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ABSTRACT

Using the Panel Study of Income Dynamics, we find that true wage changes have many fewer nominal cuts and more nominal freezes than reported nominal wage changes. The data overwhelmingly rejects a model of flexible wage changes and provides some evidence against a model of perfect downward rigidity in favor of a more general model. The more general model incorporates downward rigidity but specifies that nominal wage cuts may occur when large cuts would occur in the absence of wage rigidity. However, the results of the general model imply that nominal wage cuts are rare. We also analyze the personnel files of a large corporation and find cuts in base pay are rare and almost always associated with changes in full time status or a switch between compensation schemes involving incentives. Our evidence on the consequences of nominal wage rigidity is mixed. We find modest support for the hypothesis that workers who are overpaid because of nominal wage rigidity are less likely to quit.

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I Introduction

In this paper, we address two questions: First, to what extent are nominal wages downwardly rigid? Second, how big an effect do such rigidities have on wage levels, wage changes, and quits, layoffs, and promotions? The answers to these questions are important because macroeconomists in the Keynesian tradition have long argued that nominal rigidities play a key role in the dynamics of wages and unemployment. The answers are crucial to understanding the effects of inflation on labor markets and to determining the costs and benefits of positive inflation.

While there is a large aggregate time series literature on the interaction between wage, price, and employment dynamics, research based on data on individual wage rates is much more limited. The recent literature on nominal wage rigidity begins with a study by McLaughlin (1994), who provides references to a few earlier studies. McLaughlin concludes wages are generally flexible. He finds that 17% of annual wage changes involve nominal wage cuts and argues that, even after accounting for measurement error, at least 12% of wage changes involve nominal wage cuts. Kahn (1997) shows that the probability of a nominal wage change falling into any range with end points defined as "the median pay change minus r percentage points" and "the median pay change minus r+1 percentage points" depends negatively on an indicator variable for whether a pay change in that range involves a nominal wage cut. Card and Hyslop (1997) assume for identification that the wage change distribution is symmetric around the median wage change, and that the upper half of the distribution is unaffected by nominal rigidities. They then calculate a counterfactual bottom half of the wage change distribution that would hold in the absence of wage rigidity. They find that downward rigidity exists but has little effect on real wage growth. The methodologies of both Kahn and Card and Hyslop implicitly assume that there is no measurement error in wage reports although they both discuss likely effects of measurement error on their results.[1] However, McLaughlin (1999) provides some evidence against the assumption that wage changes are symmetric in the upper half of the distribution, suggesting Card and Hsylop's methods may overstate nominal rigidity. Akerlof, Dickens, and Perry (1996) investigate the importance of measurement error by taking a small wage sample in which there are very few nominal wage cuts and

[1] Using the Panel Study of Income Dynamics Lebow, Stockton, and Wascher (1995) also find evidence of a spike in the nominal wage change distribution at zero and interpret asymmetries in the wage change distribution as at least partially due to nominal rigidity. Blinder and Choi (1990) and Bewley (1997) provide evidence based on employer surveys that nominal wage cuts are uncommon as well as some evidence about why firms are reluctant to cut wages. We should emphasize at the outset that we have nothing new to say about the longstanding but still unresolved theoretical issue of *why* there would be downward nominal wage.

dirtying the data. They find that the mismeasured data contains a similar percentage of nominal wage cuts to that in the PSID.[2] Shea (1997) matches a sample of unionized workers in the PSID to provisions of particular long-term contracts. While only 1.3% of the sample have true nominal wage cuts, 21% report nominal wage cuts. These results imply that measurement error could account for almost all nominal wage cuts reported in survey data. Groshen and Schweitzer (1999) use inflation-induced inter-occupational wage changes to identify the effects of downward nominal rigidity.

Overall, the research since McLaughlin's important study suggests that nominal wage cuts are less common in the U.S. than he found. [3] However, it does not provide us with a precise estimate of how common nominal wage cuts are, or much information on when they occur. The main source of the difficulty is measurement error in wage reports, which makes it very difficult to identify when nominal wage cuts have occurred.

We add to the literature on the extent of nominal wage rigidity in two ways. First, we use the presumably error free wage data from the personnel files of a large financial corporation to check how many workers receive wage cuts. We find that only about 0.5 percent of salaried workers receive nominal wage reductions. Only 2.5 percent of hourly workers receive wage reductions. Furthermore, almost all of these reductions are associated with changes in part time/ full time status or switches to compensation schemes involving performance incentives. On the other hand, we do find that reductions in annual bonuses are quite common among salaried workers.

Unfortunately, the wage policies of the firm we study may not be representative of the labor market at large, which is the disadvantage of the personnel file relative to noisy data from household surveys. Our second approach builds on earlier research using the Panel Study of Income Dynamics but uses an econometric model that allows for a more general treatment of the wage setting process and of measurement error. In contrast to McLaughlin (1994), who adds normally distributed measurement error to a flexible wage model, our econometric model nests a flexible wage model, a downwardly rigid model, and a model that allows for nominal cuts in certain circumstances, and so allows us to test

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[2] They also provide a good summary of recent research on nominal wage rigidity.

[3] We do not know of much micro data based research for other countries on the issue of nominal wage cuts. However, a recent paper by Smith (1999) analyzes data from the British Household Panel and concludes that wage cuts are common in Britain. This data set contains an indicator for whether the respondents checked their pay slips when reporting pay level. Smith concludes that nominal wage cuts are common on the basis of the fact that 18% of persons who report pay based on pay stubs report nominal cuts. She assumes that pay stub based reports are accurate. A disadvantage is that the pay measure is total weekly pay rather than base hourly wage rate. Smith addresses and largely rejects the possibility that hours changes, overtime and bonuses may lead to an overstatement of wage cuts but we are not fully persuaded by this evidence, particularly with respect to hours changes.

between them. We use worker characteristics, the price level, and the state unemployment rate to gain information on what wage changes should be in the absence of nominal wage rigidity. For example, young people have faster real wage growth, which means that a nominal wage floor is less likely to be binding for them. We use estimates from the model to determine the proportion of workers who experience wage cuts and wage freezes and to measure the effects of downward rigidity on wage levels and wage changes. Our treatment of measurement error allows for the possibility that some individuals correctly report their wage rates, which is a zero probability event in a model with normal measurement error.

There is little micro data research addressing our second question---do downward nominal rigidities affect the structure of wages and labor market outcomes such as layoffs, quits, and promotions?[4] We address this question by examining the relationship between these variables and estimates from our wage model of the extent to which the wage of a particular worker is affected by nominal wage rigidity. We also analyze whether nominal wage rigidity in an environment of low inflation might compress the wage distribution. We do this by examining whether the gap in wage growth between workers who are predicted to have fast wage growth and workers who are predicted to have slow wage growth depends on the inflation rate.

Our strongest finding is that nominal wages show substantial downward nominal rigidity in the U.S. economy. We find some support for the hypothesis that workers are less likely to quit when they are protected by nominal wage floors. Our estimates of the effect of nominal wage rigidity on layoffs and promotions and estimates of the effects of inflation on relative wage growth rates are too imprecise for us to draw any conclusions.

The structure of the paper is as follows: In section II, we examine the frequency of nominal wage cuts in a large financial corporation. In Section III we describe our model of nominal wage changes and discuss how we estimate it. In section IV, we describe the data from the Panel Study of Income Dynamics that we use in this paper. In section V, we present the results of the model and the implications of the model for wage cuts, wage freezes, wage levels, and wage changes. In section VI, we test whether downward rigidity affects the probability of quits, layoffs, and promotions, and we look at the effects of nominal rigidities on the wage structure. Section VII concludes.

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[4] Ackerlof et al (1996) use a simulation model of the economy to analyze the costs of nominal wage rigidity and conclude that they are substantial.

II. Nominal Wage Cuts in a Large Corporation

If the respondents in household surveys reported wages without error, measuring the incidence of annual wage cuts would be straightforward. Unfortunately, wages are imperfectly measured in such surveys. For this reason, we make use of the personnel records from a large U.S. financial corporation to take a preliminary look at the issue.

Figure 1a presents the frequency of the change in log of annual salary from May 1996 to May 1997 for 32,138 salaried workers who were active at both points in time. As one can see, almost no workers (.47 percent) received a wage cut. 11.36 percent had no salary change, 2.49 percent had an increase between 0 and .02, 47.8 percent had an increase between .02 and .05, 21.91 percent had an increase between .1 and .2, and 3.65 percent had an increase greater than .2. A portion of the persons with a 0 change in salary may not have had a salary review take effect during the year.[5] However, the results show that nominal salary cuts are rare in this firm.

Figure 1b presents the frequency of the change in the hourly wage for the corresponding sample of 1723 persons who were paid by the hour in both years. A total of 2.5 percent of these workers experienced cuts. Interestingly, about 7/8th of these are accounted for by changes between -.2 and -.5, while only 1/8 (5 workers) received cuts between 0 and -.1.

We have cross tabulated the changes in salary and in hourly wage rates with a file containing information about the reason for the change in compensation. The results show that almost all of the reductions are associated with changes between full and part time status, or with changes in whether performance incentives are part of compensation.

Some workers in the firm receive bonuses, and it is natural to ask whether they may provide a way for the firm to circumvent nominal wage rigidity. Four percent of salaried workers experienced a nominal decline in the sum of bonus plus annual salary, with declines in bonuses accounting for most of this. Among 7494 individuals who received a bonus in both years, 8.69% experienced a nominal decline. The very term "bonus" implies that compensation is transitory, and so it is perhaps not

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[5] At Northwestern University, salary reviews of both exempt and non-exempt staff members are conducted in June. They become effective September 1st. In many cases a person who starts work on July 1 would not receive a salary change until September 1 of the following year. However, our estimate of the incidence of zero wage changes was not sensitive to restricting the analysis to workers who joined the company before May 1995. We did not eliminate observations from the PSID that involved interviews less than one year a part. The results of Lebow et al (1995) suggest that we will overstate freezes at an annual interval by a small amount.

surprising that a “bonus” is less subject to downward nominal rigidity. Few hourly workers received bonuses, so bonuses have little effect on the distribution of changes in compensation of this group.

Overall, our results suggest reductions in nominal wages are quite rare, especially for salaried workers. Changes of 0 in compensation are fairly common, which, in combination with the small number of wage cuts suggests nominal wage ridigity is real. Bonuses serve to reduce nominal rigidity, but nominal reductions that do not involve a change in part time status or a change in performance incentives are quite uncommon in this firm. An interesting subject for future research is whether the use of bonuses is positively related to the probability that a nominal wage floor would be binding.

III An Econometric Model of Wages in the Presence of Downward Rigidity and Measurement Error

Unfortunately, the wage policies of the firm we study may not be representative of the labor market at large. To study the extent of nominal wage rigidity in the U.S. economy, we work with the Panel Study of Income Dynamics and an empirical model of observed nominal wage changes that incorporates measurement error and allows for nominal rigidity but also is consistent with flexible wage changes. There are unfortunately few coherent economic models of downward nominal wage rigidity. The principle difficulty is that most economic models have the property that agents care about real rather than nominal variables. Thus, the rich literature on implicit contracts and the role of incentives in wage setting has implications for real wage dynamics but has little to say about nominal wages. One exception in this literature is MacLeod and Malcomson (1993), which provides a rationale for nominal wage contracting. Our approach is to use an empirical model of nominal wage rigidity that has some intuitive appeal. We then show that this model is consistent with the MacLeod and Malcomson formulation. In section A, we present a model that nests perfect downward wage rigidity and perfectly flexible wages as special cases. In Section B, we show that the MacLeod and Malcomson model also fits within our framework. In Section C, we incorporate measurement error into the analysis and, in Section D, we describe how we deal with problems posed by permanent heterogeneity in wages.

