density is "piled up" to a mass point at zero wage change. Thus, the effect of worker resistance to wage cuts is to yield a conditional log wage change distribution with dual censoring from above and below relative to the counterfactual[14].

The key prediction that we will test in our empirical work is the effect of the rate of inflation, π, on the compression of nominal wage increases. To this end, figure 3 presents results for the

[14] This censoring result has interesting parallels in the previous empirical literature. Altonji & Devereux (2000) estimate an econometric model similar to (15) except that they neglect the possibility of compression of wage increases.

effect of changes in the rate of inflation on the parameter u. It is clear is that the firm will reduce any active compression of wage increases as inflation rises since u falls as π rises. The intuition for this is that active compression of wage increases occurs only insofar as wage increases raise the likelihood of future costly nominal wage cuts. To see this, note that our special case in which the firm does not care about the future ($\beta = 0$) yielded no active compression of wage increases ($u = 1$). Thus, since higher inflation reduces the likelihood of future costly nominal cuts, the firm no longer needs to worry about raising the nominal wage today. A key related result that we want to emphasize is that, as inflation becomes large, $u \to 1$. That is, high inflation implies that wage increases cease to be actively compressed relative to a counterfactual world without DNWR. Thus, if the behavioral model is correct, we would expect to observe the upper tail of $f(\Delta \ln W|W_{-1})$ becoming more dispersed as inflation rises. This is illustrated in Figure 4. However, this is not the end of the story: the next section shows that there are additional reasons for there to be a compression of wage increases, even in the absence of these effects.

3.2 Latent Compression

All of the above discussion on active compression has been in terms of the nominal wage change distribution *conditional* on the lagged nominal wage. The reason for this is that the lagged wage is taken as given (is part of the state) at the time of setting the current wage, and so all theories will yield *direct* implications on the conditional distribution, $f(\Delta \ln W|W_{-1})$. However, most of the previous empirical literature has concentrated on the properties of the unconditional distribution, $f(\Delta \ln W)$, typically by estimating some measure of the increase in average wage growth due to DNWR, $E(\Delta \ln W|DNWR) - E(\Delta \ln W|no\ DNWR)$, to try to gain an impression of the effect of DNWR on the firms' real labor costs. The following proposition demonstrates that this emphasis in the previous literature may well be misleading:

Proposition 5 *DNWR has no effect on average wage growth in the long run for finite $G \equiv u/l$.*

This result can be interpreted in a number of ways. First, and closest to the form of the proof, note that the optimal wage policy (13) implies that the difference in the *levels* of the log wage with and without DNWR must be bounded (between $-\ln u < 0$ and $-\ln l > 0$). Thus, it follows that

the rates of *growth* of actual and counterfactual log wages cannot be different in the long run, as it would necessarily imply a violation of these bounds.[15]

An alternative interpretation for this result is that it is simply a requirement for the existence of a steady state in which average growth rates are equal. Since productivity shocks grow on average at a constant rate, so must wages grow at that same rate in the long run. Thus, even the model with DNWR must comply with this simple steady state condition in the long run.

How might this result come about? First, our results above indicate that firms may actively compress wage increases as a precaution against future costly wage cuts, thereby limiting the wage growth increasing effects of DNWR. However, this cannot be the whole story – we saw above that the active compression of wage increases will be less than that of wage cuts. In addition, we can find cases in which there will be no active compression of wage increases for which Proposition 5 still applies. So there must be an additional process at work.

Consider the case where $\beta = 0$. Recall that this is the case in which there is no active compression of wage increases as firms are myopic. Figure 5 shows a simulation of the unconditional wage change distribution implied by the behavioral model in this case. We can see from Figure 5 that, contrary to the assumption of previous studies, the upper tail of $f(\Delta \ln W)$ displays a compression in the presence of DNWR. Thus, the upper tail of the wage change distribution is still compressed, *even if firms do not actively compress wage increases.*

This provides an additional insight into the process by which this steady state requirement might be achieved in practice. If wage increases are not actively compressed, this means that when firms increase the wage, they increase it to the counterfactual level, A. However, recall that the existence of DNWR will tend to raise the general level of lagged wages in the economy, as firms will have been constrained in cutting wages in the past. Thus, when firms increase the wage, they do not have to increase it by as much or as often to reach the counterfactual wage level. Thus the upper tail of $f(\Delta \ln W)$ will indeed still be affected by the existence of DNWR – in particular, it will be less dispersed, as seen in Figure 5. We term this additional effect *"latent compression"*.

[15] A similar result has been established independently in the investment literature by Bloom (2000).

3.3 The Costs of DNWR to Firms

Proposition 5 has important implications with respect to the previous empirical literature. First, by not taking into account the compression of wage increases, previous empirical studies could have overstated the increase in wage growth due to DNWR. To see this, consider Figure 6. This shows three simulated wage change distributions derived from the model of section 2. The bold line shows the wage change distribution with DNWR ($c > 0$), whereas the thick dashed line illustrates the true counterfactual wage change density ($c = 0$). In addition, we include a "median symmetric" (hereafter MS) counterfactual density that is derived by imposing symmetry in the upper tail of the distribution with DNWR (according to the method of Card & Hyslop, 1997). It can be clearly seen that, by using the MS counterfactual, we obtain an *overestimate* of the increase in average wage growth due to DNWR when there is a compression of the upper tail. By neglecting this compression, previous studies could have overstated the effects of DNWR on average wage growth.

The question then arises as to how this bias is related to the implied costs of DNWR to firms. Proposition 6 addresses this issue:

Proposition 6 *To a first–order approximation around the frictionless* ($c = 0$) *case,* (i) *the true average reduction in the value of a firm due to DNWR is equal to* $-g \cdot E\left(\Delta \ln A | \Delta \ln A < 0\right) \cdot ACL^*$ *and is entirely driven by reductions in effort following wage cuts; and* (ii) *the overstatement of the increase in the value of real labor costs assuming a MS counterfactual is equal to* $g \cdot \frac{\beta}{1-\beta} \cdot ACL^*$*; where* g *is the increase in average wage growth due to DNWR assuming a MS counterfactual, and* ACL^* *is the average value of real labor costs.*

A number of points are worthy of note in the light of this. First, a corollary of Proposition 6 is that the conventional view that DNWR imposes costs on firms by increasing firms' real labor costs is incorrect in this model. The true impact of DNWR to firms occurs because nominal wage cuts substantially reduce worker effort at the margin, and thereby reduce productivity. In this way, Proposition 6 fundamentally alters the way one should think about the costs imposed on firms from being constrained in their ability to cut wages.

Moreover, Proposition 6 also allows us to compare the magnitude of the overstatement of the

costs of DNWR under the MS method relative to the true costs of DNWR implied by the current model. As an example, if $g = .01$ and $E(\Delta \ln A | \Delta \ln A < 0) = .15$[16], Proposition 6 suggests that the true costs of DNWR are approximately 0.15% of the average value of labor costs. In contrast, taking $\beta = .6$ as an example, the MS method would imply additional costs of around 1.5% of labor costs, a ten–fold overstatement. More generally, the ratio of these is given by $-\frac{\beta}{1-\beta}\left[E(\Delta \ln A | \Delta \ln A < 0)\right]^{-1}$. Figure 7 plots this ratio as a function of β, the firm's discount factor. It can be seen that for even mildly forward–looking firms the implied overstatement is quite severe. The intuition for this is quite simple – the suggestion that DNWR raises the rate of growth of wages implies that forward–looking firms will anticipate an *accumulation* of increased labor costs over time, which can be large.

One might be tempted to argue that Proposition 6 nevertheless states that the true costs of DNWR are indeed dependent on g, the MS estimate of the increase in wage growth due to DNWR, and that therefore the MS results are informative of the costs of DNWR. However, the key point is that g is informative, but not in the sense that previous studies have thought. Importantly, g does not reflect the increase in wage growth due to DNWR; and if it did, it would imply much larger economic costs on firms than is truly the case.

3.4 Turnover Effects

In addition to the above, the model of section 2 can also provide predictions on the effect of turnover on the distribution of wage changes. To see this, imagine that there is now some exogenous probability that a worker will separate from the firm each period, $\delta < 1$. The effect of this is to reduce the firm's real discount factor from β to $\beta\delta$, since there is now a lower probability that the match will survive until next period. As a result, sectors in which turnover is high (high δ) will act more myopically than sectors with low turnover. In other words, high turnover sectors should set wages more like the special case in which $\beta = 0$ (section 2.2), and low turnover sectors should act more like the forward looking firm of section 2.3. It follows that we should expect to see a greater active compression of wage increases in sectors with lower turnover.[17] Moreover, we should

[16] Card & Hyslop (1997) conclude on an estimate of $g = .01$. The value $E(\Delta \ln A | \Delta \ln A < 0) = .15$ broadly corresponds to the average real wage cut under high inflation found in the data used in section 4.

[17] Thanks to Marianne Bertrand for originally suggesting this idea to me.

also expect this effect to be stronger in periods of low inflation: when inflation is high, the firm does not have to worry about the future consequences of current wage decisions, regardless of the probability that a worker will stay at the firm. We will examine these claims in the forthcoming empirical section to which we now turn.

4 Empirical Implementation

4.1 Data

The data used in this analysis are taken from the Current Population Survey (CPS) and the Panel Study of Income Dynamics (PSID) for the US, and the New Earnings Survey Panel Dataset (NESPD) for Great Britain. For all datasets, the relevant wage measure used in this study is the basic hourly wage rate for respondents aged 16 to 65. Since the CPS and PSID are relatively well-known datasets, we only describe them briefly here.

The CPS samples are taken from the Merged Outgoing Rotation Group (MORG) files from 1979 to 2002. We link respondents across consecutive years using a method similar to that advocated by Madrian & Lefgren (1999)[18]. This method yields approximately 25,000 individual annual wage changes each year from 1980–2002, although changes in sampling method yield lower sample sizes in 1985–86 and 1995–96 (see Table 1). Unfortunately, we cannot easily differentiate between job-stayers and changers using the CPS due to a lack of information on job characteristics and tenure[19]. Additional problems arise in the CPS resulting from the survey redesign in 1994. Figure 8 illustrates the dispersion of log wage changes in the CPS over the sample period, as measured by the standard deviation, and the 90-10 and 80-20 percentile differentials. One can clearly detect a significant rise in the dispersion of wage changes starting in 1994[20]. In our ensuing empirical

[18] In particular, first we match individuals according to their personal identifiers, as well as their month of interview. We then employ Madrian & Lefgren's "s|r|a" criterion – i.e. that matched observations must report the same sex and race across years, and that the difference in their age must lie in the interval $[0, 2]$.

[19] Card & Hyslop (1997) attempt to identify job-stayers in the CPS by restricting their analysis to those respondents who do not change occupation year-on-year. We do not make such an attempt as it is complicated by changes in the occupational classification over the period. However, the sample used in this paper displays very similar properties to that of Card & Hyslop.

[20] This is likely to be due in particular to an increase in the fraction of imputed wage observations in the CPS for 1994 onwards. However, it is difficult to simply deleted such imputed observations from the analysis due to large changes in the accuracy of the CPS imputation flags over the period – see Hirsch & Schumacher (2004).

analysis we attempt to control for this.

The PSID data are taken from the random (not poverty) samples for the years 1971 to 1992. We use data on regular hourly pay rates for household heads to construct individual annual wage changes. We concentrate on the wage changes of job-stayers[21] by excluding workers with tenure of strictly less than 12 months[22], and additionally remove respondents who report that they live in a foreign country, and top-coded wage data. Our PSID sample provides us with much smaller samples than those from the CPS, with approximately 1,300–2,200 individual wage changes each year over the sample period (see Table 1).

Finally, the NESPD is an individual level panel which is collected in April of each year running from 1975 through to 2001 for Great Britain[23]. It is a 1% sample of British income tax-paying workers with a National Insurance (Social Security) number that ends in a given pair of digits. The wage measure used is the gross hourly earnings, excluding overtime, of job-stayers whose pay is unaffected by absence. Table 1 provides summary statistics for the NESPD sample. An important observation to make is that the statistics for the level of real wage changes in 1977 are vastly lower than in all other periods in the NESPD. In particular, the fraction of respondents reporting a real wage cut was 77.64% in 1977, but was never below 52% in any other year in the sample period (see Table 1). The reason for this is that the UK government of the time instituted an incomes policy in order to try to curb high inflation. In particular, these policies were remarkably successful in containing wage inflation in late 1976 to early 1977 as a result of the cooperation of the unions (see Cairncross, 1995, pp. 220–221). Despite this, however, retail price inflation remained high, thereby leading to the significant real wage losses that we observe in our data. As a result of this, we treat the 1977 data as an outlier throughout the rest of our analysis.

Since the descriptive properties of DNWR in all of these datasets have been well-explored in

[21] It should be noted that tenure in the PSID refers to the time spent with the same *employer*, except for the years 1979–80 when it refers to the time spent in the same *position*.