A. The Model

The model assumes that there is an optimal wage that the firm would like to implement **in the absence of downward nominal rigidity in period t,** but possibly taking account of the fact that the wage chosen today will constrain later wage choices. We will refer to this wage as the "notional wage".[6] The constraints imposed by downward rigidity may imply that the firm does not implement the notional wage if doing so would imply a negative nominal wage change. We write the log of the notional wage (w_{it}*) as a function of a vector of explanatory variables (x_{it}), a parameter vector (β) and a normally distributed error term (ε_{it}).

$$w_{it}* = x_{it}\, b + e_{it}$$

Let w_{it}^{0} be the log of the nominal wage at time t. We structure the model so that if the notional wage change implies a nominal wage increase, the actual wage change equals the notional wage change. It is important to keep in mind that the value of β is influenced by whether employers take into account the possibility of being constrained by downward rigidity in the future when setting current wages.[7] The value of β in a labor market characterized by nominal wage rigidity is likely to differ from the value of β when wages are perfectly flexible.

If the notional wage change is a nominal wage cut of less than $-\alpha$, then the model specifies that the actual wage change is zero. The parameter α is a positive constant to be estimated. The model allows nominal wage cuts to occur when the notional wage change is sufficiently negative. The model is

$$\begin{array}{lll} w_{it}^{0} - w_{it-1}^{0} = x_{it}\, b - w_{it-1}^{0} + e_{it} & if & 0 \le x_{it}\, b + e_{it} - w_{it-1}^{0} \\ w_{it}^{0} - w_{it-1}^{0} = 0 & if & -a \le x_{it}\, b + e_{it} - w_{it-1}^{0} \le 0 \\ w_{it}^{0} - w_{it-1}^{0} = l + x_{it}\, b - w_{it-1}^{0} + e_{it} & if & x_{it}\, b + e_{it} - w_{it-1}^{0} \le -a \end{array} \qquad (1)$$

[6] The notional wage can be thought of as an efficiency wage that maximizes expected profit for the firm. Alternatively, one can also think about it as the reservation wage of the worker in a model where this reservation wage is known to the employer, the employer makes a take it or leave it wage offer to the worker, and the firm ignores the fact that nominal wage rigidity may constrain future wage choices.

[7] The development of such a formal model of firm behavior where there are costs to reducing nominal wages and where the marginal revenue product of workers is stochastic is a natural extension of the approach we take. Given the prominent role of the assumption of nominal wage rigidity in the macro literature, it 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.

The parameter λ is a positive constant that determines the response of wage changes to the notional wage change when a cut is appropriate. When $w_{it}^* - w_{it-1}^0 = -a$, $w_{it}^0 - w_{it-1}^0 = l - a$. The heavy line in Figure 2 is the graph of the relationship between the actual wage change, $w_{it}^0 - w_{it-1}^0$, and the notional wage change, $w_{it}^* - w_{it-1}^0$ that model (1) implies. The figure also identifies the range of values of $w_{it}^* - w_{it-1}^0$ that correspond to flexible wages, freezes, and cuts. The wage model contains as special cases both a model of perfect wage flexibility and a model of perfect downward nominal wage rigidity. In the case of perfect flexibility, both λ and α are zero. For perfect downward nominal wage rigidity, λ is arbitrary and α approaches ∞. Because both these models are nested in the general model (1), one can test whether the restrictions implied by either perfect rigidity or perfect flexibility are consistent with the data.

B. The MacLeod and Malcomson model

In this section we show that our model nests MacLeod and Malcomson's (1993, hereafter MM) model of wage contracts. MM consider a fixed wage contract where $w = w^c$. Renegotiation of the wage is only by mutual consent. The worker and the firm have outside options $\underline{w}(t)$ and $\overline{w}(t)$ respectively. As long as $\underline{w}(t)$ is less than $\overline{w}(t)$, it is optimal for the match to continue. If in any period t, w^c is less than $\underline{w}(t)$ then the wage is adjusted upwards to $\underline{w}(t)$. If in any period t, w^c is greater than $\overline{w}(t)$ then the wage is adjusted down to $\overline{w}(t)$. That is, $\overline{w}(t)$ and $\underline{w}(t)$ act as constraints on the wage.

So why would firms use this kind of contract? The answer is that it does a good job of protecting both general investments and specific investments made by the firm and so the optimal investment level is achieved. Consider the following sequence of events: The contract wage w^c is agreed. The firm makes an investment I. The firm and worker can renegotiate the contract wage at any period t by mutual consent. The outside options vary across time, perhaps due to productivity shocks.

If the investment I is specific to the worker, then $\overline{w}(t)$ increases with I because greater I makes it more valuable to the firm to maintain the current worker rather than replace the worker with someone else. Thus, if sometime after the investment is made, the wage is adjusted downwards from w^c to $\overline{w}(t)$ because the firm's outside option constraint binds, then the firm would not get the full payoff to the investment. However, if the worker's outside option binds, the firm still gets the full payoff to the investment because $\underline{w}(t)$ does not depend on I. To maintain the firm's incentive to make worker-specific investments optimally, w^c should be set low enough so that the firm's outside option is never likely to bind.

The fixed wage w^c can be set in nominal or real values. In the case of specific investments, because it is fine for the worker's outside option to bind but not for the firm's to bind, the nominal wage will be better than the real wage when there is positive inflation. Inflation will help ensure that the firm's outside option never binds when the wage is fixed in nominal units.

The MM model has the contract wage (w^c) fixed in nominal terms and renegotiated up to $\underline{w}(t)$ whenever $\underline{w}(t)$ is greater than w^c, and renegotiated down to $\overline{w}(t)$ whenever $\overline{w}(t)$ is less than w^c (everything in logs). With positive inflation the latter is unlikely to occur and so nominal wages will generally be rigid downwards.

The MM model can be put in our framework as follows:

$$
\begin{aligned}
w_{it}^0 &= w_{it-1}^0 \ \text{ if } \underline{w}(t) \leq w_{it-1}^0 \leq \overline{w}(t) \\
w_{it}^0 &= \underline{w}(t) \ \text{ if } w_{it-1}^0 \leq \underline{w}(t) \\
w_{it}^0 &= \overline{w}(t) \ \text{ if } w_{it-1}^0 \geq \overline{w}(t)
\end{aligned} \qquad (2)
$$

One can specify functional forms for $\underline{w}(t)$ and $\overline{w}(t)$ in terms of the x variables and regression error and so end up with a nominal rigidity model that is quite similar in structure to the empirical model above. Consider the following parameterization of the MM model:

$$\underline{w}(t) = x_{it}\beta + \varepsilon_{it}$$

$$\overline{w}(t) = \alpha + x_{it}\beta + \varepsilon_{it} \text{ where } \alpha>0$$

We can know rewrite (2) as

$$
\begin{aligned}
w_{it}^0 &= x_{it} b + e_{it} \ \text{ if } w_{it-1}^0 \leq x_{it} b + e_{it} \\
w_{it}^0 &= w_{it-1}^0 \ \text{ if } x_{it} b + e_{it} \leq w_{it-1}^0 \leq a + x_{it} b + e_{it} \\
w_{it}^0 &= a + x_{it} b + e_{it} \ \text{ if } w_{it-1}^0 \geq a + x_{it} b + e_{it}
\end{aligned} \qquad (3)
$$

Comparing models (1) and (3), we see that model (3) is the special case of model (1) where λ is constrained to equal α. Thus, the intuitive model can be seen to encompass the MM model as a special case.

C. Measurement Error

Because wages are reported with error, we parameterize the reported wage (w_{it}) as a function of the true wage and the measurement error component u_{it}.

$$w_{it} = w_{it}^{0} + u_{it}$$

Substituting this equality into model (1), we get the following model, which is expressed in terms of the reported wage rather than the true wage.

$$\begin{aligned} w_{it} - w_{it-1} &= x_{it} b - w_{it-1} + e_{it} + u_{it} && if \quad x_{it} b - w_{it-1} + e_{it} + u_{it-1} \geq 0 \\ &= u_{it} - u_{it-1} && if \quad -a \leq x_{it} b - w_{it-1} + e_{it} + u_{it-1} \leq 0 \; (4) \\ &= l + x_{it} b - w_{it-1} + e_{it} + u_{it} && if \quad x_{it} b - w_{it-1} + e_{it} + u_{it-1} < -a \end{aligned}$$

We use two alternative specifications for the distribution of u_{it}. The first is the standard specification in which u_{it} is iid normal and independent of the other variables in the model. This implies almost immediately that u_{it}- u_{it-1} is normally distributed. This specification has two shortcomings. First, validation studies of wages and earnings show that a substantial fraction of workers report their wages exactly, which is a zero probability event if u_{it} is normally distributed with positive variance. Also, this formulation is inconsistent with the fact that a substantial fraction of respondents report the same wage in adjoining years. Consequently, most of our results are based on a specification in which u_{it} is equal to 0 with probability p and equal to a $N(0,\sigma_{u*}^{2})$ random variable u^*_{it} with probability (1-p). We sometimes refer to this specification as the mixed model of measurement error. Note that the mixed specification assumes that the probability of reporting the true wage is independent across time periods. However, we have experimented with specifications that allow for temporal correlation in the measurement error and the obtained similar results to those we report.

D. Estimation Issues

Unobserved heterogeneity poses a major problem for the estimation of this model. Because unobserved ability is correlated with w_{it-1} (cov(w_{it-1}, ε_{it}) > 0), the distribution of ε_{it} depends w_{it-1}. Hence, w_{it-1} cannot be treated as a predetermined variable in estimating the model. We deal with this problem in two ways.

In Method 1, we replace w_{it-1} with its conditional expectation given lagged values of x. The basic idea is as follows. The model (1) implies w^0_{it-1} is a function of w^0_{it-2} and w^*_{it-1}. Since w^0_{it-2} is similarly a function of w^0_{it-3} and w_{it-3}*, w^0_{it-1} is a complicated nonlinear function of the first wage on the job and all the notional values of the wage since then. Because of the complexity of the algebra required to calculate the expected value of w^0_{it-1} in this way and the additional problems posed by measurement error, we have instead approximated w^0_{it-1} by modeling its expectation $w^\wedge_{it-1}$ to be a linear function of x_{it-1}, x_{it-2}, and x_{it-3}. We then substitute for w_{it-1} in model (2) by writing

$$w_{it-1} = w^0_{it-1} + u_{it-1} = w^\wedge_{it-1} + m_{it-1} + u_{it-1}$$

After this substitution, the model (4) becomes,

$$\begin{aligned} w_{it} - w_{it-1} &= x_{it} b - w^\wedge_{it-1} + e_{it} + m_{it} && if \quad x_{it} b - w^\wedge_{it-1} + e_{it} \geq 0 \\ &= m_{it} && if \quad -a \leq x_{it} b - w^\wedge_{it-1} + e_{it} \leq 0 \qquad (5) \\ &= l + x_{it} b - w^\wedge_{it-1} + e_{it} + m_{it} && if \quad x_{it} b - w^\wedge_{it-1} + e_{it} < -a \end{aligned}$$

where $m_{it} = u_{it} - u_{it-1}$ and $e_{it} = e_{it} - m_{it-1}$. Model (5) is the model that we estimate. We assume that both m_{it} and e_{it} are normally distributed mean-zero random variables. Note that although one can always decompose w^0_{t-1} into its conditional expectation given lagged x variables and an error term, the error term μ_{it-1} is not likely to be a homoskedastic, normally distributed random variable because w^0_{t-1} is a nonlinear function of ε_{it-1}, ε_{it-2}, and so on. We assume it is normal for convenience. We estimate the model by maximum likelihood.[8]

In Method 2, we estimate the model under the assumption that w_{it-1} is approximately equal to w_{it-1}*. In this case we can use the equation $w_{it-1}=X_{it-1}\beta + \varepsilon_{it-1} + u_{it-1}$ to eliminate w_{it-1} from the model. One can justify this assumption when w_{it-1} is the first observation on the job, since the firm is not constrained by a past wage when setting the wage for a new employee. However, if nominal wage rigidity constrained the wage change between t-2 and t-1, w_{it-1} may exceed w_{it-1}^*. For this reason we prefer Method 1 on a priori grounds.