[22] A selection issue arises when excluding job-changers. In particular, previous research has shown that "displaced" workers often accept significant reductions in earnings on re-employment (see Jacobson, LaLonde & Sullivan, 1993). Thus, by concentrating on job-stayers, our results might overstate the true extent of DNWR. However, it is also the case that much of the previous literature has focused on job-stayers, so our analysis will be comparable to that of other studies. We leave these empirical issues for future research.

[23] However, much of our analysis requires the use of consistent industry and occupation coding, which we have up to 1999 only.

previous analyses – Card & Hyslop (1997) for the CPS, Kahn (1997) and Altonji & Devereux (2000) for the PSID, and Nickell & Quintini (2003) for the NESPD – we do not seek to provide a full descriptive account of DNWR. For reference, though, Tables 1 and 2 present summary statistics for wage changes and the key variables that will be used in the forthcoming analysis. The primary aim of the current section is rather to assess the validity of the predictions of the model presented in section 2. However, the relative merits of the datasets used are worth mentioning, especially since the NESPD data have not been as widely used to date.

The NESPD data for Great Britain have a number of key advantages for our purposes, especially in comparison with the CPS and PSID samples for the US. The first, and most obvious, is that the NESPD provides us with comparatively very large sample sizes: we obtain sample sizes of 60–80,000 wage change observations each year. The second advantage of the NESPD data is its sample period: from 1975–2001. This is particularly useful for our purposes given that we seek to use variation in the rate of inflation to gauge the impact of DNWR on wage changes, since the UK experienced significant variation in inflation over this period relative to the US. Figure 9 displays the time-series of the leading UK inflation indicator – the Retail Price Index (RPI) – and the CPI-U inflation rate for the US, over the relevant periods. It can be seen that the UK inflation rate varied substantially, with rates over 20% in the 1970s down to below 2% in the 1990s. Inflation in the US, on the other hand, displays much less variation, with rates no higher than 11%.

The final key advantage of the NESPD sample is that measurement error in these data is likely to be less of a problem relative to individually reported data of the CPS and PSID samples. The reason for this is that the NESPD is collected from employers' payroll records, thereby leaving less scope for error due to imperfect memory etc. (see Nickell & Quintini, 2003, for more on this)[24]. This is important because the existence of measurement error in hourly wages has been shown in previous empirical studies to act as a key impediment to inferring the extent of DNWR. As emphasized throughout this analysis, the existence of a spike at zero in the distribution of nominal wage changes is a key characteristic of DNWR. Classical measurement error in wages and hours

[24] Indeed validation studies of leading panel datasets have used matched data from employer surveys to assess the extent of measurement error in worker reported earnings data. In particular, Bound & Krueger (1990) and Card & Hyslop (1997) both seek to assess the importance of measurement error in the CPS via this method.

data would yield an *understatement* of the spike by rendering true wage freezes to be observed as (small) wage changes (Akerlof et al., 1996). In contrast, previous studies have also stressed that individuals may *round* their reported wages yielding an *overstatement* of the extent of DNWR as small true wage changes are reported as wage freezes (Smith, 2000). Whilst some existing studies have attempted to circumvent this by explicitly modelling measurement error, or using data from payroll records of individual establishments (Altonji & Devereux, 1999; Fehr & Goette, 2003), the relative accuracy of the NESPD allows us to avoid these difficulties and preserve a more representative sample, and is thus an important virtue in this context[25].

A final note worth making in the context of our datasets is that inflation stayed at persistently low levels in the US and UK from 1992 onwards, with an average inflation rate of 2.56% for the US 1992–2002 and 2.69% for the UK 1992–2001. This is important, as a criticism levelled at previous studies of DNWR has been that individuals will get used to receiving nominal wage cuts when inflation has remained low for some time (Gordon, 1996, and Mankiw, 1996). Such a criticism becomes less compelling when the inflation rate has stayed low for the 9–10 years observed in our samples for the CPS and the NESPD.

4.2 Does DNWR Increase Aggregate Wage Growth?

In order to test our hypotheses, we need a way of modelling empirically the wage change distribution, $f(\Delta \ln W)$. In what follows, we will focus on the analogous real wage change distribution counterparts to these[26]. In order to motivate our preferred method, let us begin by considering some naive approaches. First, we might think of simply looking at the differences between the wage change distributions in high inflation periods and low inflation periods to see if the predictions of section 3 are confirmed at this basic level. To this end, figures 10(a) and 11(a) present estimates of the density of log real wage changes for periods with different inflation rates using the PSID for the

[25] It should be noted that hourly earnings in the NESPD are derived from dividing weekly earnings by weekly hours, thereby potentially exacerbating any underlying measurement error. However, Nickell & Quintini (2003) have compared the accuracy of hourly wage changes in the NESPD with those obtained from a sample whose payslip was checked in the British Household Panel Study and found remarkably similar properties in both datasets.

[26] Note that this does not alter any substantive aspects of the analysis, since these are exactly the same shaped distributions, just shifted to the left by a constant, $\Delta \ln (W/P) = \Delta \ln W - \Delta \ln P \cong \Delta \ln W - \pi$ where π is the rate of inflation. However, focusing on real wage changes does allow greater ease of comparison across years with different inflation rates.

US, and the NESPD for Britain (the redesign of the CPS renders this a less useful exercise for the CPS data). Notice that lower inflation leads to a compression of the lower and, more importantly for our purposes, the *upper* tail of the wage change distribution, precisely in accordance with the predictions of section 4.1[27].

However, one could argue that at least some of the observed differences were due to changes in other variables that affect wage changes. For example, there have been changes in the industrial, age, gender, regional etc. compositions of the workforce in both the US and Britain over these time periods. So, we should control for factors such as these before attributing any differences to DNWR. To address this, we introduce a set of micro-level control variables for each dataset, summarized in Table 2. In particular, we control for changes in micro-level variables by re-weighting the observed wage change distributions according to the method of DiNardo, Fortin & Lemieux (1996) (henceforth DFL)[28]. To do this, we first define a "base year", T – for all datasets this will be the final sample year – and re-weight each year's observed wage change distribution to obtain an estimate of what the wage change distribution would have looked like if the distribution of micro-level characteristics were identical to that at date T. In particular, if we define the log wage change as Δw, micro-level characteristics as x, and the year of the relevant x distribution as t_x, we derive:

$$f\left(\Delta w_t; t_x = T\right) = \int f\left(\Delta w|x\right) dF\left(x|t_x = T\right) = \int f\left(\Delta w|x\right) \cdot \psi \cdot dF\left(x|t_x = t\right) \tag{16}$$

for all $t < T$. The key insight of DFL is that this is simply a re-weighted version of the observed date t wage change distribution, with weights ψ given by:

$$\psi = \frac{dF\left(x|t_x = T\right)}{dF\left(x|t_x = t\right)} = \frac{\Pr\left(t_x = T|x\right)}{\Pr\left(t_x = t|x\right)} \cdot \frac{\Pr\left(t_x = t\right)}{\Pr\left(t_x = T\right)} \tag{17}$$

where the second equality follows from Bayes' Rule. The conditional probabilities in (17) can then

[27] It should be noted that the existence of the spike in the lower tail of the real wage change distribution (at approximately minus the rate of inflation) can lead to an overstatement of lower tail compression. However, our emphasis is on the effects on the upper tail, which are not subject to this problem.

[28] An important benefit of the DFL methodology is that it requires few parametric assumptions on the impact of the x variables. Given the intrinsically non-linear character of the wage policy (13), this is especially helpful.

be estimated simply via a probit model.

Figures 10(b) and 11(b) displays density estimates of the DFL re-weighted distribution of log real wage changes for different inflation periods, again for the PSID and NESPD. Again, it can be seen clearly that lower rates of inflation are associated with a compression both of tails of the wage change distribution, in line with the predictions of section 3.

However, even having controlled for such factors, it is still not necessarily legitimate to attribute *all* the residual difference in the wage change distributions to DNWR. Thus we need a way of ensuring that only the variation in wage change distributions that varies *systematically* with DNWR is attributed. To do this, we estimate regressions of the form:

$$P_{n_{rt}} = \beta_{0n} + \beta_{1n} P_{50_{rt}} + \eta_n \boldsymbol{\pi}_t + z'_{rt}\gamma_n + \varepsilon_{n_{rt}} \tag{18}$$

where $P_{n_{rt}}$ is the nth percentile of the real wage change distribution in region r at time t, $\boldsymbol{\pi}_t$ is the rate of inflation at time t and thereby measures the prominence of nominal zero in the distribution of log real wage changes, and z_{rt} is a vector of aggregate controls that could potentially affect the distribution of wage changes. $P_{50_{rt}}$ is included on the RHS of (18) in order to control for changes in the central tendency of the distribution of wage changes. That is, it "re-centres" the distributions over time in order to make them comparable. We estimate (18) by Least Squares, where we weight by the size of the region at each date[29].

The measure of inflation used will be the CPI-U-X1 series for the US, and the April to April log change in the Retail Price Index for Great Britain. The aggregate controls will be as follows. First, we control for any distortion to the wage change distributions caused by peculiarities of the datasets used. So, to control for the effects of survey redesign issues after 1994 in the CPS, we include a dummy variable that takes value one for all years from 1994 onwards when we estimate (18) for the CPS. In addition, to control for the incomes policies implemented in 1977 in the UK, we include a dummy that takes value one for the year 1977 in our NESPD regressions.

[29] Formal quantile regression (Least Absolute Deviation) estimators were also tried with little difference in results. However, such is the computational intensity involved in estimating the correct standard errors for these estimators, we opted for simple OLS instead.

In addition, we control for the absolute change in the rate of inflation. This is motivated by the hypothesis that greater inflation volatility will yield greater dispersion in relative wages regardless of the existence of DNWR (see Groshen & Schweitzer, 1999). We also include both current and lagged regional unemployment rates. This is motivated by the idea that the existence of DNWR might lead to unemployment – indeed, as mentioned before, this is one of the principal reasons for interest in the topic. Since unemployment will lead to workers "leaving" the wage change distribution, it is important to control for any resulting distributional consequences. We also include lagged regional unemployment in accordance with the wage curve hypothesis of Blanchflower & Oswald (1994) that the *level* of wages is empirically associated with the level of unemployment. If this is true, then we would expect the change in unemployment to affect the change in wages, and so we include lagged regional unemployment to control for this possibility.

It should be noted that the empirical method described above is robust to a number of possible concerns. First, the specification is robust to the existence of rigidity in *real* wages. The reason is that real wage rigidity, in its traditional form, will be invariant to inflation by definition. An exception to this is the argument put forward by Akerlof, Dickens & Perry (2001) that real wage rigidity is amplified as inflation rises because it becomes optimal for workers to direct their scarce attention to maintaining their real wage. However, if anything, such a possibility would work *against* the claim of the model in section 2, as it would predict that the upper tail of wage changes would become *more* compressed as inflation rises. If this were the case, any evidence we find for the predictions of section 3 could be interpreted as lower bounds on the true effects. A similar reasoning applies to any concerns one might have about the impact of skill-biased technical change (SBTC). Under SBTC, we might expect that workers obtaining high wage increases early in our samples will obtain even higher wage increases later on as technical change increasingly favors those in skilled sectors. However, since inflation is in practice declining over the sample periods of our data, SBTC would, if anything, work against the predictions of section 3.

Clearly, the coefficients of interest in (18) for the purposes of estimating the effects of DNWR are η_n. In particular, the predictions of section 3.2 indicate that η_n should be negative for low percentiles, and positive for high percentiles. The reasoning is that higher inflation should lead to

an increased dispersion of wage changes, and thereby decrease negative percentiles, and increase positive ones.

Recall that we would like to obtain an estimate of the increase in average wage growth due to DNWR, $\lambda \equiv E(\Delta w|DNWR) - E(\Delta w|no\ DNWR)$. Such an estimate can be obtained using the estimates obtained from regressions of the form (18). In order to use this information to get an estimate of λ, we obtain an estimate of the predicted average wage change when inflation is very low (e.g. 1.3% in 1993 for Britain) and subtract the analogous average wage change when inflation is very high (e.g. 21.8% in 1980 for Britain)[30]:

$$\hat{\lambda} = \hat{E}(\Delta w|\boldsymbol{\pi} = 1.3\%, x, z) - \hat{E}(\Delta w|\boldsymbol{\pi} = 21.8\%, x, z) \tag{19}$$

To obtain these estimates using (18), we can use the fact that the quantiles of a random variable are uniformly distributed. In particular, if we estimate k equi-spaced percentiles of $f(\Delta w)$ then a best guess of the predicted average wage change is:

$$E(\Delta w|\boldsymbol{\pi}, x, z) \approx \frac{1}{2(k-1)} \sum_{i=1}^{k-1} \left(\hat{P}_i + \hat{P}_{i+1} \right) \tag{20}$$

where i is an ascending index of the percentiles, with $i = 1$ indicating the lowest percentile, $i = 2$ the second lowest etc., and the $\hat{P}$s are the predicted values of these percentiles obtained from estimating equation (18).

Since our predicted percentiles allow us to sketch out a discretization of the whole distribution of wage changes, we can also decompose the increase in average wage growth due to DNWR into two components. The first is the increase in average wage growth due to compressed nominal wage cuts, which we refer to as "lower tail losses"; the second is the decrease in average wage growth due to compressed wage increases, "upper tail gains". In practice, we will perform this procedure on 99 estimated wage change percentiles, $\hat{P}_1, \hat{P}_2, ..., \hat{P}_{99}$, for the specification detailed above.

The Effects of Measurement Error

[30] Note that this involves out-of-sample predictions for the US data.

As mentioned previously, the impact of measurement error on our ability to infer the effects of DNWR has received substantial attention in the literature. Whilst we have attempted to mitigate this as a problem by using the relatively clean data in the NESPD for Britain, the question arises as to the effects of measurement error on the methodology detailed above. Proposition 7 answers this question:

Proposition 7 *If measurement error is independent of the rate of inflation, then* (i) *estimates of* η_n *in (18) report attenuated estimates of the corresponding true effects;* (ii) *this attenuation vanishes for sufficiently high and low percentiles; and* (iii) *estimates of* λ *based on (20) will nonetheless remain consistent.*

The intuition for this is quite straightforward[31]. The existence of measurement error will render some true negative wage changes to be observed as positive wage changes (and vice versa). Thus, measurement error will lead to a partial conflation of the effects of inflation on the upper tail with those in the lower tail. However, as we proceed further into the tails of the wage change distribution, the likelihood of measurement error having displaced observations in this way becomes smaller. Thus, such attenuation will disappear for sufficiently high or low percentiles.

Whilst this result does mean that a certain caution should be afforded to our interpretation of the *magnitude* of the estimated coefficients from (18), the main question we are trying to answer is a qualitative one: does lower inflation compress both the upper and lower tails of the wage change distribution? To this end, the above attenuation result will actually *reduce* our ability to observe any such compression, should it exist. Thus, any evidence of compression we might find would be found despite the existence of measurement error, rather than because of it.

The final part of Proposition 7 results from the fact that, by definition, the mean of any random variable can always be expressed as an unweighted average of its percentiles. Since measurement error is assumed independent of the rate of inflation (see Gottschalk, 2004, for evidence that this is empirically the case), it follows that the observed mean wage change at any given rate of inflation

[31] It should be noted, however, that this attenuation result is quite distinct from the traditional attenuation bias resulting from errors in variables when implementing OLS.

will be unaffected by the existence of measurement error. Thus, our estimates of λ based on (20) should also be unaffected by measurement error.

Empirical Results

We estimate (18) in three specifications. First, we simply include controls for the median wage change, P_{50}, and for any dataset peculiarities such as the CPS survey redesign from 1994 onwards and incomes policies of 1977 in the NESPD. We then include controls for the absolute change in the rate of inflation, and for regional current and lagged unemployment rates. Finally, we implement a specification with full controls that estimates (18) using percentiles of the DFL *re-weighted* wage change distributions so we can control for an array of micro-level characteristics as well.

The results from estimating our three specifications of (18) for each dataset are reported in Tables 3–5. First, consider the results obtained for the CPS in Table 3. In all three specifications it can be seen that the estimated impact of inflation is negative for the 20th–30th percentiles, with strongest effects around the 30th percentile; and positive for the 40th–90th percentiles, with strong effects in the 60th–90th percentiles. Thus, these results are in line with the hypothesis that higher inflation reduces the compression of both tails of the wage change distribution. Moreover, we see that the estimated effects of inflation at different points in the distribution are generally significant and fairly stable across all specifications. In addition, Table 3 presents estimates of the lower tail losses and upper tail gains due to DNWR. It can be seen that in all specifications there are substantial savings due to compressed wage increases, some of which even outweigh the costs from compressed wage cuts.

Table 4 reports the analogous estimates for the PSID data. We can see that in all specifications the effect of inflation is negative for the 10th–20th percentiles, and positive for the 40th–90th percentiles. However, here the estimated effects are strongest in the 10th, and particularly the 20th, percentiles in the lower tail in contrast to the CPS results. The differences in the lower tail effects between the CPS and PSID results are likely to reflect the differences in the position of nominal zero in the respective wage change distributions, due to higher rates of inflation in the PSID sample period. In the CPS, nominal zero appears mostly between the 20th and 35th

percentiles, whereas it appears at around the 10th–35th percentile in the PSID sample. Thus, the point at which DNWR binds differs across these two datasets.

The PSID results are broadly as significant as those for the CPS, with both lower and upper tail effects remaining significant, and fairly stable across specifications. In addition, the coefficient estimates in the upper tail are comparable to those obtained in the CPS results, and we again observe that there are large savings from the compression of the upper tail. In particular, we find an estimated increase in average wage growth due to lower tail losses of around $1 - 1.2\%$ which is offset by a reduction in average wage growth due to upper tail compression of $0.9 - 1.1\%$. It should however be noted that for the PSID, and to some extent the CPS data, these estimates are constructed from a number of regressions for which no significant inflation effect was detected. This is likely due to the relative lack of observations and inflation variation in the CPS and PSID compared to the NESPD. Thus, we do not want to place too much stock in the actual quantitative estimates obtained from this dataset. Rather, we consider our estimates of upper tail gains and lower tail losses for the PSID to be instructive of the fact that there is some significant compression of the upper tail of wage changes, and that this compression is of similar significance and magnitude to the compression of the lower tail due to DNWR.

The results for the NESPD data are reported in Table 5. Again we observe that inflation has a negative impact on lower percentiles (10th–40th) and a positive impact on higher percentiles (60th–90th). Moreover, we obtain highly significant estimates for almost all percentiles and in all specifications. As mentioned above, this greater significance in comparison to the results for the PSID and the CPS is likely to be due to the larger sample sizes, more precise wage information, and large variation in inflation in the NESPD. In addition, we again observe substantial upper tail gains due to compression of wage increases relative to lower tail losses, which are more consistent across specifications than those obtained for the CPS and the PSID. In particular, our results suggest that 75–95% of the lower tail losses due to DNWR is saved by restricting wage increases in the upper tail in the NESPD data, and that the increase in average real wage growth due to DNWR is of the order 0.04–0.3% – much lower than results obtained previously.

Together, these results provide strong evidence for the prediction that the upper tail of the

wage change distribution will be less dispersed as a result of DNWR – in all specifications and for all datasets we see that wage increases become more restricted as inflation falls. As a result, by allowing both the upper and lower tails of the wage change distribution to be affected by DNWR, the estimated increase in average wage growth due to DNWR becomes much reduced and closer to zero – precisely in line with the predictions of section 3 and Proposition 5.

4.3 Does Higher Turnover Reduce the Compression of Wage Increases?

In addition to the above, recall that section 3.3 established the claim that higher turnover sectors should act more myopically, will thus feel more at liberty to raise nominal wages, but that such an effect should fade as inflation rises[32]. We test this hypothesis in a manner similar to that employed in section 4.2. First, we define a measure of "turnover" as the fraction of workers that changes jobs each year in a given occupation, region group. In a steady state this should closely match the fraction of workers who separate, and thus correspond to the parameter δ in section 3.3.

To gain an initial impression for whether such effects exist, Figure 12(a) plots density estimates of the nominal wage change distribution for job stayers in high and low turnover (respectively above and below median turnover) occupations using the NESPD data[33]. It can be seen from this simple comparison that low turnover occupations seem to be compressing wage increases much more than high turnover occupations.

However, recall that section 3.3 noted that these effects should be manifested through changes in the *active* compression of wage increases, and thus relate to the properties of the conditional distribution, $f\left(\Delta \ln W | W_{-1}\right)$. To address this issue, Figure 12(b) replicates the exercise but controls for the covariates listed in Table 2, as well as the lagged wage[34], using the DFL technique. This allows us to control for any micro-level factors that may be driving the results, as well to focus

[32] Note that we are making a claim about the effects of turnover on the *distribution* of wage growth rather than mean wage growth. Clearly, there are additional concerns that would relate turnover to mean wage growth such as effects on tenure, match quality etc. These are not however related to compression of the distribution of wage growth emphasized here.

[33] Since tenure is not reported in the CPS, and is subject to changes in definition in the PSID, we concentrate on the NESPD data for this section.

[34] In particular, we use the lagged wage adjusted for inflation and productivity growth, $\tilde{W}_{t-1} \equiv W_{t-1} \cdot \frac{P_{T-1}}{P_{t-1}} \cdot \frac{a_{T-1}}{a_{t-1}}$, where a_s is measured as GDP per hour in year s. This is legitimate provided that DNWR has no impact on either price or productivity growth. Given Proposition 5 and the results of section 4.2 this does not seem unreasonable.

on variation in the distribution of wage changes conditional on the lagged wage. Again, however, we see low turnover occupations compressing wage increases more than high turnover occupations, consistent with the predictions of section 3.3.

To complete the picture, panels (c) and (d) of Figure 12 attempt to assess whether the effect of turnover on the distribution of wage changes varies with the rate of inflation. It can be seen that the compression of wage increases due to lower turnover appears to be stronger in periods of low inflation, as predicted by the model of section 2. Together, then, there appears to be suggestive evidence that the predictions of section 3.3 are supported in the data.

To identify these turnover effects more formally, however, we run Least Squares regressions of the form:

$$P_{n_{ort}} = \alpha_{0n} + \alpha_{1n} P_{50_{ort}} + \theta_n \boldsymbol{\tau}_{ort} + \eta_n \boldsymbol{\pi}_t + \iota_n \left(\boldsymbol{\pi}_t \times \boldsymbol{\tau}_{ort}\right) + z'_{rt}\psi_n + \varepsilon_{n_{ort}} \tag{21}$$

where $P_{n_{ort}}$ now refers to the nth percentile of nominal wage changes for job stayers, re-weighted for micro covariates and the lagged wage, in occupation o, region r, at time t. The variable of interest is $\boldsymbol{\tau}_{ort}$ which denotes the fraction of job changers in an occupation, region, year cell. Under the predictions of section 3.3, we would expect that the coefficients on turnover, θ_n, to be positive, the coefficients on inflation, η_n, to be positive, and the coefficients on the interaction term, ι_n, to be negative, for all positive percentiles of nominal wage changes.

Table 6 summarizes these estimates for the 60–90th percentiles. In the first specification (column 1), we include only basic controls for the median wage change and a dummy for 1977 to control for the incomes policies of that time, and exclude the rate of inflation and its interaction with turnover. It can be clearly seen that turnover has a positive and highly significant impact on the 60–90th percentiles of nominal wage changes. The two additional columns address potential concerns one might have about the simple specification of column (1).

In particular, one concern might be that we would expect sectors with greater DNWR to have greater rates of turnover due to workers being made unemployed more often. In addition, we

would also expect sectors with greater DNWR to exhibit a greater compression of wage increases, and thus create downward pressure on percentiles of wage increases. In this sense there may be an omitted variable – the extent of DNWR – that will lead to bias in the estimates of column (1). In particular, it would imply a *downward* bias to the estimates. Column (2) seeks to assess this possibility by including the current and lagged regional unemployment rates as controls. It can be seen, however, that this actually *reduces* the estimated effects of turnover, which nevertheless remain positive and highly significant. Thus, there does not appear to be strong evidence for an omitted variables problem of this type.

However, we may still be concerned that the measure of turnover is more generally cyclical, and thus potentially correlated with the rate of inflation. Thus we may be worried that we are attributing to turnover the effects due to declining inflation. We may also be concerned that the effects of inflation found in section 4.2 are not robust to the addition of turnover as a control. Finally, we would like to assess whether the effect of turnover on the distribution of wage changes diminishes as inflation rises. To address these concerns, specification (3) includes the rate of inflation and the interaction between inflation and turnover as regressors. It can be seen that introducing these controls does not significantly alter the estimated effects of turnover, and that the coefficients remain positive and highly significant. Moreover, we find that the effect of inflation is robust to the addition of turnover as a control, and remains positive and highly significant for all except the 90th percentile of wage changes.

The coefficients on the interaction between turnover and inflation are less successful, however. In particular, the estimates are positive, but close to zero and highly insignificant for the 60th and 70th percentiles. The estimates for the 80th and 90th percentiles are negative and much larger in magnitude – in accordance with the "behavioral" model's predictions – though the effect is significant only at the 80th percentile. However, these latter estimates imply that a 20% inflation rate will reduce the effect of turnover on the distribution of wage changes by 50–60%.

We thus find robust evidence that increased turnover leads to an increased dispersion of wage increases, that the effects of turnover and inflation are mutually robust, and that there is sugges-

tive evidence that the effects of turnover are reduced as inflation rises, broadly in line with the predictions of section 3.3.