Using either of these approaches to estimation, there are several caveats that must be acknowledged. First, unobserved heterogeneity may bias the coefficients on the x variables. Fortunately, the coefficient values are not the central focus of this paper. All that matters for our purposes is that the wage estimates capture the total effect of particular variables on wages controlling for the other observed variables. Consequently, we define the coefficient vector β and the error term ε_{it} so that ε_{it} is orthogonal to x_{it}. Second, the sample is selected in that persons included worked for the same employer in two consecutive years. Below we use our model to study the impact of wage rigidity on quits and layoffs. Since quits and layoffs are likely to be affected by whether or not nominal wage floors are binding, there is the potential for selection bias when we perform the analysis on stayers. Our empirical results suggest tentatively that this may be an issue for quits and does not support strong conclusions regarding layoffs.[9] Third, the model assumes that there is no upward nominal rigidity in that it assumes that if the notional wage is greater than the previous wage, the notional wage is implemented. We are skeptical on a priori grounds that menu costs would cause firms to avoid adjusting wages at an annual interval. It is possible, however, that some firms do not bother adjusting wages when the notional increase is very small, say less than a half percent. Finally, we ignore possible effects of rounding of actual wage or reported wages.

IV Implications of the Model

A. <u>The Effects of Wage Rigidity on Wage Levels and Changes</u>

The model permits us to estimate several interesting variables that cannot be inferred directly from the wage data because of the problem of measurement error. First, we can estimate the proportion

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[8] Another approach would be to specifiy a functional form for the distribution of $(\varepsilon_{it}|w_{it-1},X_{it-1})$ and estimate the model parameters along with the parameters of this conditional distribution by maximum likelihood. A problem is that the model suggests that conditional distribution is unlikely to be normal.

[9] In principle, one could estimate the model controlling for selection in who stays in the firm, but we have not attempted to do so.

of workers taking nominal wage cuts each year. Given estimates of the model parameters we calculate the probability that $x_{it} b - w_{it-1} + e_{it} < -a$ conditional on x_{it} and $\hat{w}_{it-1}$ when we use Method 1 to estimate the model and conditional on x_{it} and x_{it-1} when we use Method 2. We then take the average of the probabilities over the sample members. Given estimates of β and α from the model, we calculate the probability that a worker with a given x takes a nominal wage cut and, hence, the proportion of workers that take wage cuts. Similarly, we use the model to estimate the proportion that have a nominal wage freeze and the proportion that have a nominal wage increase in each year. The sum of the proportion who take cuts and the proportion who have a wage freeze equals the proportion of workers whose wage changes are constrained by downward nominal rigidities.

The model implies that the actual wage will be at least as great as the notional wage. The measured wage may be less than the notional wage in some cases because of measurement error, but on average it will exceed the notional wage because measurement error has mean zero. We use the model estimates to calculate $E(w^0_{it} - w_{it}^*)$, the expected value of the difference between the actual value of the wage and the notional value. The details of the calculation of $E(w^0_{it} - w_{it}^*)$ are in Appendix 1.

Finally, we calculate what wage changes would be if there were no downward rigidities in the current period. The model provides estimates of both the notional value of the wage at t and of the actual value of the wage at t-1 and, by subtracting the latter from the former, we estimate the notional wage change. By comparing this estimate to reported nominal wage changes, we estimate the effects of downward rigidities on the average nominal wage change. Keep in mind, however, that the term notional wage is the optimal wage in the absence of wage rigidity in the current period but with the effect of current wage setting on future constraints accounted for. For example, wage offers to new employees may lie below what they would be in the absence of nominal wage rigidity even if firms are free to choose the starting wage subject to the constraint of competition for new workers.[10] In particular, the gap between that actual wage and the notional wage is not the effect that nominal wage rigidity in the economy has on wage levels. Consequently, these calculations should be taken with a grain of salt.

F. Real Effects of Nominal Rigidity on Layoffs, quits, and promotions

[10] The simulation model that Akerlof et al (1997) use to assess the effects of nominal simulation model ignores this type of behavior, which would mediate the tendency of downward nominal wage rigidity to produce a real wage distribution that is too high.

One of the main reasons to be interested in nominal wage rigidity is that it may have effects on real variables such as layoffs, quits, and promotions. Indeed, if nominal wage rigidity exists but has no effect on employment dynamics, it is unlikely that it is of any substantive importance to the economy. Solon, Whately, and Stevens (1997) present a case where the effects of wage stickiness within positions is massively attenuated by the strategic use of promotions and demotions within firms.[11]

The approach we take is to use the estimates from the structural estimation to calculate a predicted value of $w_{it}*$ for all workers employed at t-1 and then to examine how the probability of layoff depends on $w_{it}*$ and w_{it}. Since w_{it} is unobserved for workers who are not with the same employer at t and at t-1, we use the wage adjustment model to estimate the expected value of $w^0_{it} - w_{it}*$, as discussed in Appendix 1. One would expect that the larger is the expected value of $w^0_{it} - w_{it}*$, the higher the probability that the employer will lay off the worker, and the lower the probability that the worker will quit to a new job. We regress layoff, quit, and promotion indicators on a full set of explanatory variables and on either $E(w^0_{it} - w_{it}*)$ or the probability of a wage freeze for each worker.

V Data

We use data drawn from the 1971-1992 waves of the PSID. The sample is composed of workers who are in the same job at t as at t-1, who are paid by the hour, who are not self employed, and who are aged between 18 and 64. The PSID contains both a random sample and a poverty sample. We use only the random sample in our analysis. We work with two subsamples. The first is restricted to white non-union men. The second or "combined" subsample includes men and women of all races and both union and nonunion workers. However it is drawn from the 1976-1992 waves because the information available for wives is inadequate for our purposes before 1976.

There are two potential wage measures for hourly workers in the PSID: the reported hourly wage rate and annual average hourly earnings. We have chosen to use the reported wage, as it is specific to the current job and is also likely to be more accurately reported. Since the hourly wage is topcoded at $9.98 before 1978, many wage observations in those years are topcoded. In these cases average hourly

[11] Devereux (2000) forthcoming shows that there is little evidence that position changes are used to adjust wages to the business cycle in recent data from the PSID.

earnings for the year have been used to predict the wage.[12] Wage growth in year t is computed as the difference in the log of the wage in year t and in year t-1. We use the GNP deflator to construct our measure of inflation. We compute the inflation rate as the difference in the log of this index from t-1 to t.

We trim both of the subsamples of stayers by excluding 1% of wage change observations in each tail. The logic for trimming is to exclude wage changes that are most probably affected by measurement error. The effect of the trim on the white male sample is to exclude wage changes of stayers that involve cuts of greater than 23% and increases of greater than 51%. For the large sample, the equivalent numbers are cuts greater than 26% and increases of greater than 45%. In table 1, we present means of selected variables for both subsamples.

Figures 3 and 4 provide some further description of the nature of wage changes in the PSID data. In figure 3, we draw a bar graph for the sample of non-union white men that is equivalent to figure 1's representation of wage changes in the personnel data. The striking difference between figure 3 and figure 1 is that the number of nominal wage cuts is much higher in the PSID than in the personnel files. The key question we attempt to address with our econometric model is whether these are real or reflect measurement error in wage reports. Figure 4(a) shows how the proportion of cuts and freezes in the white non-union men sample varies with the inflation rate. The pattern is clear; the proportion of cuts and freezes are inversely related to the inflation rate. Figure 4(b) graphs the mean nominal wage change against the inflation rate. Once again there is a clear relationship with nominal wage increases being higher in periods of high inflation.

For the sample of non-union white men, the x variables include a constant, years of education, tenure at t, tenure squared at t, labor market experience at t, labor market experience at t squared, the log of the price level, the state unemployment rate at t, a time trend , and separate indicators for whether the worker was married at t and lived in the south at t. In the combined sample we add controls for female, white, female interacted with tenure and experience, and white interacted with tenure and experience.

When estimating $\hat{w}_{it-1}$, we regress w_{it-1} on tenure at t-1, tenure squared at t-1, union member at t-1, married at t-1, experience at t-1, experience squared at t-1, the price level at t-1, south at t-1, state

[12] There are very few topcoded wages in our samples: In the small sample, there is one topcoded value in 1978. In the combined sample, there is one case in 1977 and one case in 1978. We exclude salaried workers from the analysis because some salary changes may be due to changes in work hours, and standard hours levels are used to convert the salary reports into an hourly rate. The results of Lebow et al (1995), Kahn (1997), and McLaughlin (1999) for the PSID all suggest there may be less wage nominal rigidity for salaried workers but none of these studies deal with the possibility of differential measurement error or hours variation. The results for our one firm, if anything, suggest the opposite.

unemployment rate at t-1, and year at t-1. The regressions also include the state unemployment rate at t-2, the inflation rate at t-1, the inflation rate at t-2, and interactions of these three variables with each other and with all the other explanatory variables. $\hat{w}_{it-1}$ is then the predicted value of w_{it-1} from this regression. The full regression is reported in table A1.

VI Results for Wages and Wage Changes

We begin by showing that the amount of measurement error required to square the hypothesis of perfect downward wage rigidity to the data is consistent with the available evidence. We then show that the log likelihood of the flexible wage model is much lower than the general model and the perfect downward wage rigidity model. Finally, we discuss estimates of the key parameters of the wage rigidity models.

A Measurement Error and the Hypothesis of Perfect Downward Wage Rigidity

How much measurement error is required to reconcile the nominal wage cuts observed in the data with the hypothesis that there really are no nominal wage cuts? We address this question by comparing the estimates of the variance of measurement error for the special case of the model where there is perfect downward nominal wage rigidity to estimates based on matched employer/employee reports of wages. For this perfect wage rigidity model using the combined sample the MLE estimates of the probability p of reporting the true wage in each period is .5823 and the standard deviation of u_t^* is .0697 (table II, column 4).[13] The implied estimate of the variance of the measurement error u_t-u_{t-1} in the wage change is .0041.[14] The variance of reported wage changes in this sample is .0081. Thus, measurement error comprises (0.0041/0.0081)*100 = 51 percent of the variance of wage changes for the combined sample of men and women who are paid by the hour and don't change employers.[15] For the sample of white male nonunion workers, the MLE estimate of p is .666 and the estimate of σ_{u*} is .0675

[13] Asymptotic standard errors do not account for serial correlation among the observations on a given individual or, in the case of Method 1, for the fact that $\hat{w}_{t-1}$ is an estimated regressor.

[14] This estimate is obtained using the formula $\mathrm{var}(u_t - u_{t-1}) = 2 \cdot p \cdot (1-p) s_{u^*}^2 + 2 \cdot (1-p)^2 \cdot s_{u^*}^2$.

[15] While models with normally distributed measurement error and p=0 are inconsistent with the spike in wage change distribution at 0, it is nevertheless interesting to compute measurement error variance implied by those models. In Appendix Table 2 for the combined sample we report that the standard deviation of σ_u is .0466. The implied estimate of $\mathrm{var}(u_t\text{-}u_{t-1})=2(.0466)^2$ =.0043, which is about 6 percent higher than the estimate of .0041 based on the mixed measurement error model. The estimate of $\mathrm{var}(u_t\text{-}u_{t-1})$ is substantially higher in the untrimmed sample, as demonstrated in Altonji and Devereux (1998).