5 Conclusions

This paper has sought to clarify our interpretation of the widespread empirical evidence for DNWR. In particular, it has shown that the conventional view that the costs of DNWR are driven by increases in real labor costs in times of low inflation is misguided. In particular, we have shown as a theoretical issue that DNWR should have no effect of aggregate wage growth in the long run, and that this is achieved because firms compress wage increases as well as wage cuts when DNWR binds. Moreover, we show that a neglect of the compression of wage increases leads to an overstatement of the increase in aggregate wage growth due to DNWR, and thereby a substantial implied overstatement of the costs of DNWR to firms.

Taking these predictions to a range of micro–data on wages from the US and Great Britain we find that firms do indeed compress wage increases as well as wage cuts. Furthermore we find that accounting for the compression of wage increases reduces the estimated increase in aggregate wage growth due to DNWR to be much closer to zero, consistent with the predictions of the model. In this way, the theoretical and empirical results of the paper can help explain the limited evidence for large macroeconomic effects despite robust micro–level evidence for DNWR.

A number of issues arise in the light of the results of this paper. From a theoretical perspective, one may ask whether the findings of section 4 can be explained by models of nominal rigidity other than DNWR. In particular, one may contend that a standard model of menu costs can explain the observed compression of wage increases in times of low inflation. In particular, Sheshinski & Weiss (1977) show that, in a deterministic price-setting model, increased inflation will result in more extreme price increases – as firms increase prices less often to avoid successive payment of menu costs in high inflation environments, when they do increase the price, they will increase it by more. However, if this were the correct model, we would expect to see "holes" either side of zero in the density of nominal wage changes, and moreover that these holes would *widen* as inflation rises.

Whilst previous empirical work has found some evidence for menu cost effects, these effects have only a modest impact on wage changes around zero (Card & Hyslop, 1997), and certainly are not accentuated in times of high inflation. So it is by no means clear that a menu cost model could explain the results presented in section 4.

An additional question that arises is whether the model of DNWR presented here is subject to the Barro (1977) critique that wage rigidity should not be allocational. As presented, the model indeed implies that DNWR imposes costs on firms, and that labor demand is actually reduced when inflation is low. The reason is that, although long–term wage contracts allow the firm to smooth out labor costs over time, the wage is not only a side payment in this model but also a determinant of productivity. However, a number of points should be made at this point. First, it should nevertheless be noted that a key result of this paper is that the costs of DNWR are likely to be not nearly as large as previously envisaged. Second, an issue arises as to whether firms can obviate the costs of DNWR, for instance by replacing incumbent workers with new workers without a wage history. The results of this paper suggest that this option is not universally at firms' disposal. In particular, if firms could operate such a policy DNWR would not be a concern to them, and they would not feel the need to constrain wage increases when inflation is low. An implication is that it is not easy for firms to replace existing workers, perhaps due to costs of searching for, hiring, and training new workers. In the light of this, then, fruitful areas of future research might be to try to quantify the productivity costs of DNWR, and to analyze how these costs vary with the costs of replacing workers.[35]

From an empirical perspective, the results of this paper have important implications with respect to our ability to identify a valid contemporaneous counterfactual wage change distribution. In particular, we have highlighted the problems inherent in previous studies which have sought to obtain identification from the observed distribution of wage increases. A fundamental question arises in the light of this as to how one can obtain estimates of the effects of DNWR with a clean

[35] In addition, the theory presented in section 2 neglected the possibility of DNWR motivated by factors such as the legal requirement in many countries (notably excluding the US) that wage contracts be renegotiated by mutual consent of the firm and the worker (MacLeod & Malcomson, 1993; Holden, 1994). The current paper does not seek to deny the existence of such motivations, but merely to draw out and test the implications of behavioral foundations to DNWR. Indeed, as pointed out in Holden (2004), contract and behavioral motivations may even reinforce one another in explaining DNWR.

causal interpretation. This paper has shown that there are strong theoretical and empirical reasons to conclude that such a question is more difficult to solve than previously envisaged, and therefore that future work on this issue is needed.

In addition, there is a need in the literature on DNWR for an empirical model that can conform well with an explicit theory of wage setting as well as with the structure of the data. In particular, whilst the empirical methods of section 4 allow the data to speak more – by allowing different effects of inflation at different points in the wage change distribution, and by assuming little about the parametric form of counterfactual wage changes – they do not provide us with direct estimates that can be related back to a model of wage setting. The current paper seeks to contribute to this process by showing how one can write down models of wage setting based on worker resistance to wage cuts, and by also providing empirical evidence that can inform future, more complex, models of DNWR. This will hopefully enable the formulation of more realistic structural models of DNWR that can be successfully estimated with meaningful parameter estimates.

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7 Appendix

A Lemmas and Proofs

Lemma 1 The value function defined in (5) has the following properties:

$$\begin{aligned} v_W^- \left(W, A'\right) &= -c\frac{A'}{W} \\ v_W^0 \left(W, A'\right) &= \frac{A'}{W} - 1 + \frac{\beta}{1+\pi} D\left(W, A'\right) \\ v_W^+ \left(W, A'\right) &= 0 \end{aligned} \tag{22}$$

Proof. First, note that standard application of the Envelope Theorem implies:

$$\begin{aligned} v_W^- \left(W, A'\right) &= c\frac{A'}{W} \\ v_W^+ \left(W, A'\right) &= 0 \end{aligned} \tag{23}$$

It is only slightly less obvious what happens when $\Delta W' = 0$, i.e. when the wage is not adjusted. In this case, $W' = W$ and this implies that:

$$v^0 \left(W, A'\right) = A' \ln\left(\frac{W}{B'}\right) - W + \frac{\beta}{1+\pi} \int v\left(W, A''\right) dF\left(A''|A'\right) \tag{24}$$

It therefore follows that:

$$v_W^0 \left(W, A'\right) = \frac{A'}{W} - 1 + \frac{\beta}{1+\pi} \int v_W\left(W, A''\right) dF\left(A''|A'\right) \tag{25}$$

Since, by definition $D\left(W, A'\right) \equiv \int v_W\left(W, A''\right) dF\left(A''|A'\right)$, the statement holds as required. ■

Proof of Proposition 2. First, note that we can re-write the continuation value conditional on each of the three possible continuation regimes:

$$v\left(W, A'\right) = \begin{cases} v^-\left(W, A'\right) & \text{if} \quad A' < A'_l \\ v^0\left(W, A'\right) & \text{if} \quad A' \in [A'_l, A'_u] \\ v^+\left(W, A'\right) & \text{if} \quad A' > A'_u \end{cases} \tag{26}$$

where superscripts $^{-/0/+}$ refer to whether the nominal wage is cut, frozen, or raised tomorrow.

Note also that, due to the recursive nature of the problem:

$$\begin{aligned} A_l \equiv l\left(W_{-1}\right); \qquad & A_u \equiv u\left(W_{-1}\right) \\ \Longrightarrow A_l' \equiv l\left(W\right); \qquad & A_u' \equiv u\left(W\right) \end{aligned} \tag{27}$$

Thus we can write[36]:

$$\int v\left(W,A'\right)dF\left(A'|A\right) = \int_0^{l(W)} v^-\left(W,A'\right)dF + \int_{l(W)}^{u(W)} v^0\left(W,A'\right)dF + \int_{u(W)}^{\infty} v^+\left(W,A'\right)dF \tag{28}$$

Taking derivatives with respect to W and recalling the definition of $D\left(\cdot\right)$, and noting that, since $v\left(W,A'\right)$ is continuous, it must be that $v^-\left(W,l\left(W\right)\right) = v^0\left(W,l\left(W\right)\right)$ and $v^0\left(W,u\left(W\right)\right) = v^+\left(W,u\left(W\right)\right)$ yields:

$$D\left(W,A\right) = \int_0^{l(W)} v_W^-\left(W,A'\right)dF + \int_{l(W)}^{u(W)} v_W^0\left(W,A'\right)dF + \int_{u(W)}^{\infty} v_W^+\left(W,A'\right)dF \tag{29}$$

Finally, using the Envelope conditions in Lemma 1, and substituting into (29) we obtain (9) in the main text:

$$D\left(W,A\right) = \int_{l(W)}^{u(W)} \left[\frac{A'}{W} - 1\right]dF - \int_0^{l(W)} c\frac{A'}{W}dF + \frac{\beta}{1+\pi}\int_{l(W)}^{u(W)} D\left(W,A'\right)dF \equiv \left(\mathbf{C}D\right)\left(W,A\right) \tag{30}$$

To verify that $\mathbf{C}$ is a contraction mapping over the "relevant range" (to be defined shortly), we confirm that Blackwell's sufficient conditions for a contraction hold here (see Stokey & Lucas, 1989, p.54). First, note that any values for (W,A) that render $\mathbf{C}$ unbounded cannot obtain under optimality, since they will necessarily violate the conditional first-order condition, (6). Thus, we can restrict our attention to a subset of values for (W,A) around the optimum for which $\mathbf{C}$ is bounded. This is what we define as the "relevant range". That $\mathbf{C}$ then maps the space of bounded functions into itself over this range holds by definition. Given this, monotonicity and discounting are straightforward to verify. To verify monotonicity, fix $(W,A) = \left(\bar{W},\bar{A}\right)$, and take $\hat{D} \geq D$. Then note that:

$$\begin{aligned} \int_{l(\bar{W})}^{u(\bar{W})} \hat{D}\left(\bar{W},A'\right)dF\left(A'|\bar{A}\right) - & \int_{l(\bar{W})}^{u(\bar{W})} D\left(\bar{W},A'\right)dF\left(A'|\bar{A}\right) \\ = & \int_{l(\bar{W})}^{u(\bar{W})} \left[\hat{D}(\bar{W},A') - D\left(\bar{W},A'\right)\right]dF\left(A'|\bar{A}\right) \geq 0 \end{aligned} \tag{31}$$

Since $\left(\bar{W},\bar{A}\right)$ were arbitrary, it thus follows that $\mathbf{C}$ is monotonic in D. To verify discounting, note that:

$$\begin{aligned} \left[\mathbf{C}\left(D+a\right)\right]\left(W,A\right) &= \left(\mathbf{C}D\right)\left(W,A\right) + \frac{\beta}{1+\pi}a\left[F\left(u\left(W\right)|A\right) - F\left(l\left(W\right)|A\right)\right] \\ &\leq \left(\mathbf{C}D\right)\left(W,A\right) + \frac{\beta}{1+\pi}a \end{aligned} \tag{32}$$

[36] Henceforth, "dF" without further elaboration is to be taken as "$dF\left(A'|A\right)$".

Since we know that $\frac{\beta}{1+\pi} < 1$ it follows that $\mathbf{C}$ is a contraction over the relevant range. It therefore follows from the Contraction Mapping Theorem that $\mathbf{C}$ has a unique fixed point over the relevant range. ■

Proof of Proposition 5. Denote the counterfactual nominal wage at time t as $W_t^* = A_t$. We seek the properties of the difference in average wage growth between the "actual" (with DNWR) and counterfactual cases, which we define as λ:

$$\begin{aligned} \lambda_T &\equiv \frac{1}{T}\sum_{s=t+1}^{t+T} \ln\left(\frac{W_s}{W_{s-1}}\right) - \frac{1}{T}\sum_{s=t+1}^{t+T} \ln\left(\frac{W_s^*}{W_{s-1}^*}\right) \\ &= \frac{1}{T}\left\{\ln\left(\frac{W_{t+T}}{W_{t+T}^*}\right) - \ln\left(\frac{W_t}{W_t^*}\right)\right\} \end{aligned} \tag{33}$$

Then note that, from the optimal wage policy of the firm, (13), it follows that the log–difference between the actual and counterfactual wages must be bounded, $\ln\left(W_t/W_t^*\right) \in [-\ln u, -\ln l]$. Thus:

$$\begin{aligned} \sup \lambda_T &= \frac{1}{T}\left[\ln u - \ln l\right] = \frac{1}{T}\ln G \\ \inf \lambda_T &= \frac{1}{T}\left[\ln l - \ln u\right] = -\frac{1}{T}\ln G \end{aligned} \tag{34}$$

Therefore, for finite G, $\lim_{T\to\infty} \sup \lambda_T = 0 = \lim_{T\to\infty} \inf \lambda_T$. ■

Proof of Proposition 6. (i) Define the increase in wage growth due to DNWR in the median–symmetric (MS) case as:

$$g \equiv \int_{-\infty}^{0} x f(x)\, dx - \int_{-\infty}^{0} x \hat{f}(x)\, dx \tag{35}$$

where f, and $\hat{f}$ are respectively the observed, and MS counterfactual log wage change densities. Further define the true average cost of DNWR as $\mathbb{C} \equiv E(v^* - v)$ where v^* denotes the true frictionless value of a firm. Taking a first order Taylor series approximation around $c = 0$ yields:

$$\mathbb{C} \overset{c=0}{\approx} \left.\frac{d\mathbb{C}}{dc}\right|_{c=0} \cdot c \tag{36}$$

Note first that $d\mathbb{C}/dc|_{c=0} = -\left[\frac{\partial v}{\partial c} + \frac{\partial v}{\partial W}\cdot\frac{\partial W}{\partial c}\right]\Big|_{c=0} = -E\sum_{t=0}^{\infty}\beta^t a_t \Delta \ln A_t \mathbf{1}_t^- = -E\left(\Delta \ln A \mathbf{1}^-\right)\frac{E(a)}{1-\beta}$, since $\partial v/\partial W|_{c=0} = 0$, and where the second equality follows from the stationarity of, and the independence of increments to, the shock process. It follows from this that the costs of DNWR are entirely driven by reductions in effort following wage cuts. Next, note that by using the definition of G in (55) one can show that $\ln G \overset{c=0}{\approx} c$.

The final step for part (i) is to show that $\ln G \approx g/\Pr\left(\Delta \ln A < 0\right)$. Recall that in the wage cut regime, $\Delta \ln W = \ln\left(A/W_{-1}\right) - \ln l = \ln\left(A/A_{-s}\right) - \ln l + \ln \kappa_{-s}$, since $W_{-1} = A_{-s}/\kappa_{-s}$, where s is the number of periods since the last wage change, and $\kappa_{-s} \in \{l, u\}$ depending on whether the last wage change was negative or positive. Then note that if one were to assume that observed wage increases were unaffected by DNWR, one implicitly assumes that the counterfactual wage change is given by $\Delta \ln \hat{A} = \ln\left(A/W_{-1}\right) - \ln u = \ln\left(A/A_{-s}\right) - \ln u + \ln \kappa_{-s}$. If follows that, for a given s,

$\hat{f}(x) = f(x + \ln G)$. Thus, for a given s,

$$g = \int_{-\infty}^{0} x f(x)\, dx - \int_{-\infty}^{\ln G} (x - \ln G) f(x)\, dx \overset{\ln G=0}{\approx} \ln G \cdot \Pr(\Delta \ln A < 0) \tag{37}$$

which is independent of s. It follows that $g \approx \ln G \cdot \Pr(\Delta \ln A < 0)$ as required. Piecing these components together we obtain:

$$\mathbb{C} \overset{c=0}{\approx} -g \cdot E(\Delta \ln A | \Delta \ln A < 0) \cdot \frac{E(a)}{1-\beta} \tag{38}$$

(*ii*) Previous studies asserted that DNWR is costly because it raises the cost of labor, and thereby sought to estimate these increased costs. In this way, they identify the costs of DNWR to be equal to:

$$\mathbb{W} \equiv E \sum_{t=0}^{\infty} \beta^t (\omega_t - \omega_t^*) = E \sum_{t=0}^{\infty} \beta^t \prod_{n=1}^{t} (1 + g_n)\, \omega_0 - \frac{E(a)}{1-\beta} \tag{39}$$

where g_n is the (random) growth rate of real wages in period n. From the above definition of g, $E(g_n) \equiv g$. Together with the stationarity of the increments to the shock process, this implies that $E(g_n | g_{n-1}, g_{n-2}, ...) = g$; i.e. that the gs are independent of each other. Thus:

$$\mathbb{W} = \frac{E(\omega_0)}{1 - \beta - \beta g} - \frac{E(a)}{1-\beta} \tag{40}$$

According to Proposition 5, of course, the "true" value of g is zero. Thus the overstatement of the wage costs of DNWR under the MS methodology is given by:

$$\frac{\beta g}{1 - \beta - \beta g} \frac{E(\omega_0)}{1-\beta} \overset{g=0}{\approx} g \cdot \frac{\beta}{1-\beta} \cdot \frac{E(a)}{1-\beta} \tag{41}$$

as required. ■

Proof of Proposition 7. (*i*) By definition, $F(P_n) \equiv n/100$. Totally differentiating yields:

$$\frac{\partial P_n}{\partial \pi} = \left. \frac{-\partial F / \partial \pi}{f(\Delta w)} \right|_{\Delta w = P_n} \tag{42}$$

Now, defining measurement error as ϵ with cdf $G(\epsilon)$, it follows that the cdf of observed wage changes, Δw, is given by $F(\Delta w) = \int F^*(\Delta w^* | \epsilon)\, dG(\epsilon)$ where F^* is the cdf of true wage changes, $\Delta w^* \equiv \Delta w - \epsilon$. It follows that:

$$\frac{\partial F}{\partial \pi} = \int f^*(\Delta w^* | \epsilon) \frac{\partial(\Delta w^*)}{\partial \pi} dG(\epsilon) \tag{43}$$

where f^* is the pdf of true wage changes. To make the exposition most stark, consider the following case:

$$\frac{\partial(\Delta w^*)}{\partial \pi} = \upsilon \cdot \mathbf{1}(\Delta w^* > 0) + \lambda \cdot \mathbf{1}(\Delta w^* < 0) \tag{44}$$

where $\mathbf{1}(\cdot)$ is the indicator function, and $\upsilon > 0$ and $\lambda < 0$. In this case, it is straightforward to

show that:

$$\frac{\partial P_n}{\partial \pi} = \upsilon \cdot \varphi(n) + \lambda \cdot [1 - \varphi(n)] \tag{45}$$

where $\varphi(n) = \left.\frac{\int_0^\infty f(\Delta w|\Delta w^*)dF(\Delta w^*)}{\int_{-\infty}^\infty f(\Delta w|\Delta w^*)dF(\Delta w^*)}\right|_{\Delta w = P_n} \in (0,1)$. Thus, the estimates of (18) report a weighted average of upper and lower tail effects, and are thus attenuated.

(*ii*) It is also true, however, that as $n \to 100$ (resp. 0) then $\varphi \to 1$ (resp. 0). Therefore, this attenuation vanishes for both sufficiently high and low percentiles.

(*iii*) Note that, since ϵ is independent of π, then $E(\Delta w|\pi) = E(\Delta w^*|\pi)$: i.e. the mean conditional wage change is unaffected by measurement error. To complete the proof, note that, by definition:

$$E(\Delta w|\pi) \equiv \frac{1}{100}\int P_n dn = E(\Delta w^*|\pi) \tag{46}$$

■

B Technical Details of Proposition 3

The following lemma will turn out to be useful in what follows:

Lemma 2 If $\ln x \sim N\left(\mu, \sigma^2\right)$ then it follows that:

$$\int_{\underline{x}}^{\overline{x}} x dF(x) = \exp\left[\mu + \frac{1}{2}\sigma^2\right]\left\{\Phi\left[\frac{\ln \overline{x} - \mu}{\sigma} - \sigma\right] - \Phi\left[\frac{\ln \underline{x} - \mu}{\sigma} - \sigma\right]\right\} \tag{47}$$

where $\Phi(\cdot)$ is the c.d.f. of the standard Normal.

Proof. Since x is log-Normally distributed, the p.d.f. of x is given by $f(x) = \frac{1}{\sigma x}\phi\left(\frac{\ln x - \mu}{\sigma}\right)$, where $\phi(\cdot)$ is the p.d.f. of the standard Normal. It follows that:

$$\int_{\underline{x}}^{\overline{x}} x dF(x) = \int_{\underline{x}}^{\overline{x}} x \frac{1}{\sigma x\sqrt{2\pi}} \exp\left[-\frac{1}{2}\left(\frac{\ln x - \mu}{\sigma}\right)^2\right] dx \tag{48}$$

Defining $z \equiv \ln x - \mu \Longrightarrow dx = \exp(\mu + z)\, dz$, we obtain:

$$\int_{\underline{x}}^{\overline{x}} x dF(x) = \int_{\ln \underline{x} - \mu}^{\ln \overline{x} - \mu} \frac{1}{\sigma\sqrt{2\pi}} \exp\left[\mu + z - \frac{1}{2\sigma^2} z^2\right] dz \tag{49}$$

Completing the square for the term in brackets and substituting back into the former expression:

$$\begin{aligned} \int_{\underline{x}}^{\overline{x}} x dF(x) &= \int_{\ln \underline{x} - \mu}^{\ln \overline{x} - \mu} \frac{1}{\sigma\sqrt{2\pi}} \exp\left[\mu + \frac{1}{2}\sigma^2 - \frac{1}{2}\left(\frac{z - \sigma^2}{\sigma}\right)^2\right] dz \\ &= \exp\left[\mu + \frac{1}{2}\sigma^2\right]\left\{\Phi\left[\frac{\ln \overline{x} - \mu}{\sigma} - \sigma\right] - \Phi\left[\frac{\ln \underline{x} - \mu}{\sigma} - \sigma\right]\right\} \end{aligned} \tag{50}$$

as required. ■

B.1 Obtaining the functions $D(W, u(W))$ and $D(W, l(W))$

We proceed by using the method of undetermined coefficients. We conjecture that $D(W, A)$ is of the form:

$$D(W, A) = \alpha_1 \frac{A}{W} + \alpha_2 \tag{51}$$

and verify that this will indeed be the case for $A = u(W)$ or $l(W)$, using Lemma 2 to solve out the integrals in (9). Following this method yields[37]:

$$\begin{aligned} D(W, u(W)) &= (1+\pi)\frac{u(W)}{W}\left[\frac{\kappa_1 - (1+c)\Phi_1}{1-\beta(\kappa_1 - \Phi_1)}\right] - \frac{\kappa_2 - \Phi_2}{1 - \frac{\beta}{1+\pi}(\kappa_2 - \Phi_2)} \\ D(W, l(W)) &= (1+\pi)\frac{l(W)}{W}\left[\frac{\Phi_3 - (1+c)\kappa_1}{1-\beta(\Phi_3 - \kappa_1)}\right] - \frac{\Phi_4 - \kappa_2}{1 - \frac{\beta}{1+\pi}(\Phi_4 - \kappa_2)} \end{aligned} \tag{52}$$

where:

$$\begin{array}{ll} \Phi_1 = \Phi\left[\frac{1}{\sigma}\left(-\ln G - \ln(1+\pi) + \frac{1}{2}\sigma^2\right) - \sigma\right] & \Phi_2 = \Phi\left[\frac{1}{\sigma}\left(-\ln G - \ln(1+\pi) + \frac{1}{2}\sigma^2\right)\right] \\ \Phi_3 = \Phi\left[\frac{1}{\sigma}\left(\ln G - \ln(1+\pi) + \frac{1}{2}\sigma^2\right) - \sigma\right] & \Phi_4 = \Phi\left[\frac{1}{\sigma}\left(\ln G - \ln(1+\pi) + \frac{1}{2}\sigma^2\right)\right] \\ \kappa_1 = \Phi\left[\frac{1}{\sigma}\left(-\ln(1+\pi) + \frac{1}{2}\sigma^2\right) - \sigma\right] & \kappa_2 = \Phi\left[\frac{1}{\sigma}\left(-\ln(1+\pi) + \frac{1}{2}\sigma^2\right)\right] \end{array} \tag{53}$$

and we define $G \equiv \frac{u(W)}{l(W)}$, the geometric gap between the two trigger values for A.

B.2 Obtaining the functions $u(W)$ and $l(W)$

It is now straightforward to solve for the functions $u(W)$ and $l(W)$ by substituting the above (52) into the equations (8) to obtain after some algebra:

$$\begin{aligned} u(W) &= \left[\frac{1-\beta(\kappa_1 - \Phi_1)}{1 - \frac{\beta}{1+\pi}(\kappa_2 - \Phi_2)} \cdot \frac{1}{1 - c\beta\Phi_1}\right] \cdot W \\ l(W) &= \left[\frac{1-\beta(\Phi_3 - \kappa_1)}{1 - \frac{\beta}{1+\pi}(\Phi_4 - \kappa_2)} \cdot \frac{1}{1 + c - c\beta\Phi_3}\right] \cdot W \end{aligned} \tag{54}$$

These two equations clearly depend on $G \equiv \frac{u(W)}{l(W)}$, which is unknown so far. However, we can determine G using our expressions for $u(W)$ and $l(W)$ above:

$$\frac{u(W)}{l(W)} \equiv G = \frac{1+\pi-\beta(\Phi_4 - \kappa_2)}{1+\pi-\beta(\kappa_2 - \Phi_2)} \cdot \frac{1-\beta(\kappa_1 - \Phi_1)}{1-\beta(\Phi_3 - \kappa_1)} \cdot \frac{1+c-c\beta\Phi_3}{1-c\beta\Phi_1} \equiv T(G) \tag{55}$$

Note that all the terms on the RHS of this equation are functions of G, and not of W. Obtaining the relevant value of G requires solving for the fixed point(s) of the mapping defined by this equation. Given the relevant value of G, this implies that the Φ_is, $i = 1, ..., 4$, will be given constants, as will the coefficients on W in (54), and it follows that $u(W) = u \cdot W$ and $l(W) = l \cdot W$ as stated in the main text.

[37] Technical details of this derivation are available on request from the author.