(table IV, column 3). The implied estimate of the variance of measurement error in wage changes is 0.0030. This compares to the variance of nominal wage changes in this sample of .0083. Thus, the variance of measurement error is about 36 percent of the variance of nominal wage changes.

To the best of our knowledge, the particular wage measure that we use here, a reported hourly wage rate for hourly workers, has never been validated in a measurement error study. However, in their study of measurement error in earnings, Bound and Krueger (1991) estimated that the variance of measurement error was 28% of the total earnings change variance. The findings of Bound, Brown, Duncan, and Rodgers, 1994 (BBDR) are similar for log earnings. However for the change in the log of earnings divided by hours, BBDR found that measurement error composes 62% of the variance of total wage changes for stayers. Our measurement variances are in the middle of these estimates. Unsurprisingly, when we restore the trimmed wage observations to the samples, the variance of measurement error rises considerably. In the sample of white nonunion men, the estimate rises to .012 and contributes 67 percent of the wage change variance of .0178. This measurement error estimate is just outside the range suggested by the validation studies. The overall conclusion from the rigid model is that, after trimming of extreme values, the variance of measurement error required for no actual nominal wage cuts is well within the range of estimates provided by the validation studies that we are aware of.

Card and Hyslop (1997) report results from a January 1977 CPS validation study that collected self-reported wage information from workers and matching information from their employers. They report that among hourly workers paid above the minimum wage, 44% of employers and employees report exactly the same wage. Given that the employer report may be mismeasured but presumably are at least as accurate as worker reports, this suggests that the proportion of workers who accurately report their wage is between 44% and 66.3%.[16] Our findings, using method 1, of a value of p equal to 0.666 is at the upper end of this range. Our method 2 estimates are above the range.

B Likelihood Ratio Tests of the Flexible Wage Model

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[16] Let p_f denote the probability that the firm responds correctly and assume that firms and workers make independent errors. Then $p \cdot p_f = .44$. If $p \leq p_f$, (firms are at least as accurate as workers), then $p = p_f = .663$ is the upper bound for p.

Since both the perfectly rigid and the flexible models are nested in the general model, one can use standard testing procedures to test the restrictions implied by the nested models. We conducted these tests for the two samples using the normal measurement error specification in which the probability that the worker reports the correct wage is 0. In this case, measurement error in wage changes is normally distributed. We chose this measurement error specification because it is more favorable to the flexible wage model.[17] The estimates for the combined sample using estimation method 1 are in table A2. The log likelihood values for the general model, rigid wage model, and flexible wage models are 14580, 14520, and 12124 respectively, and a likelihood ratio test overwhelmingly rejects the flexible model in favor of the general model. The perfect wage rigidity model is also strongly rejected against the general model at conventional significance levels, but the log likelihood value for the perfect rigidity model is much higher than that for the flexible model, implying that perfect downward rigidity is a much better approximation to reality than perfect flexibility. We obtain similar results when we estimate the rigid wage models using the sample of non-union white men. 2 (See Table A3).

A comparison of the estimates of the variance parameters from the flexible and rigid wage models provides some insight as to how the rigid model accounts for the variance in measured wage changes. In the flexible wage model, the variance of measurement error u_t-u_{t-1} and the error e_t in w^*_t - w^*_{t-1} are not separately identified but the total error variance is. The estimated variance of e for the general model is $.123^2$=.0151, which actually exceeds the estimate of $.109^2$=.0119 obtained for the total variance $(s_e^2 + 2s_u^2)$ based on the flexible wage model. (See Table A2). The estimate of the total variance $(s_e^2 + 2s_u^2)$ in the flexible model is only about 60% of the total error variance in the other models. The higher total variance of the error of the notional wage process in the rigid model arises

[17] The flexible model cannot account for the fact that the wage change is exactly 0 for 12 percent of the observations unless the distribution of the change in the stochastic component ε_t is not normally distributed and depends on the distribution of wage index (x_t-x_{t-1}) in a way that generates a mass point at 0. When ε_t is normal the likelihood function for the flexible model is 0 for the 0 wage change observations, and so the log likelihood for the flexible wage model sample is undefined. The models that allow for rigid wages are consistent with the mass point at 0 provided that some fraction of the respondents report the correct wage in both periods. Restricting both models to the case of classical, normally distributed measurement error presumably levels the "playing field" between the flexible and rigid model. We did not compare the rigid wage model to a symmetric menu costs model in which the true wage change is 0 if the notional wage change is near 0, although such a comparison would be interesting. The fact that there are many small positive wage changes in our firm sample and in the PSID raises doubts about a symmetric menu costs model. On the other hand, Smith's (1999) analysis of UK panel data suggests such a model has empirical relevance. Note that the values of the likelihood function for the two measurement error specifications are not comparable. The fact that the likelihood value for the rigid model with normal measurement error in Table A2 exceeds that for the mixed measurement error model is not a cause for concern or an indication that the true value of p is 0. Indeed, the likelihood

because downward rigidity reduces the variance in actual wage changes by taking what would be wage cuts and turning them into wage freezes.

C. Estimates of the wage rigidity parameters α and λ

We now turn to the estimates of the alternative models of wage rigidity. We return to mixed specification for measurement error. A key fact to keep in mind is that if wage cuts are rare, which means that the wage cut threshold parameter, α, is large, then the specific value of α and the value of λ are poorly identified, especially given that we have trimmed the data. Basically, a value of α of .35 and a value of 5000 both imply that wage cuts almost never happen. If wage cuts almost never happen, then the parameter λ does not influence the distribution of wage changes. The estimated standard errors for α and λ should be taken with a big grain of salt. We experienced numerical problems in the white nonunion sample when we estimated the general model, which does not restrict either parameter. We believe that this is a reflection of the fact that true nominal wage cuts are rare.

Table II presents model estimates for the combined sample using Method 1, which involves using the predicted value of w_{it-1} given lagged value of x to address the endogeneity of lagged wages in the model. The log likelihood for the general model is 7015, which is slightly above the values of 7011 for λ=0 case and 7010 for the rigid model. A likelihood ratio test rejects the perfectly rigid model against the general model at the .01 significance level. The estimate of α is .6536 (.1703) and the estimate of λ is .0442 with a huge estimated standard error. This imprecision in the estimate of λ is unsurprising because, as discussed above, λ is unidentified when α is large. The estimates α and λ are .319 (.0171) for the MM specification, which restricts α to equal λ. The log likelihood value for the MM specification is somewhat below that for the other specifications. The use of the mixed measurement error specification improves the relative performance of the perfectly rigid model, as comparison of the results in Tables II and A 2 demonstrates. The estimated probability of a nominal wage cut is essentially 0 for all four specifications and, as we shall see, the estimates imply that a large proportion of workers experience nominal wage freezes.

Table II also reports estimates of β, the coefficients on the x variables that appear in the equation (1) for w_{it}*. These estimates are not sensitive to the specification of wage rigidity. They are very similar to what we obtain when we estimate the flexible wage model (not shown). For example, the coefficient

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value for the rigid wage model with the mixed error structure is only -5669 when we estimate the model with p on education is .06, the coefficient on union is .2486, and the coefficient on female is -.2612. The value of 1.226 for the coefficient on the log price level is typical of what we obtain. The fact that the price level exceeds 1.00 is probably related to the fact that we obtain an unusually large negative coefficient of -.023 on the time trend, since the price level has a strong positive trend. One should keep in mind when interpreting the coefficients on the price level, the time trend, and the state unemployment rate that we have made no attempt to account for dynamics or for the fact that the current aggregate price level may not be predetermined in a nominal wage regression. Thus, one should not give these coefficients a structural interpretation.

Table III reports estimates based on Method 2, which uses the assumption that $w^0_{it-1} = w_{it-1}*$ and so eliminates w_{it-1} from the model. With Method 2, the general model and the MM model yield very similar results. Both specifications imply estimates of α and λ of about .06. The smaller value of α obtained using Method 2 goes hand in hand with a larger estimate (.035) of the probability of a cut and a smaller estimate (.149) of the probability of a freeze. There is an inverse relationship between the values of p and α that permit the model to fit the number of observations with 0 wage changes. The smaller value of α means that the probability the wage falls in the range in which there is a freeze is reduced, and so a larger value of p is "required" for the model to match the percentage of observations with 0 reported wage changes. The likelihood values of the general model and the perfectly rigid model are 9201 and 9113 respectively, so there is strong evidence against perfect rigidity.

In contrast to the Method 1 results in Table II, the coefficients on x variables in Table III refer to effects on the notional wage change rather than the notional wage level. They are not directly comparable to the wage level coefficients in Table II. These estimates are not sensitive to the wage rigidity specification or to whether we use the mixed or normal measurement error specification. The small point estimates bring attention to an unfortunate fact of life in surveys such as the PSID---it is very difficult to predict nominal wage changes with the available regressors. The R^2 of an OLS regression of the nominal wage change on the x variables used in Table III is only .068, and the smallest value of the predicted nominal wage change in the combined sample is .016. (The sample mean is .070). These facts imply that we are relying heavily on the assumption that ε_{it} is normally distributed for identification.

Table IV and V report corresponding estimates for the white nonunion male sample for 1971-1992. Basically, the estimates are broadly similar to those obtained for the combined sample and we

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constrained to .01, while the unconstrained maximum is 7010.

omit a detailed discussion to save space.[18] This fact suggests that union contracts are not the main source of nominal wage rigidity. Only a handful of observations on w_{it-1} are at or below the federal minimum wage in this nonunion white male sample or in the combined sample, and only 3 of these involves a 0 wage change. Consequently, minimum wages do not appear to be an important factor in nominal wage rigidity either.

D Implications of the Model Estimates for Wage Growth, Wage Freezes, and Wage Cuts

In this section we further explore the implications of the model estimates for wage growth. We will use the estimates from the general model for this exercise but since the estimates are similar across models it does not matter much which model we choose. In the case of the combined sample with estimation method 1, we were unable to estimate the general model. For this case, we use the estimates from the MM model. In columns 2 and 3 of table VI, we present means by year of the inflation rate and the reported nominal wage change $w_{it} - w_{it-1}$ for the combined sample. The rest of the columns report the mean of the notional wage change, which is the estimate from the model of $w_{it}* - w_{it-1}$, the probability of a wage cut, the probability of a freeze, and the mean of $E(w^0_{it} - w_{it}*)$, which is the conditional expectation of $w^0_{it} - w_{it}*$ given current and lagged x values. (See Appendix 1). The results in Table VI refer to the combined sample and are for the general model, which has the highest likelihood value. Columns 4, 5, 6, and 7 are based on the Method 1 parameter estimates while columns 8-11 refer to Method 2.[19]

In table VI, we see that the estimated probability of a nominal wage cut is essentially 0 in each of the years when we use Method 1. The estimates are below the value of 2.5% that we obtained for hourly workers using the firm personnel file. (As we noted earlier, almost all of the nominal cuts for that firm are associated with a change between combined and part time status or to a lessor extent a change in whether compensation is incentive based.) The very low estimates of the pay cut probability obtained from the model contrast sharply with the 11.1 % and 13.6% of reported nominal wage changes that

[18] For this sample we were not able to estimate the general model using Method 1. The large value of α means that there are no cuts to identify λ separately from α. See Table V, column 2. In the first column we note that we experienced difficulties in computing standard errors for the model with α free and λ constrained to 0.

[19] Results for the other versions of the rigid wage model are very similar when we use Method 1. In the case of Method 2, the probability of a cut is lower (essentially 0) than the results for the general model in the table when we restrict λ to equal 0. Table VI reports results for the combined sample. The patterns are similar to those for the combined sample and so we focus on the results for the combined sample.

involve a pay cut in the combined and nonunion white male samples (respectively). The model implies that essentially all of the reported nominal wage cuts are due to measurement errors rather than actual nominal cuts. The estimates also imply a high probability of a nominal wage freeze. For the combined sample across all years, 43% of wage changes are estimated to involve a nominal wage freeze. This estimate is very high in comparison to the percentage of freezes in reported wages of 11.7% in the combined sample and 13.6% in the sample of nonunion white men. Given measurement error one would expect the estimated probability of freeze to exceed the fraction observed in the raw data, but the estimates are also much higher than the value of 12.71 percent that we obtained for hourly workers from the particular firm. Using Method 2 the implied percentage of freezes is 14.9 percent, which is still quite substantial. We do not fully understand why Method 1 and II yield such different estimates of the percentage of wage freezes. As a matter of arithmetic, the difference is closely related to the fact that with Method 2 we get much higher estimates of p and lower estimates of α.

There is a lot of variation across years in the proportion of nominal wage cuts and/or nominal wage freezes. Pay cuts and freezes are more likely to occur when inflation is low such as during the mid-eighties and early nineties. When we regress the yearly estimates based on Method 2 of the probability of a freeze or a cut on the inflation rate and the change in the state unemployment rate we obtain

$$\text{Prob(Cut)} = \underset{(0.0076)}{0.0585} - \underset{(0.0013)}{0.0067} * \text{Inflation rate} + \underset{(0.0023)}{0.0022} * \text{Change in State Unemployment Rate}$$

$$\text{Prob(Freeze)} = \underset{(0.0258)}{0.3331} - \underset{(0.0080)}{0.0338} * \text{Inflation rate} + \underset{(0.0074)}{0.0080} * \text{Change in State Unemployment Rate} .$$

For Method 1, we only report the regression for Prob(Freeze) because the odds of a cut are essentially 0 in all years. The regression estimate is

$$\text{Prob(Freeze)} = \underset{(0.0501)}{0.5733} - \underset{(0.0089)}{0.0399} * \text{Inflation rate} - \underset{(0.0153)}{0.0021} * \text{Change in State Unemployment Rate}$$

What is the effect of nominal wage rigidity on the wage growth rates of stayers? The Method 1 estimates in column 3 and 4 of table VI imply that the actual nominal wage change for stayers exceeds notional wage growth $w_t^*- w_{t-1}$ by 0.0469 on average. This average difference seems large us and may not be precisely estimated, but it should be kept in mind that the difference in growth does not accumulate over time. In keeping with the model, the size of the gap is a negative function of the inflation rate, which reflects the fact that freezes are less common when inflation is high. The estimates based on Method 2 in columns 8 and 9 follow the same pattern, but the difference between the actual wage growth and notional wage growth is smaller and less sensitive in absolute terms to inflation. This reflects the fact that the probability of a freeze is much smaller when we use this estimation strategy and cuts are more common. Because of the variation in the estimates, we do not draw strong conclusions about the size of the effect of downward rigidities on average wage changes of stayers.

VII Real Effects of Nominal Rigidity on Layoffs, Quits, Promotions, and Relative wages

A. Effects of Nominal Rigidity on Layoffs, Quits, and Promotions:

In this section we test whether layoffs, quits, and promotions that occur between t-1 and t are affected by the amount the worker is going to be overpaid at t because of downward nominal rigidity. One might expect that the more overpaid a worker will be at t, the more likely his employer will lay him off between t-1 and t. The worker should also be less likely to quit if he is overpaid on the current job. The effect of downward rigidity on promotions is less clear cut but it has been suggested that employers use position changes to adjust wages when wages within positions are rigid. In this case, the degree of rigidity would affect the probability of promotion.

In Table VIII we estimate separate linear probability models for the probability of a layoff, a quit, and a promotion. The right hand side variables are $E(w^0_{it} - w_{it}^*)$ and a set of control variables consisting of education, tenure at t-1, tenure squared at t-1, experience at t, experience squared at t, married, south at t-1, change in state unemployment rate at t, and a full set of year dummies. We suppress the coefficients on the control variables. Equations estimated using the combined sample have additional controls for female, white, and union member. In some specifications, we replace $E(w^0_{it} - w_{it}^*)$ with the estimated probability of a nominal wage freeze for the worker. The $E(w^0_{it} - w_{it}^*)$ variable and the probability of freeze variable are calculated using estimates from the general model of wage rigidity. Each cell of the table represents a separate regression. We report the coefficient (standard error) on the variable of interest for the specified dependent variable, sample, and estimation method.

The results of the estimation differ considerably depending on whether Method 1 or Method 1I is used in the estimation of the wage model. The Method 1 results suggest only a small effect of nominal rigidity on quits. The coefficient on $E(w^0_{it} - w_{it}^*)$ is -.087 (.098) in the combined sample (Panel A) and -

.162 (.192) in the nonunion white men sample (Panel C), but neither is statistically significant. Prob(freeze) enters negatively and is statistically significant at the .01 level in the combined sample. The coefficients in the layoff equation are mixed in sign, very small in magnitude and never statistically significant.[20] The signs of the coefficients on the rigidity measures are mixed in the promotion regressions but are positive and significant at the .10 level in the case of nonunion white men (Panel C).

When Method 2 is used, the coefficients in the quit regressions are always negative and are statistically significant when Prob(freeze) is used as the measure of wage rigidity. They are also very large. For example, for the combined sample, the 10th percentile of $E(w^0_{it} - w_{it}^*)$ is 0.0013 and the 90th percentile is 0.0101. Thus, the coefficient of -3.6 in the specification with the time trend suggests that if a worker changed from being at the 10th percentile to being at the 90th percentile, the probability of quitting would fall by about 0.03 (3.6 · .009). Given that the mean quit probability is about 0.10, this amounts to a substantial effect of nominal rigidities on quit probabilities. The estimates for layoffs from Method 1 are smaller in magnitude and mixed in sign. The promotion estimates are generally insignificant but the signs suggest that workers who are overpaid are less likely to be promoted.

Given the fact that the magnitude of the estimates differ greatly between Method 1 and Method 2, we have not learned that much about the magnitude of the effect of nominal rigidities on quits and the other transitions. However, given the agreement in the signs across the quit specifications, and statistical significance in many cases, there is modest support for the hypothesis that when workers are relatively overpaid because of nominal rigidities, they are less likely to quit. This is in keeping with our expectations and implies that downward nominal rigidity does have real effects on labor market outcomes.

B. Effects of Inflation on the Wage Structure:

Partition x_t into (x_{1t}, π_t) where π_t is the inflation rate and x_{1t} consists of the other wage growth determinants. The notional wage change equation in (1) can be rewritten as

$$w_t^* - w_{t-1}^* = (x_{1t} - x_{1t-1})b_1 + p_t b_{t-2} + e_t - e_{t-1} \ .$$

The model of nominal wage rigidity implies that holding the inflation rate constant, the wage growth of workers with larger values of $x_t\beta$ are less likely to be constrained by nominal wage floors. This means that nominal wage floors tend to compress differences in within job wage growth. Furthermore, the

[20] Standard errors are probably understated because we have not taken account of the fact that the variable of interest is a predicted value or for heteroskedasticity in the linear probability models.

higher the inflation rate, the lower the probability that the wage floor binds for any given value of $x_t\beta$ and the smaller the degree of compression. This reasoning implies that the interaction between π_t and the index $z_t b_1$ will enter negatively in a within job wage equation of the form

$$w_t - w_{t-1} = z_t b_1 + p_t b_2 + (z_t b_1 \cdot p_t) \cdot b_3 + n_t$$

where z_t contains current and lagged values of labor market variables and person and job match specific variables that predict wage growth within a firm and ν_t is a composite error term. In the table below we report nonlinear least squares estimates of b_3 for three samples. The first is the sample of hourly nonunion white males used elsewhere in the paper. The second is combined sample consisting of observations on hourly males and females for 1976-1992. The third consists of hourly or salaried males and females for whom we have reported wage data from 1971-1992, which permits us to maximize sample size as well variation in the inflation rate. (We do not have data on salaried workers prior to 1975 and lack data on married females prior to 1976.) We measure inflation as the change in the log of the price index.

The Effect of Inflation on the Response of Actual Wage Growth to Predicted Wage Growth (b_3)	
Sample	*Estimate (standard error) of b_3*
Nonunion White Males, Hourly, 1971-1992	2.35 (4.72)
Males and Females, hourly 1976-1992.	-3.98 (1.48)
Male and Female Household heads, hourly 1971-1974, hourly and salaried 1975-1992, Married Females, hourly and salaried 1976-1992.	4.0 (4.8)

The point estimate of 4.0 for in the third row for the largest sample implies that an increase in the log price change by .10 would increase a .05 difference in the log wage change when inflation is 0 to about .07, a large effect. However, the estimates of b_3 vary in sign across the samples and are imprecisely estimated. Consequently, we are not able to draw conclusions about whether nominal wage rigidity interacts with inflation to effect relative wage growth.

VI Conclusions and Future Research

In this paper we explore the prevalence of downward nominal wage rigidity in the U.S. economy. We do so using two very different types of wage data. The first type, personnel data, is probably free of measurement error data but is not nationally representative. The second type is wage data from the PSID, which is nationally representative but contains measurement error.

We approach the PSID data by estimating a model of wage determination that nests downward nominal rigidity as well as flexible wages and directly incorporates a model of measurement error in wage reports. In keeping with the evidence from replications studies, we find that measurement error in wages is better described by a mixed distribution in which respondents sometimes report wages accurately than by the more standard assumption of normally distributed measurement error. However, we should point out that we do not take account of rounding in the data or possibility that measurement error is correlated with the actual wage.

The PSID data overwhelmingly rejects the hypothesis of perfect flexibility. There is also some evidence against the hypothesis of perfect downward nominal wage rigidity in favor of a more general model that incorporates downward rigidity but acknowledges that some nominal wage cuts may occur, although many specifications of the general model imply that the probability of a nominal wage cut is essentially zero. We find that true wage changes have many fewer nominal cuts than reported wage changes have, with our point estimates ranging from essentially 0 for our preferred estimation methodology to .035. We also find more nominal freezes than reported nominal wage changes. Unfortunately, the estimates range widely, from .15 to .43, depending on the estimation methodology and sample.

Our analysis of the wage and salary data from personnel files of a large financial corporation also suggests that reductions in nominal wages are quite rare, especially for salaried workers. Changes of zero in compensation are fairly common in the firm and wage cuts are infrequent. These findings suggest that downward nominal wage rigidity is present. Bonuses serve to reduce nominal rigidity, but nominal reductions in base salary or hourly pay that do not involve a change in part time status or a change in performance incentives almost never occur in this firm. An interesting subject for future research is whether the use of bonuses is positively related to the probability that a nominal wage floor is binding, such a period of low inflation. Our finding that nominal wage cuts are rare is consistent with the findings of Baker, Gibbs, and Holmstrom (1994), who also use personnel data. Further work with this kind of data for a broad set of firms is needed to provide more precise information about the frequency of cuts and freezes, the role of rounding and symmetric menu costs, and the circumstances in which cuts and freezes occur than is possible with noisy data from household surveys.

The cumulative evidence from previous research and the present paper constitutes a strong case that nominal wage rigidity is pervasive in the U.S. labor market. Much less is known about whether nominal wage rigidity matters. We obtain modest support for the hypothesis that workers who are overpaid because of nominal wage rigidities are less likely to quit. However, our analysis of the effects of nominal wage rigidity on layoffs, promotions, and relative wage growth rates is inconclusive. Research on whether nominal wage rigidity has real effects on employment, mobility patterns, and relative wages deserves a high research priority.