B.3 Properties of the Map $T(G)$

Simulations of the mapping $T(G)$ in (55) reveal that, whilst there always exists at least one fixed point for $T(G)$, there is not, in general, a unique fixed point. Thus, in the case where there exists more than one fixed point, we need a criterion for identifying which fixed point value of G maximizes the value function, which is provided by the following proposition:

Proposition 8 *Where there exist multiple fixed points for the mapping $T(G)$, the wage policy that maximizes the value function is that associated with $G^1 \equiv \min\{G : G = T(G)\}$.*

Proof. Define the multiple fixed points of $T(G)$ as $G^1 < G^2 < G^3 < ...$, and the associated value functions as $v^1, v^2, v^3, ...$ We claim that the following must be true:

$$v^1 \geq v^2 \geq v^3 \geq ... \tag{56}$$

To see this, note first that a higher value of G only serves to restrict the firm's choice of W by widening the region in which wages are not changed. In particular, under a lower value of G, the firm can always choose a W arbitrarily close to W_{-1}, and hence replicate the wage policy under a higher G, if it wishes. In general, though, the firm can do better than this under lower values of G. Thus the statement must hold. ■

Figure 1: The Effort Function

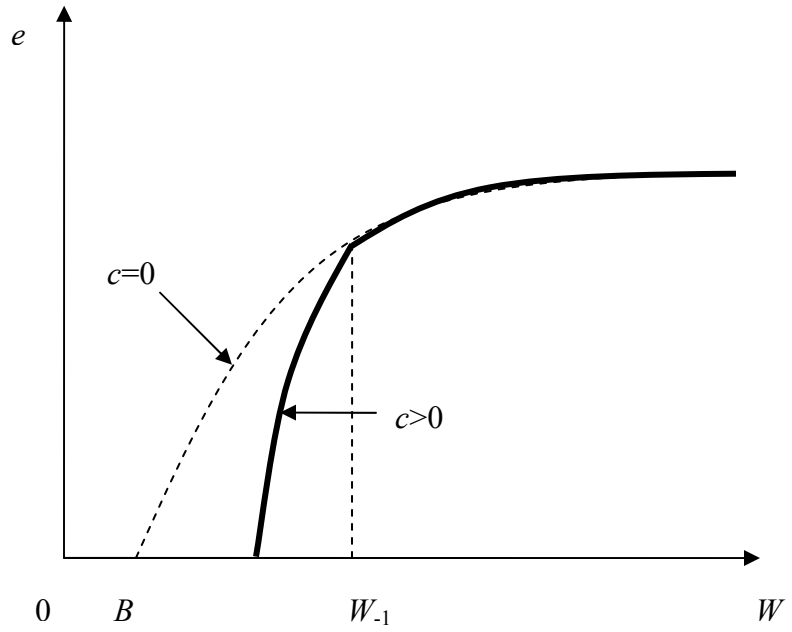
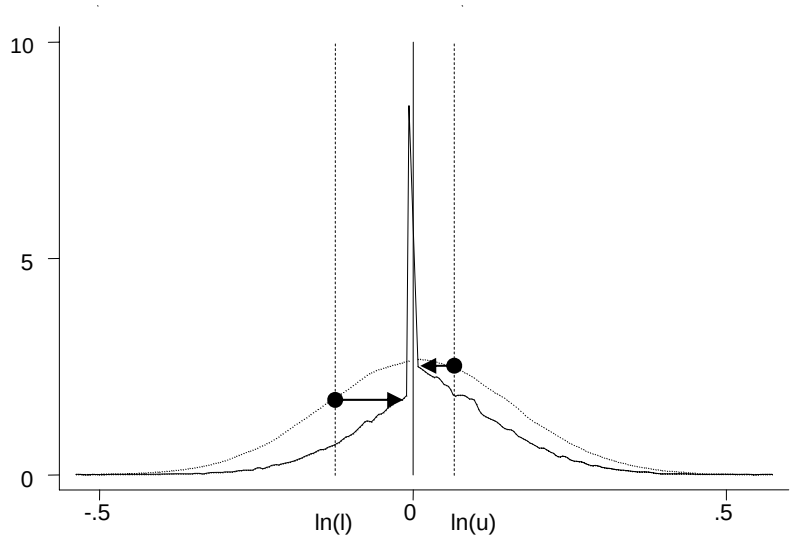


Figure 2: $f\left(\Delta \ln W \mid W_{-1}\right)$ **implied by Theory**



_____: Actual Density; ----------: Counterfactual Density

Figure 3: Properties of the Optimal Wage Policy Parameters

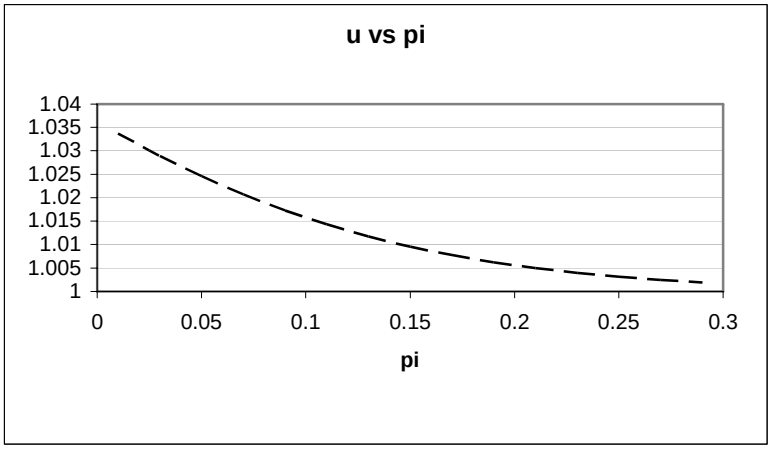


Figure 4: Theoretical $f(\Delta \ln W \mid W_{-1})$ for Different Rates of Inflation

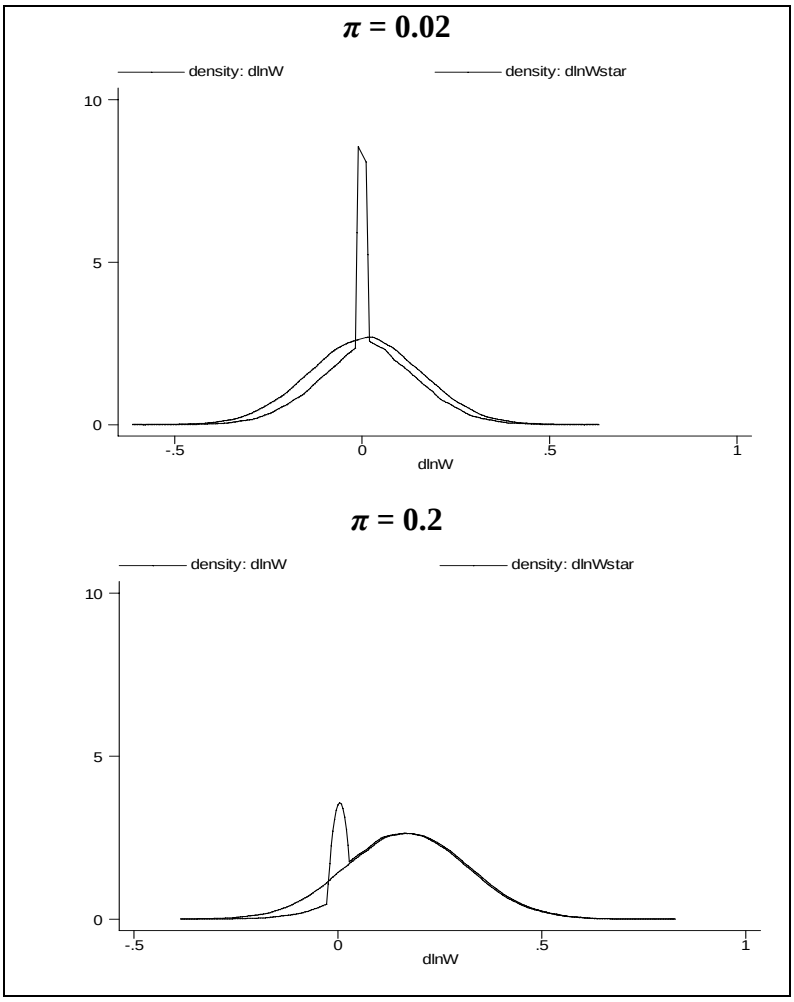
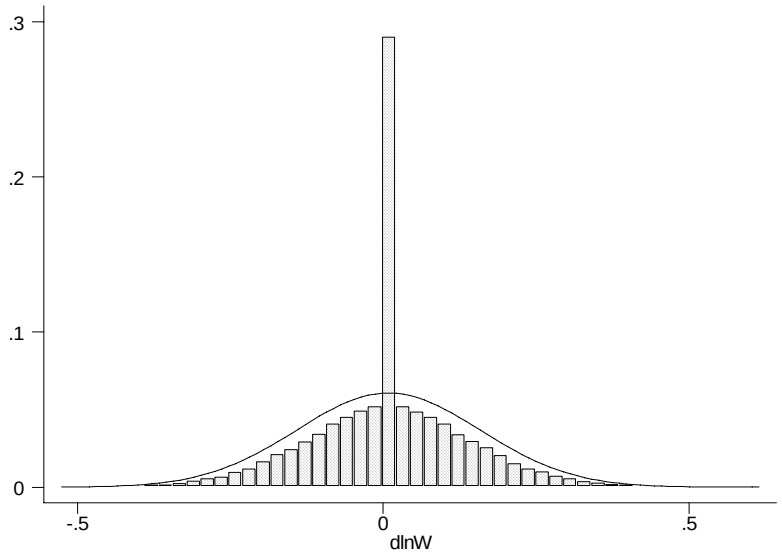


Figure 5: $f(\Delta \ln W)$ Implied by Theory when β=0



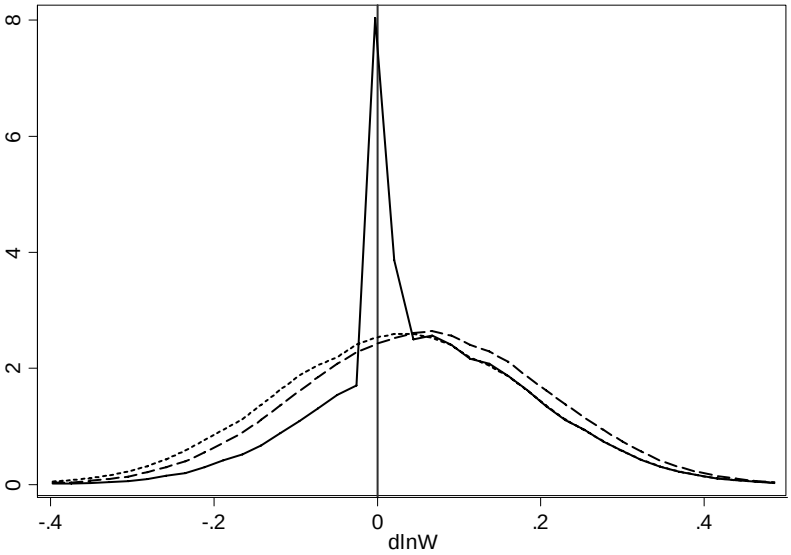
Notes:

Histogram: Actual Density; _____: Counterfactual Density

Simulations of $f(\Delta \ln W)$ predicted by model of 50,000 worker-firms after 10 periods.

Parameter values: $c = 0.1$, $\beta = 0$, $\pi = 0.02$, $\sigma = 0.15$.

Figure 6: Overstatement of Costs of DNWR



_____ Actual Density, _ _ _ _ _ _ True Counterfactual, ... Median-Symmetric Counterfactual

Figure 7: Overstatement of the Increase in Labor Costs as a Fraction of the True Costs of DNWR

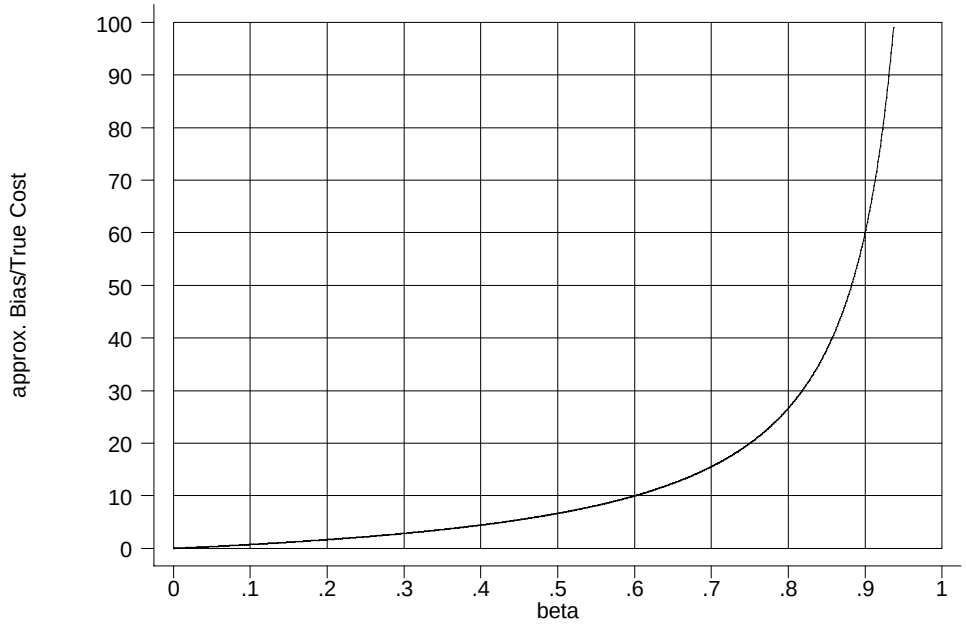
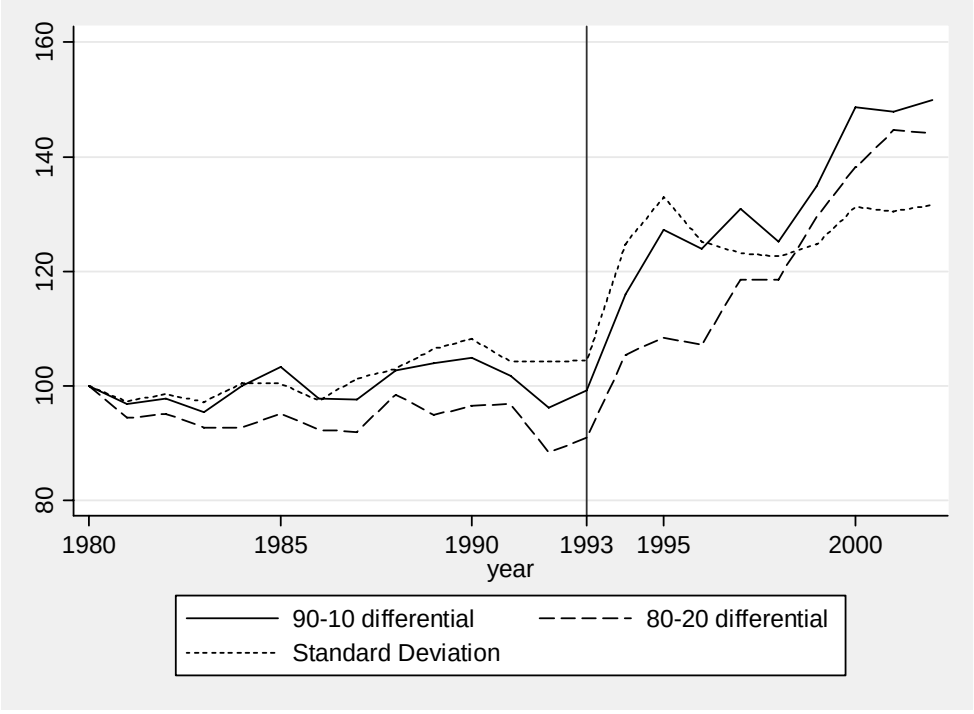
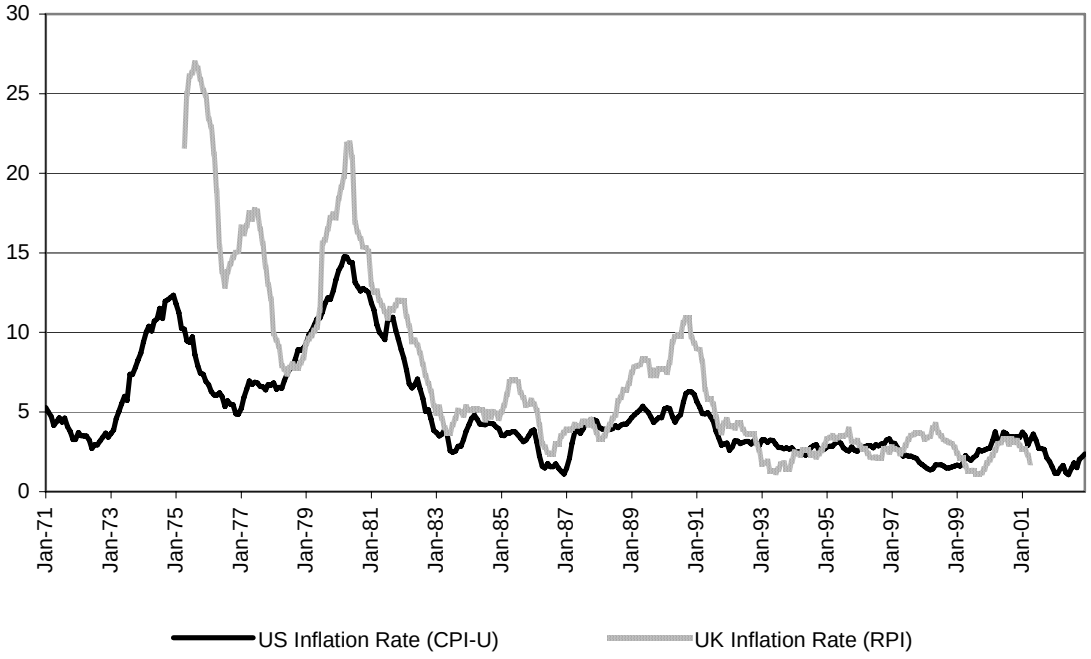


Figure 8: The Dramatic Increase in the Dispersion of Real Wage Changes after CPS survey redesign in 1994 (CPS)



Notes:
We measure dispersion by the standard deviation and the 90-10 and 80-20 percentile differentials of log real wage changes. All measures of dispersion are normalised to equal 100 in 1980.

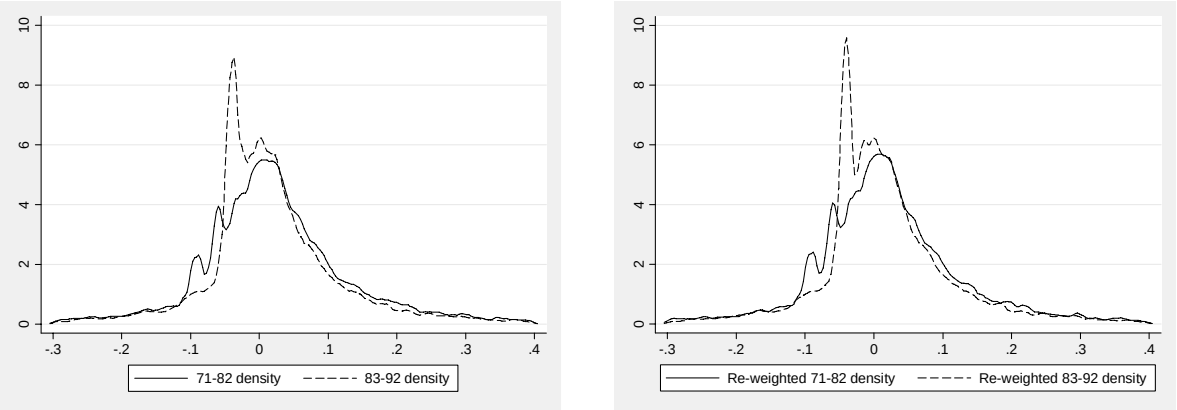
Figure 9: US & UK Inflation over the Sample Periods



Source: US CPI data was obtained from http://data.bls.gov. UK RPI data was obtained from www.statistics.gov.uk/statbase.

Figure 10: Density Estimates of Log Real Wage Change Distributions (PSID)

a) Without Re-weighting *b) With Re-weighting*[b]



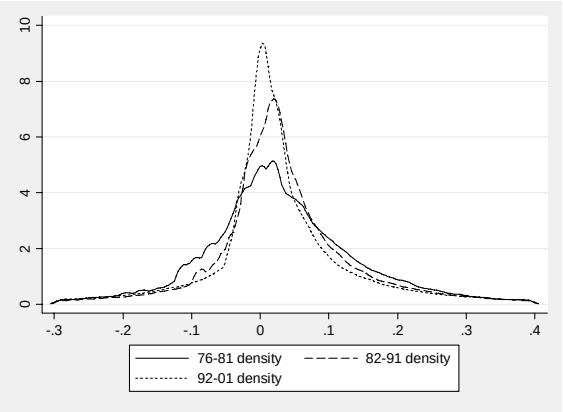
Notes:

a. Kernel density estimates using an Epanechnikov kernel, over 250 data points, and a bandwidth of 0.005.

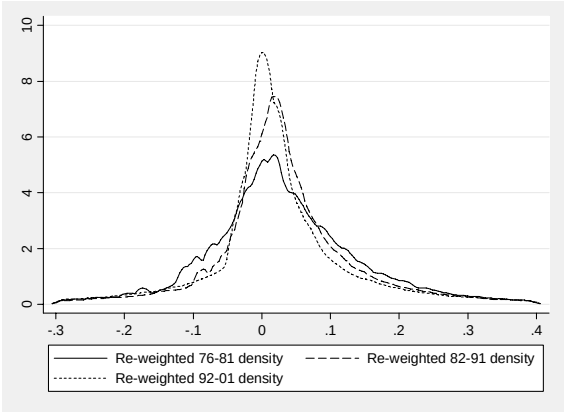
b. "Re-weighting" refers to the use of the DiNardo, Fortin & Lemieux (1996) re-weighting technique to control for changes in age, age^2, sex, education, 1-digit industry, 1-digit occupation, region, self employment, and tenure.

Figure 11: Density Estimates of Log Real Wage Change Distributions (NESPD)

a) Without Re-weighting



b) With Re-weighting



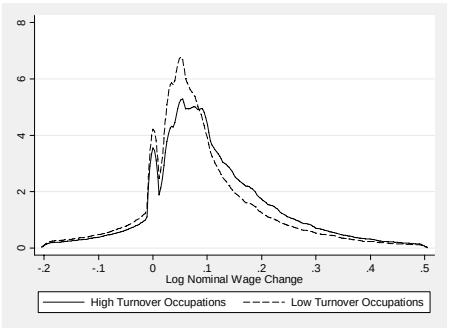
Notes:

a. Kernel density estimates using an Epanechnikov kernel, over 250 data points, and a bandwidth of 0.005.

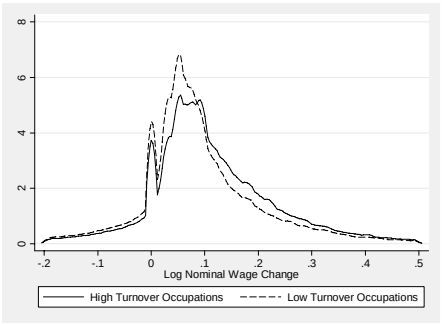
b. "Re-weighting" refers to the use of the DiNardo, Fortin & Lemieux (1996) re-weighting technique to control for changes in age, age^2, sex, region (including London dummy), 2-digit industry, 2-digit occupation, and major union coverage.

Figure 12: Turnover Effects on the Distribution of Log Nominal Wage Changes for Job Stayers (NESPD)

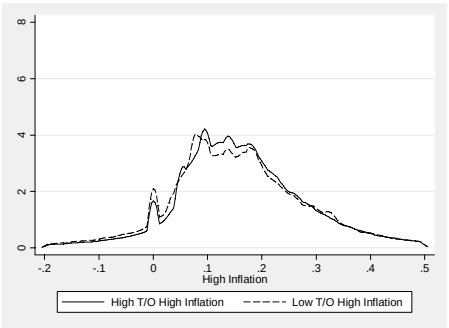
a) Without Re-Weighting



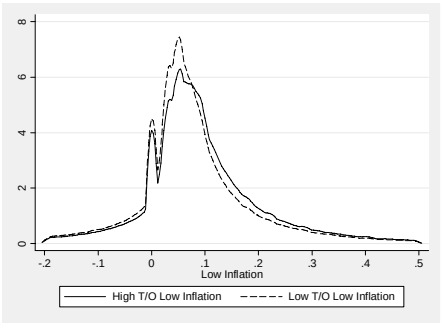
b) With Re-Weighting



c) Re-Weighted – High Inflation



d) Re-Weighted – Low Inflation



Notes:

a. Kernel density estimates using an Epanechnikov kernel, over 250 data points, and a bandwidth of 0.005.

b. "Re-weighting" refers to the use of the DiNardo, Fortin & Lemieux (1996) re-weighting technique to control for changes in adjusted lagged wages, age, age^2, sex, region (including London dummy), 2-digit industry, 2-digit occupation, and major union coverage.

c. High turnover refers to occupations for which the fraction of job changers in any given year exceeds the median.

d. High inflation refers to years for which the inflation rate exceeded 10%.