Finally, a more structural approach to the study of wage rigidity is needed. In an environment in which it is costly to cut nominal wages, firms should take account of the possibility of future constraints when setting current wages. A structural model of wage setting behavior in presence of nominal wage rigidity might be a useful tool with which to study wage and employment dynamics and the effects of inflation on the labor market.

Appendix 1: Calculating E(w^0_{it} - w_{it}*)

Consider model (5) once more.

$$
\begin{aligned}
w_{it} - w_{it-1} &= x_{it} b - \hat{w}_{it-1} + e_{it} + m_{it} && if \quad x_{it} b - \hat{w}_{it-1} + e_{it} \geq 0 \\
&= m_{it} && if \quad -a \leq x_{it} b - \hat{w}_{it-1} + e_{it} \leq 0 \\
&= l + x_{it} b - \hat{w}_{it-1} + e_{it} + m_{it} && if \quad x_{it} b - \hat{w}_{it-1} + e_{it} < -a
\end{aligned}
$$

As before, $w_{it} * = x_{it} b + e_{it}$

Together these imply that

$$
\begin{aligned}
w_{it} - w_{it} * &= u_{it} && if \quad x_{it} b - \hat{w}_{it-1} + e_{it} \geq 0 \\
&= u_{it} - (x_{it} b - \hat{w}_{it-1} + e_{it}) && if \quad -a \leq x_{it} b - \hat{w}_{it-1} + e_{it} \leq 0 \\
&= l + u_{it} && if \quad x_{it} b - \hat{w}_{it-1} + e_{it} < -a
\end{aligned}
$$

Using the fact that the expected value of u_{it} is zero and the assumption of a zero covariance between u_{it} and e_{it}, one arrives at the following expression, expression (6)

$$
\begin{aligned}
&E(w^0{}_{it} - w_{it} *) = -E(x_{it} b - \hat{w}_{it-1} + e_{it} \mid -a \leq x_{it} b - \hat{w}_{it-1} + e_{it} \leq 0) \Pr ob(-a \leq x_{it} b - \hat{w}_{it-1} + e_{it} \leq 0) + \\
&l \Pr ob(x_{it} b - \hat{w}_{it-1} + e_{it} \leq -a)
\end{aligned}
\qquad (6)
$$

The terms in the expression can be further defined as follows:

$$
E(x_{it} b - \hat{w}_{it-1} + e_{it} \mid -a \leq x_{it} b - \hat{w}_{it-1} + e_{it} \leq 0) = x_{it} b - \hat{w}_{it-1} - s_e \frac{A}{B}
$$

$$
\Pr ob(-a \leq x_{it} b - \hat{w}_{it-1} + e_{it} \leq 0) = B
$$

$$
A = j\left(\frac{\hat{w}_{it-1} - x_{it} b}{s_e}\right) - j\left(\frac{\hat{w}_{it-1} - x_{it} b - a}{s_e}\right)
$$

$$B = \Phi\left(\frac{\hat{w}_{it-1} - x_{it} b}{s_e} \right) - \Phi\left(\frac{\hat{w}_{it-1} - x_{it} b - a}{s_e} \right)$$

$$\Pr ob(x_{it} b - \hat{w}_{it-1} + e_{it} \leq -a) = \Phi\left(\frac{\hat{w}_{it-1} - x_{it} b - a}{s_e} \right)$$

Ignoring selection issues, for any group of workers who are employed at t-1, expression (6) can be used to predict what the difference between the notional wage and the actual wage at t will be if they stay with the same employer. In order to evaluate the expression we need to know the value of the x variables at t for all workers, including those who do not stay with the same employer. The value of the inflation rate and the state unemployment at t do not pose a problem because they are known for all workers and do not depend on whether or not the worker changes jobs. For tenure and experience, we take the value of tenure and experience at t-1 and add 1 in order to have the values of these variables assuming the workers stay with the same employer. We assume that union status would not change between t-1 and t if the worker were to stay in the same job until t. In this way we get a set of x values at t for all hourly workers who are employed at t-1. We then use equation (6) to estimate the value of $E(w^0_{it} - w_{it}*)$ for each worker under the assumption that they stay with the employer they have at t-1.

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Table I: Descriptive Statistics for the Sample of Nonunion white males (4485 Observations)

Variable	Mean	Standard Deviation	Minimum	Maximum
Log of nominal hourly wage	2.0123	0.4422	0.4055	4.1744
change in the log of hourly wage	0.0746	0.0910	-0.2231	0.5108
Years of Education	11.7939	2.1826	2.5	17
Tenure	8.0290	6.9073	1	42.2440
Experience	18.4964	10.8935	3	47.3333
Married	0.8832	0.3213	0	1
South	0.3579	0.4794	0	1
Year	83.7652	5.6910	72	92
Log of price level	4.4334	0.3107	3.7087	4.8251
Change in Log of Price Level	0.0527	0.0191	0.0296	0.0983
Change in State Unemployment rate	0.0698	1.3782	-5.9333	6.4333
State Unemployment rate	6.6898	2.0211	2.2667	15.2
Indicator for nominal wage cut	0.0883	0.2838	0	1
Indicator for nominal wage freeze	0.13601	0.3428	0	1

Descriptive Statistics for the Combined Sample (15234 Observations)

Variable	Mean	Standard Deviation	Minimum	Maximum
Log of nominal hourly wage	2.0388	0.45515	0	4.576668
change in the log of hourly wage	0.0703	0.089932	-0.2551	0.448025
Years of Education	11.8948	2.078078	2	17
Tenure	8.9540	7.610207	1.083333	45.6916
Experience	18.1277	10.16682	2	47.5
Married	0.8222	0.382323	0	1
Female	0.4177	0.493194	0	1
White	0.8861	0.317688	0	1
Union member	0.3543	0.478308	0	1
South	0.2966	0.456791	0	1
Year	84.9183	4.695698	76	92
Log of price level	4.5063	0.235794	3.975936	4.825109
Change in Log of Price Level	0.051381	0.018662	0.029569	0.098345
Change in State Unemployment rate	0.022489	1.394761	-5.93333	6.933333
State Unemployment rate	6.998659	2.116549	2.266667	15.2
Indicator for nominal wage cut	0.111461	0.314713	0	1
Indicator for nominal wage freeze	0.117435	0.321948	0	1

Table II : Estimates from Wage Models using Combined Sample (Method 1)

Variable	General Model	$\lambda = 0$	$\lambda = \alpha$ (MM model)	Rigid Model (no cuts)
Constant	-2.7551 (0.0290)	-2.7378 (0.0290)	-2.7096 (0.0284)	-2.7684 (0.0288)
Education	0.0604 (0.0007)	0.0604 (0.0007)	0.0603 (0.0007)	0.0606 (0.0007)
Tenure	0.0185 (0.0009)	0.0196 (0.0009)	0.0233 (0.0009)	0.0185 (0.0009)
Tenure Squared/100	-0.0409 (0.0022)	-0.0409 (0.0022)	-0.0436 (0.0021)	-0.0418 (0.0022)
Married	0.0226 (0.0036)	0.0248 (0.0036)	-0.0029 (0.0035)	0.0176 (0.0036)
Experience	0.0210 (0.0008)	0.0205 (0.0008)	0.0154 (0.0008)	0.0208 (0.0008)
Experience Squared/100	-0.0322 (0.0014)	-0.0320 (0.0014)	-0.0285 (0.0014)	-0.0313 (0.0014)
Log of Price Level	1.2259 (0.0371)	1.1641 (0.0371)	1.2084 (0.0364)	1.1627 (0.0370)
south	-0.0381 (0.0031)	-0.0381 (0.0031)	-0.0365 (0.0030)	-0.0368 (0.0031)
state unemployment rate	-0.0016 (0.0007)	-0.0014 (0.0007)	-0.0019 (0.0007)	-0.0008 (0.0007)
year	-0.0227 (0.0019)	-0.0197 (0.0019)	-0.0215 (0.0019)	-0.0193 (0.0019)
union member	0.2486 (0.0032)	0.2484 (0.0032)	0.2476 (0.0032)	0.2489 (0.0032)
female	-0.2612 (0.0058)	-0.2600 (0.0058)	-0.2663 (0.0057)	-0.2624 (0.0058)
white	0.1850 (0.0079)	0.1871 (0.0079)	0.1492 (0.0077)	0.1883 (0.0078)
female * tenure	0.0044 (0.0006)	0.0046 (0.0006)	0.0012 (0.0006)	0.0046 (0.0006)
female * experience	-0.0038 (0.0004)	-0.0039 (0.0004)	-0.0023 (0.0004)	-0.0038 (0.0004)
white * tenure	0.0027 (0.0007)	0.0014 (0.0007)	-0.0007 (0.0007)	0.0030 (0.0007)
white * experience	-0.0043 (0.0005)	-0.0038 (0.0005)	-0.0004 (0.0005)	-0.0046 (0.0005)
σ_e	0.1129 (0.0013)	0.1131 (0.0013)	0.1120 (0.0013)	0.1129 (0.0013)

Table II, continued

σ_{u*}	0.0697 (0.0008)	0.0695 (0.0008)	0.0702 (0.0009)	0.0697 (0.0008)
α	0.6536 (0.1703)	0.6135 (0.6340)	0.3193 (0.0171)	
λ	0.0442 (2094.8934)			
p	0.5825 (0.0067)	0.5804 (0.0067)	0.5855 (0.0068)	0.5823 (0.0067)
Prob (freeze)	0.4303	0.4314	0.4188	0.4296
Prob (cut)	0.00001	0.00003	0.0087	0
log(L)	7015	7011	6914	7010

Table III: Estimates from Wage Models using Combined Sample (Method 2)

Variable	General Model	$\lambda = 0$	$\lambda = \alpha$ (MM model)	Rigid Model (No cuts)
Constant	0.2497 (0.0179)	0.2209 (0.0174)	0.2497 (0.0178)	0.2192 (0.0172)
Education	0.0012 (0.0003)	0.0012 (0.0003)	0.0012 (0.0003)	0.0013 (0.0003)
Tenure	-0.0029 (0.0003)	-0.0025 (0.0003)	-0.0029 (0.0003)	-0.0025 (0.0003)
Tenure Squared/100	0.0068 (0.0009)	0.0061 (0.0009)	0.0068 (0.0009)	0.0061 (0.0009)
Change in marital Status	0.0003 (0.0036)	-0.0003 (0.0034)	0.0003 (0.0036)	-0.0011 (0.0034)
Experience	-0.0011 (0.0003)	-0.0009 (0.0003)	-0.0011 (0.0003)	-0.0009 (0.0003)
Experience Squared/100	0.0016 (0.0007)	0.0010 (0.0006)	0.0016 (0.0007)	0.0011 (0.0006)
Inflation rate	0.6448 (0.0475)	0.6254 (0.0457)	0.6402 (0.0473)	0.6311 (0.0456)
south	0.0010 (0.0015)	0.0016 (0.0014)	0.0010 (0.0015)	0.0016 (0.0014)
change in state unemployment rate	0.0023 (0.0005)	0.0017 (0.0005)	0.0022 (0.0005)	0.0017 (0.0005)
year	-0.0025 (0.0002)	-0.0022 (0.0002)	-0.0025 (0.0002)	-0.0022 (0.0002)
Union member	0.0048 (0.0016)	0.0085 (0.0016)	0.0048 (0.0016)	0.0083 (0.0016)
female	0.0047 (0.0014)	0.0045 (0.0014)	0.0047 (0.0014)	0.0045 (0.0014)
white	-0.0003 (0.0020)	0.0002 (0.0020)	-0.0003 (0.0020)	0.0019 (0.0020)
σ_e	0.0619 (0.0007)	0.0581 (0.0006)	0.0617 (0.0007)	0.0581 (0.0006)
σ_{u*}	0.1255 (0.0030)	0.1111 (0.0017)	0.1253 (0.0031)	0.1111 (0.0017)
α	0.0596 (0.0000)	0.2159 (11.89)	0.0597 (0.0025)	
λ	0.0590 (0.0033)			

Table III, continued

p	0.8497 (0.0059)	0.7931 (0.0046)	0.8487 (0.0067)	0.7933 (0.0046)
Prob (freeze)	0.1493	0.1689	0.1491	0.1689
Prob (cut)	0.0358	0	0.0353	0
log(L)	9201	9113	9193	9113

Table IV: Estimates from Wage Models using White Non-union Male Sample (Method 1)

Variable	λ = 0*	$\lambda = \alpha$ (MM model)	Rigid Model (No cuts)
Constant	-2.3152	-2.3603 (0.0389)	-2.3437 (0.0391)
Education	0.0538	0.0538 (0.0011)	0.0536 (0.0011)
Tenure	0.0263	0.0263 (0.0010)	0.0264 (0.0010)
Tenure Squared/100	-0.0491	-0.0486 (0.0035)	-0.0493 (0.0035)
Married	0.0897	0.0927 (0.0067)	0.0913 (0.0067)
Experience	0.0216	0.0212 (0.0009)	0.0215 (0.0009)
Experience Squared/100	-0.0475	-0.0469 (0.0020)	-0.0474 (0.0020)
Log of Price Level	1.4826	1.3417 (0.0518)	1.3932 (0.0520)
south	-0.0688	-0.0680 (0.0047)	-0.0690 (0.0047)
state unemployment rate	-0.0008	0.0002 (0.0012)	0.0013 (0.0012)
year	-0.0395	-0.0316 (0.0029)	-0.0346 (0.0029)
σ_e	0.1096	0.1093 (0.0021)	0.1095 (0.0021)
σ_{u*}	0.0674	0.0678 (0.0019)	0.0675 (0.0019)
α	1.3573	0.4417 (0.5004)	
λ			
p	0.6653	0.6678 (0.0126)	0.6665 (0.0125)
Prob (freeze)	0.3817	0.3792	0.3804
Prob (cut)	0.0000	0.0002	0.0000
log(L)	2007	2003	2005

* Could not compute standard errors because of singularities

Table V: Estimates from Wage Models using White Non-union Male Sample (Method 2)

Variable	General Model	$\lambda = 0$	$\lambda = \alpha$ (MM model)	Rigid Model (No cuts)
Constant	0.2322 (0.0253)	0.2299 (0.0247)	0.2363 (0.0252)	0.2300 (0.0247)
Education	0.0014 (0.0006)	0.0016 (0.0006)	0.0014 (0.0006)	0.0016 (0.0006)
Tenure	-0.0043 (0.0006)	-0.0039 (0.0006)	-0.0043 (0.0006)	-0.0040 (0.0006)
Tenure Squared/100	0.0112 (0.0020)	0.0103 (0.0020)	0.0111 (0.0020)	0.0108 (0.0020)
Change in marital Status	0.0034 (0.0055)	0.0044 (0.0053)	0.0032 (0.0055)	0.0037 (0.0053)
Experience	-0.0011 (0.0005)	-0.0013 (0.0005)	-0.0011 (0.0005)	-0.0009 (0.0005)
Experience Squared/100	0.0015 (0.0011)	0.0002 (0.0010)	0.0014 (0.0011)	0.0008 (0.0010)
Inflation rate	0.7636 (0.0774)	0.7023 (0.0749)	0.7496 (0.0769)	0.7077 (0.0750)
south	0.0032 (0.0025)	0.0042 (0.0025)	0.0030 (0.0025)	0.0026 (0.0025)
change in state unemployment rate	0.0001 (0.0009)	-0.0001 (0.0009)	0.0001 (0.0009)	0.0001 (0.0009)
year	-0.0023 (0.0003)	-0.0023 (0.0003)	-0.0023 (0.0003)	-0.0023 (0.0003)
σ_e	0.0633 (0.0011)	0.0609 (0.0010)	0.0631 (0.0011)	0.0609 (0.0010)
σ_{u*}	0.1362 (0.0059)	0.1216 (0.0036)	0.1351 (0.0059)	0.1214 (0.0036)
α	0.0725 (0.0000)	0.2213 (0.0001)	0.0734 (0.0053)	
λ	0.0715 (0.0064)			
p	0.8775 (0.0086)	0.8396 (0.0072)	0.8752 (0.0093)	0.8391 (0.0072)
Prob (freeze)	0.1601	0.1746	0.1602	0.1748
Prob (cut)	0.0254	0.0000	0.0242	0.0000
log(L)	2585	2561	2582	2561

Table VI: Predicted Cuts, Predicted Freezes, Reported Wage Changes, and Actual Wage Changes by Year for Combined Sample

			Method 1				Method 2			
Year (1)	Inflation Rate (2)	Mean Reported Wage Change (3)	Mean Notional Wage Change (4)	Predicted Prob. of Cut (5)	Predicted Prob. of Freeze (6)	$E(w^0_{it} - w_{it}^*)$ (7)	Mean Notional Wage Change (8)	Predicted Prob. of Cut (9)	Predicted Prob. of Freeze (10)	$E(w^0_{it} - w_{it}^*)$ (11)
76	5.3963	0.0925	0.0779	0	0.2705	0.0236	0.0818	0.0145	0.0925	0.0029
77	6.1838	0.0977	0.0806	0	0.2714	0.0235	0.0853	0.0123	0.0834	0.0026
78	6.8169	0.0875	0.0365	0	0.3857	0.0366	0.0867	0.0115	0.08	0.0024
79	8.0676	0.0985	0.055	0	0.3323	0.0306	0.0936	0.0084	0.0652	0.0019
80	9.8345	0.1067	0.0866	0	0.2416	0.0189	0.1083	0.0042	0.041	0.0011
81	8.3216	0.099	0.0809	0	0.2688	0.0233	0.0954	0.0077	0.0617	0.0018
82	5.3066	0.087	0.0498	0	0.3538	0.0354	0.0777	0.0169	0.1024	0.0033
83	4.1207	0.0644	0.0222	0	0.4287	0.0396	0.0662	0.0277	0.1379	0.0049
84	3.6244	0.0618	0.002	0	0.4918	0.0502	0.0541	0.0458	0.1829	0.0072
85	3.7122	0.0626	0.006	0	0.4809	0.0501	0.0569	0.0407	0.1719	0.0066
86	2.9569	0.0448	-0.0244	0	0.5715	0.066	0.0505	0.0526	0.1975	0.008
87	3.9781	0.053	0.0012	0	0.496	0.055	0.0529	0.0477	0.1874	0.0074
88	4.2101	0.056	-0.0032	0	0.5072	0.0563	0.0519	0.0497	0.1917	0.0076
89	4.8653	0.0606	0.0165	0	0.4472	0.0443	0.054	0.0455	0.1829	0.0071
90	5.0745	0.0588	0.0115	0	0.4645	0.0463	0.0539	0.0458	0.1835	0.0072
91	4.3319	0.0611	0.0018	0	0.4942	0.05	0.0505	0.0527	0.1973	0.008
92	3.512	0.0514	-0.0191	0	0.5615	0.0609	0.0425	0.0722	0.2317	0.0101
Overall	5.1381	0.0703	0.0234	0	0.4303	0.0437	0.0657	0.0358	0.1493	0.0057

Notes: The results in columns 4-7 are means by year of estimates of the general model using Method 1 (Table 2). The results in columns 8-11 are based on estimates of the general model using Method 2 (Table 3). See Section VI D.

Table VII: Predicted Cuts, Predicted Freezes, Reported Wage Changes, and Actual Wage Changes by Year for the Non-union White Male Sample

			Method 1				Method 2			
Year (1)	Inflation Rate (2)	Mean Reported Wage Change (3)	Mean Notional Wage Change (4)	Predicted Prob. of cut (5)	Predicted Prob. of Freeze (6)	$E(w^0_{it} - w_{it}^*)$ (7)	Mean Notional Wage Change (8)	Predicted Prob. of Cut (9)	Predicted Prob.of Freeze (10)	$E(w^0_{it} - w_{it}^*)$ (11)
72	3.238	0.0877	0.0675	0	0.2948	0.0241	0.0717	0.0165	0.1374	0.0048
73	4.552	0.0709	0.0891	0.0001	0.2525	0.0223	0.08	0.0114	0.1105	0.0037
74	8.955	0.1117	0.1331	0	0.1541	0.0109	0.1095	0.0028	0.0463	0.0013
75	7.823	0.1174	0.126	0	0.1669	0.0114	0.0997	0.0045	0.0625	0.0018
76	5.3963	0.0961	0.0659	0	0.2906	0.0214	0.0857	0.0091	0.0952	0.0031
77	6.1838	0.0983	0.0654	0.0001	0.3041	0.0261	0.0892	0.0077	0.0859	0.0027
78	6.8169	0.106	0.044	0.0001	0.3585	0.0314	0.0934	0.0063	0.076	0.0023
79	8.0676	0.109	0.0626	0	0.3003	0.0231	0.1001	0.0046	0.0624	0.0018
80	9.8345	0.1202	0.0826	0	0.2403	0.0169	0.1098	0.0028	0.0455	0.0013
81	8.3216	0.0957	0.0833	0	0.2515	0.0181	0.0958	0.0054	0.0699	0.0021
82	5.3066	0.0856	0.037	0	0.3895	0.0332	0.0739	0.0145	0.1279	0.0044
83	4.1207	0.0561	0.046	0	0.3653	0.0308	0.0634	0.0229	0.1662	0.0062
84	3.6244	0.063	0.0175	0.0001	0.4545	0.044	0.0577	0.0293	0.1899	0.0074
85	3.7122	0.0689	0.024	0.0003	0.429	0.0434	0.0558	0.0319	0.1984	0.0079
86	2.9569	0.0445	-0.0027	0.0004	0.503	0.0504	0.0493	0.0415	0.2279	0.0095
87	3.9781	0.0556	0.0139	0.0003	0.4492	0.0433	0.0525	0.0364	0.213	0.0087
88	4.2101	0.0582	0.019	0.0002	0.4341	0.0396	0.0526	0.0362	0.2125	0.0086
89	4.8653	0.0613	0.0308	0.0002	0.3928	0.0346	0.0546	0.0334	0.2035	0.0081
90	5.0745	0.0601	0.0313	0.0002	0.3911	0.0336	0.0542	0.0341	0.2054	0.0083
91	4.3319	0.0554	0.0114	0.0002	0.4576	0.0412	0.0483	0.0442	0.2335	0.0099
92	3.512	0.0439	-0.0211	0.0007	0.5669	0.0585	0.0411	0.0593	0.2693	0.0123
Overall	5.2707	0.0746	0.0397	0.0002	0.3792	0.0341	0.0688	0.0254	0.1601	0.0063

Notes: The results in columns 4-7 are means by year of estimates of the MM model using Method 1 (Table 4). The results in columns 8-11 are based on estimates of the general model using Method 2 (Table 5). See Section VI D.

Table VIII: The Effects of the Probability of a Wage Freeze and $E(w_t^0 - w_t^*)$ on Quits, Layoffs, and Promotions. Linear Probability Estimates (Standard Errors in Parentheses).