Table 1: Descriptive Statistics of Wage Changes (CPS, PSID, NESPD)

Year	US Data: CPS				PSID			British Data: NESPD			
	Π	Obs.	ΔW=0	Δω<0	Obs.	ΔW=0	Δω<0	Π	Obs.	ΔW=0	Δω<0
1971	4.4				1,520	10.39	34.41				
1972	3.0				1,527	11.59	32.35				
1973	6.3				1,599	8.88	46.34				
1974	10.0				1,676	8.35	56.74				
1975	8.3				1,733	7.39	42.07				
1976	5.7				1,471	7.48	34.33	18.9	60,318	0.67	41.00
1977	6.4				1,468	8.65	36.72	17.5	64,838	1.43	77.64
1978	6.8				1,605	7.35	37.57	7.9	66,168	2.15	33.73
1979	9.6				1,704	6.51	51.35	10.1	65,619	2.33	38.39
1980	11.2	25,626	5.70	53.39	1,756	4.38	52.51	21.8	66,574	0.44	46.81
1981	9.5	28,343	5.79	48.07	1,746	7.22	50.29	12	70,431	2.62	40.53
1982	6.1	27,426	10.41	45.76	1,664	8.17	38.58	9.4	75,745	3.01	49.34
1983	4.2	26,521	12.73	45.99	1,606	14.51	44.46	4	77,910	2.06	19.93
1984	4.3	26,675	12.76	46.29	1,621	12.95	46.33	5.2	75,652	5.09	41.62
1985	3.6	13,122	12.28	43.72	1,702	11.16	41.07	6.9	75,311	1.69	50.80
1986	1.9	6,935	13.67	40.63	1,830	15.30	42.51	3	74,487	1.39	18.88
1987	3.7	27,348	13.68	45.94	1,801	15.16	49.53	4.2	74,848	2.52	24.97
1988	4.1	26,825	12.59	46.43	1,848	15.42	50.87	3.9	73,440	1.55	20.57
1989	4.8	26,736	11.99	47.90	1,863	13.96	53.30	8	72,278	2.13	44.91
1990	5.4	28,045	11.14	49.11	1,815	12.01	54.66	9.4	70,752	2.49	50.33
1991	4.2	28,688	11.61	46.52	2,441	13.93	49.77	6.4	72,065	2.75	26.40
1992	3.0	28,521	13.43	44.94	2,441	16.39	45.60	4.3	76,335	4.87	30.87
1993	3.0	28,468	13.25	45.73				1.3	78,171	6.95	27.91
1994	2.6	26,584	11.88	44.49				2.6	78,167	6.36	48.14
1995	2.8	10,227	12.20	45.32				3.3	79,644	5.55	51.37
1996	3.0	8,458	11.46	44.68				2.4	82,489	1.53	32.31
1997	2.3	25,386	10.67	41.53				2.4	80,221	1.71	33.52
1998	1.6	25,255	10.31	38.00				4	76,999	4.08	51.19
1999	2.2	25,489	9.80	41.02				1.6	77,227	4.38	25.93
2000	3.4	25,215	9.68	44.19				3	76,806	4.35	39.69
2001	2.9	24,574	9.32	42.65				1.8	79,689	0.00	32.69
2002	1.6	26,575	10.32	42.38							

Notes:
"Π" denotes the rate of inflation in a given year. This is measured by the CPI-U-X1 for the US, and the RPI for the UK.
"Obs." Refers to the number of non-missing wage change observations each year.
"ΔW=0" reports the percentage of nominal wage changes each year that are exactly zero.
"Δω<0" reports the percentage of real wage cuts implemented each year.

Table 2: Summary Statistics (CPS, PSID, NESPD)

(a) CPS:	Obs	Mean	Std. Dev.	Min	Max
Change in log real wage	547042	0.025061	0.310781	-5.72642	4.562072
Age	547042	38.01381	12.52692	16	65
Female	547042	0.501651	0.499998	0	1
Education:					
< High School	546516	0.175773	0.380628	0	1
High School	546516	0.451021	0.497596	0	1
Some College	546516	0.287882	0.452776	0	1
College Degree	546516	0.071936	0.258382	0	1
Advanced Degree	546516	0.013388	0.114931	0	1
Metropolitan area	521083	0.710509	0.453527	0	1
Non-white	547042	0.204385	0.403252	0	1
Self-employed	546877	0.000104	0.010209	0	1

(b) PSID:					
Change in log real wage	33283	0.022087	0.337482	-3.72463	4.619859
Age	33283	38.24457	11.53739	18	65
Female	33283	0.196617	0.397446	0	1
Education:					
0-5 grades	30671	0.036256	0.186929	0	1
6-8 grades	30671	0.114375	0.318272	0	1
9-11 grades	30671	0.218382	0.413155	0	1
12 grades	30671	0.448469	0.497346	0	1
Some College	30671	0.130253	0.336587	0	1
College degree	30671	0.040201	0.196433	0	1
Advanced degree	30671	0.012064	0.109171	0	1
Tenure:					
[1, 1.5] years	30536	0.092907	0.290306	0	1
(1.5, 3.5) years	30536	0.204546	0.403376	0	1
[3.5, 9.5) years	30536	0.354008	0.47822	0	1
[9.5, 19.5) years	30536	0.236999	0.425249	0	1
19.5 years +	30536	0.111541	0.314805	0	1
Self-employed	33257	0.014313	0.118779	0	1

(c) NESPD:					
Change in log real wage	1922184	0.026539	0.190503	-9.9292	9.757886
Age	1922184	41.01464	11.85092	16	65
Female	1922184	0.409069	0.491662	0	1
Major union coverage	1922029	0.426511	0.49457	0	1
London dummy	1919091	0.144433	0.351528	0	1

Notes:
CPS sample also contains 2-digit industry classifications, and 50 regional dummies.
PSID sample also contains 1-digit industry and 1-digit occupation classifications, and 6 region dummies.
NESPD sample also contains 2-digit industry and 2-digit occupation classifications, and 10 region dummies. Major union coverage variable does not include more disaggregated union agreements.

Table 3: Regressions of Percentiles of Real Wage Changes on the Rate of Inflation and Controls (CPS, 1980 – 2002)

Percentile	Coefficient on Inflation Rate[a]					
	No Controls[b]		Aggregate Controls[c]		Full Controls[d]	
10th	-0.062	[0.134]	0.06	[0.148]	-0.067	[0.151]
20th	-0.254	[0.070]***	-0.157	[0.090]*	-0.15	[0.091]
30th	-0.326	[0.074]***	-0.275	[0.074]***	-0.366	[0.068]***
40th	0.04	[0.040]	0.049	[0.043]	0.035	[0.038]
60th	0.053	[0.024]**	0.053	[0.025]**	0.045	[0.028]
70th	0.123	[0.050]**	0.123	[0.051]**	0.127	[0.049]**
80th	0.178	[0.093]*	0.198	[0.100]*	0.154	[0.092]
90th	0.162	[0.159]	0.262	[0.172]	0.373	[0.131]***
Lower Tail Losses	+0.98%		+0.37%		1.10%	
Upper Tail Gains	-0.98%		-1.45%		-1.57%	
↑ in $\Delta\overline{w}$ due to DNWR[e]	0%		-1.07%		-0.47%	

Notes:

a. Reports Least Squares estimates (weighted by region size) of real wage change percentiles on the rate of inflation and controls.

b. Includes a dummy for the years 1994 onwards to control for the increase in dispersion of real wage changes following introduction of CAPI.

c. As b, but includes additional controls for the absolute change in the rate of inflation, and the contemporaneous and lagged state unemployment rate.

d. As c, but uses real wage change percentiles re-weighted for changes in age, age^2, sex, race, region (including metropolitan dummy), 2-digit industry, education, public sector employment, and self-employment.

e. Predicted effect on real wage growth of a change in inflation from 22% (maximum NESPD sample inflation, 1980) down to 1.3% (minimum NESPD sample inflation, 1993). Computed from estimation of 99 percentile regressions of the form summarised in the Table using the method outlined in the main text.

f. Standard errors in brackets: robust to non-independence within years.

g. * significant at the 10% level; ** significant at the 5% level; *** significant at the 1% level.

Table 4: Regressions of Percentiles of Real Wage Changes on the Rate of Inflation and Controls (PSID, 1971 – 92)

Percentile	Coefficient on Inflation Rate[a]					
	No Controls		Aggregate Controls[b]		Full Controls[c]	
10th	-0.228	[0.076]***	-0.227	[0.077]***	-0.236	[0.100]**
20th	-0.596	[0.090]***	-0.57	[0.099]***	-0.585	[0.100]***
30th	0.018	[0.034]	0.003	[0.041]	-0.026	[0.040]
40th	0.017	[0.024]	0.016	[0.027]	0.012	[0.030]
60th	0.029	[0.023]	0.027	[0.027]	0.015	[0.033]
70th	0.096	[0.047]*	0.11	[0.051]**	0.066	[0.054]
80th	0.177	[0.074]**	0.173	[0.081]**	0.16	[0.081]*
90th	0.313	[0.123]**	0.33	[0.122]**	0.301	[0.118]**
Lower Tail Losses	+1.05%		+1.13%		+1.18%	
Upper Tail Gains	-0.90%		-1.11%		-1.04%	
↑ in $\Delta\overline{w}$ due to DNWR[d]	+0.15%		+0.02%		+0.14%	

Notes:

a. Reports Least Squares estimates (weighted by region size) of real wage change percentiles on the rate of inflation and controls.

b. Controls for the absolute change in the rate of inflation.

c. As b, but uses real wage change percentiles re-weighted for changes in age, age^2, sex, education, 1-digit industry, 1-digit occupation, region, self employment, and tenure.

d. Predicted effect on real wage growth of a change in inflation from 22% (maximum NESPD sample inflation, 1980) down to 1.3% (minimum NESPD sample inflation, 1993). Computed from estimation of 97 percentile regressions of the form summarised in the Table using the method outlined in the main text – bottom and top percentiles are trimmed away as these yield extreme results.

e. Standard errors in brackets: robust to non-independence within years.

f. * significant at the 10% level; ** significant at the 5% level; *** significant at the 1% level.

Table 5: Regressions of Percentiles of Real Wage Changes on the Rate of Inflation and Controls (NESPD, 1976 – 2001)

Percentile	Coefficient on Inflation Rate[a]					
	No Controls[b]		Aggregate Controls[c]		Full Controls[d]	
10th	-0.138	[0.061]**	-0.057	[0.061]	-0.103	[0.068]
20th	-0.286	[0.033]***	-0.227	[0.018]***	-0.239	[0.021]***
30th	-0.197	[0.023]***	-0.149	[0.020]***	-0.143	[0.023]***
40th	-0.101	[0.016]***	-0.089	[0.016]***	-0.089	[0.011]***
60th	0.088	[0.008]***	0.076	[0.009]***	0.074	[0.008]***
70th	0.165	[0.015]***	0.147	[0.017]***	0.148	[0.014]***
80th	0.206	[0.028]***	0.176	[0.027]***	0.172	[0.023]***
90th	0.117	[0.053]**	0.066	[0.047]	0.103	[0.043]**
Lower Tail Losses	+1.26%		+0.71%		+1.16%	
Upper Tail Gains	-0.96%		-0.67%		-1.04%	
↑ in $\Delta\overline{w}$ due to DNWR[e]	+0.30%		+0.04%		+0.13%	

Notes:

a. Reports Least Squares estimates (weighted by region size) of real wage change percentiles on the rate of inflation and controls.

b. Includes a dummy for the year 1977 to control for the dramatic fall in real wage growth due to the incomes policies implemented in the UK at that time.

c. As b, but includes additional controls for the absolute change in the rate of inflation, and the contemporaneous and lagged regional unemployment rate.

d. As c, but uses real wage change percentiles re-weighted for changes in age, age^2, sex, region (including London dummy), 2-digit industry, 2-digit occupation, and major union coverage.

e. Predicted effect on real wage growth of a change in inflation from 22% (maximum NESPD sample inflation, 1980) down to 1.3% (minimum NESPD sample inflation, 1993). Computed from estimation of 99 percentile regressions of the form summarised in the Table using the method outlined in the main text.

f. Standard errors in brackets: robust to non-independence within years.

g. * significant at the 10% level; ** significant at the 5% level; *** significant at the 1% level.

Table 6: Effect of Turnover on Percentiles of Nominal Wage Increases for Job Stayers (NESPD)

Percentile	(1)	(2)	(3)[c]		
	τ	τ	τ	π	π*τ
60th	0.023 [0.002]***	0.016 [0.002]***	0.011 [0.004]***	0.081 [0.010]***	0.057 [0.037]
70th	0.047 [0.003]***	0.032 [0.003]***	0.028 [0.008]***	0.133 [0.020]***	0.024 [0.075]
80th	0.074 [0.005]***	0.054 [0.005]***	0.066 [0.010]***	0.153 [0.027]***	-0.2 [0.102]*
90th	0.113 [0.008]***	0.092 [0.008]***	0.116 [0.020]***	0.077 [0.064]	-0.29 [0.207]
Controls	Median Wage Change 1977 dummy	(1) + Current & Lagged Regional U/E Rates	(2) + Inflation & Inflation*Turnover		

Notes:

a. Report Least Squares estimates, weighted by cell (occupation, region, year) size.

b. Uses nominal wage change percentiles for job stayers, re-weighted for changes in adjusted lagged wage, age, age^2, sex, region (including London dummy), 2-digit industry, 2-digit occupation, and major union coverage.

c. Standard errors robust to non-independence within years.

d. * significant at the 10% level; ** significant at the 5% level; *** significant at the 1% level.