		Dependent Variable		
Sample, Treatment of endogeneity of wt-1, model of wage rigidity	Wage Rigidity Measure:	Quit	Layoff	Promo-tion
A. Combined Sample, Method 1, General Model	$E(w_t^0 - w_t^*)$	-0.0862 (0.0982)	-0.0055 (0.0550)	-0.0235 (0.0656)
	Prob (freeze)	-0.0374*** (0.0153)	-0.0084 (0.0085)	0.0005 (0.0102)
B. Combined Sample, Method 2, General Model.	$E(w_t^0 - w_t^*)$	-3.5691 (2.6559)	2.6834* (1.4872)	-2.5851 (1.7752)
	Prob (freeze)	-0.6675*** (0.2288)	0.1939 (0.1282)	-0.2108 (0.1530)
C. Nonunion white men, Method 1, Malcolmson and MacLeod Model	$E(w_t^0 - w_t^*)$	-0.1619 (0.1922)	-0.0003 (0.1197)	0.1728* (0.1207)
	Prob (Freeze)	-0.0256 (0.0286)	0.0027 (0.0178)	0.0318* (0.0180)
D. Nonunion white men, Method 2, General Model	$E(w_t^0 - w_t^*)$	-5.4599 (3.5954)	-0.6039 (2.2383)	-2.2212 (2.2602)
	Prob (freeze)	-0.6782** (0.3072)	-0.1267 (0.1913)	-0.2583 (0.1931)

Notes: The coefficients (standard errors) in the table are linear probability coefficients from separate regressions of the dependent variable in the column on $E(w^0_{it} - w_{it}*)$ or Prob(freeze) and a set of controls. $E(w^0_{it} - w_{it}*)$ is the expectation of the gap between the actual wage and the notional wage given the model parameters and the characteristics of worker i in period t. The creation of this variable is described in appendix 1. Prob(freeze) is the conditional probability of a wage freeze. All equations control for year dummies and for education, tenure, tenure squared, experience, experience squared, change in marital status, change in state unemployment rate, and residence in the south. Equations estimated using the combined sample have additional controls for female, white, and union member. The sample size is 19841 for combined sample and 6134 for the nonunion white men sample. The symbols *, **, and *** denote statistical significance at the .10, .05, and .01 level on the basis of a two-tailed test.

Appendix Tables

Table A1: REGRESSIONS USED TO CREATE $\hat{w}_{t-1}$

	SMALL SAMPLE		COMBINED SAMPLE	
	Parameter	Standard	Parameter	Standard
Variable	Estimate	Error	Estimate	Error
INTERCEPT	-3.71411	0.414638	-3.81148	0.764871
Education	0.053684	0.010309	0.110686	0.006122
Tenure at t-1	0.005948	0.009048	0.034051	0.005118
Tenure at t-1 squared	0.000785	0.000336	-0.00057	0.000164
Union member at t-1			0.313941	0.029309
married at t-1	0.03643	0.064861	-0.02211	0.030753
Years of experience at t-1	0.015013	0.008689	0.030913	0.005328
Years of experience at t-1 squared	-0.00043	0.000191	-0.00053	0.000118
Log of price level at t-1	1.584433	0.196958	1.502994	0.282144
South at t-1	0.117196	0.04993	0.172135	0.029028
State Unemployment rate at t-1	0.057668	0.020008	0.035224	0.010861
State Unemployment rate at t-2	0.112079	0.058388	0.036295	0.045164
Inflation rate at t-1	12.78718	8.917637	12.27066	12.15867
Inflation rate at t-2	-4.99105	7.871795	-1.48044	7.746704
Female			-0.09782	0.027641
White			0.261568	0.041613
Inflation rate at t-1 * Education	-0.15233	0.1701	-0.24002	0.109356
Inflation rate at t-1 * tenure at t-1	0.230539	0.162102	-0.04785	0.091465
Inflation rate at t-1 * tenure at t-1 squared	-0.01083	0.00596	-0.00025	0.003057
Inflation rate at t-1 * union member at t-1			-0.38581	0.513628
Inflation rate at t-1 * married at t-1	0.485202	1.168798	0.113039	0.572249
Inflation rate at t-1 * years of experience at t-1	-0.36863	0.141367	-0.24835	0.090887
Inflation rate at t-1 * years of experience at t-1 squared	0.007386	0.003088	0.005683	0.002077
Inflation rate at t-1 * log of price level at t-1	-1.21082	1.922158	-0.74169	2.462397
Inflation rate at t-1 * south at t-1	-1.41131	0.77933	-1.53012	0.497981
Inflation rate at t-1 * state unemployment rate at t-1	-0.68811	0.364525	-0.42076	0.209464
Inflation rate at t-1 * state unemployment rate at t-2	0.526387	0.388167	0.297496	0.244019
Inflation rate at t-1 * inflation rate at t-1	-33.1245	27.45393	-16.9217	20.09195
Inflation rate at t-1 * inflation rate at t-2	46.3455	40.87874	20.24589	39.00604
Inflation rate at t-2 * Education	0.011547	0.160364	-0.34724	0.101167

Table A1 continued

Inflation rate at t-2 * tenure at t-1	-0.04568	0.151336	-0.14397	0.084354
Inflation rate at t-2 * tenure at t-1 squared	-0.00135	0.005491	0.002708	0.002818
Inflation rate at t-2 * union member at t-1			-0.80352	0.476699
Inflation rate at t-2 * married at t-1	0.802646	1.087046	0.131119	0.527249
Inflation rate at t-2 * years of experience at t-1	0.231664	0.13249	0.073915	0.084007
Inflation rate at t-2 * years of experience at t-1 squared	-0.00461	0.002918	-0.0026	0.001921
Inflation rate at t-2 * log of price level at t-1	0.742176	1.618211	1.189473	1.528412
Inflation rate at t-2 * south at t-1	0.312444	0.727527	-0.26624	0.456651
Inflation rate at t-2 * state unemployment rate at t-1	0.015985	0.307743	0.181332	0.199188
Inflation rate at t-2 * state unemployment rate at t-2	-0.34621	0.332731	0.042817	0.224306
Inflation rate at t-2 * inflation rate at t-2	-14.3927	23.42115	-14.276	24.03076
Inflation rate at t-1 * white			-0.87038	0.68078
Inflation rate at t-1 * female			-0.91415	0.488754
Inflation rate at t-2 * white			-0.78849	0.626616
Inflation rate at t-2 * female			-1.18442	0.46007
female * tenure at t-1			0.004427	0.000941
female * years of experience at t-1			-0.00436	0.000676
white * tenure at t-1			0.00404	0.001307
white * years of experience at t-1			-0.00476	0.000977
			0.002083	0.003785
			-0.00624	0.002529
State unemployment rate at t-2 * Education	0.000823	0.001027	-1.3E-05	1.49E-05
State unemployment rate at t-2 * tenure at t-1	0.002375	0.000863	0.002884	0.002784
State unemployment rate at t-2 * tenure at t-1 squared	-0.00011	3.09E-05	0.003891	0.002987
State unemployment rate at t-2 * union member at t-1			0.000157	0.000497
State unemployment rate at t-2 * married at t-1	-0.00011	0.006345	4.28E-06	1.14E-05
State unemployment rate at t-2 * years of experience at t-1	0.002228	0.000848	-0.0065	0.008506
State unemployment rate at t-2 * years of experience at t-1 squared	-3.3E-05	1.91E-05	-0.01578	0.002826
State unemployment rate at t-2 * log of price level at t-1	-0.03246	0.010644	-0.00286	0.000887
State unemployment rate at t-2 * south at t-1	-0.01951	0.004987	0.001593	0.000763

Table A1 continued

State unemployment rate at t-2 * state unemployment rate at t-1	-0.00257	0.001664	-0.00267	0.000592
State unemployment rate at t-2 * state unemployment rate at t-2	0.000508	0.001418	-0.00031	0.000459
Year	-0.03019	0.009079	-0.03562	0.007839
Number of Observations	6134		19841	

Table A2: Estimates from Model where Measurement error is normally distributed
Combined sample. (standard errors in parenthesis).

Variable	General Model	Flexible Model	Rigid Model
σ_e	0.1319 (0.0014)		0.1230 (0.0015)
σ_u	0.0379 (0.0004)		0.0466 (0.0003)
α	0.3248 (0.0067)		
λ	0.1726 (0.0114)		
$\sigma_e^2 + 2\sigma_u^2$		0.0119 (0.0001)	
Prob (freeze)	0.4456	0.0000	0.4789
Prob (cut)	0.0158	0.2888	0.0000
log(L)	14580	12124	14520

Note: See Table II for a list of the variables in the models.

Table A3: Estimates from Model where Measurement error is normally distributed
Non-union white males sample. (standard errors in parentheses)

Variable	General Model	Flexible Model	Rigid Model
σ_e	0.1213 (0.0018)		0.1194 (0.0024)
σ_u	0.0286 (0.0008)		0.0410 (0.0006)
α	0.2337 (0.0090)		
λ	0.1187 (0.0123)		
$\sigma_e^2 + 2\sigma_u^2$		0.0103 (0.0001)	
Prob (freeze)	0.3658	0.0000	0.4330
Prob (cut)	0.0247	0.2635	0.0000
log(L)	4490	3902	4473

Note: See Table IV for a list of the other variables in the models.

Figure 1a

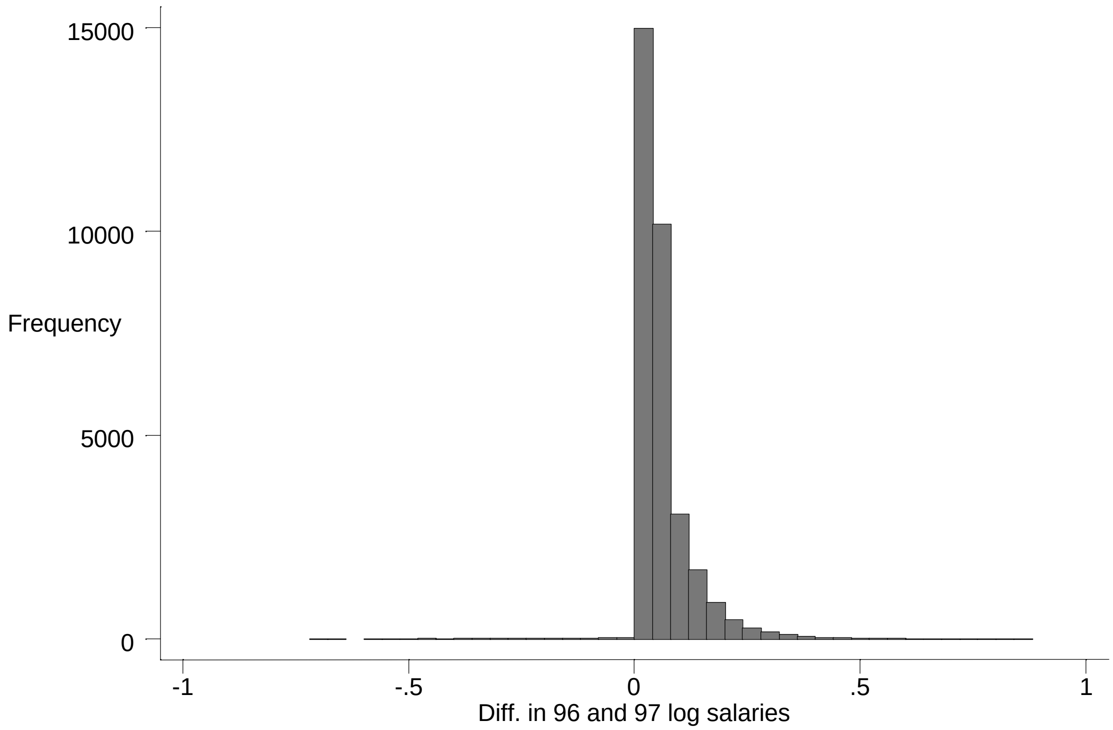

Figure 1a: Frequency of the change in log annual salary, Salary Workers.

Figure 1b

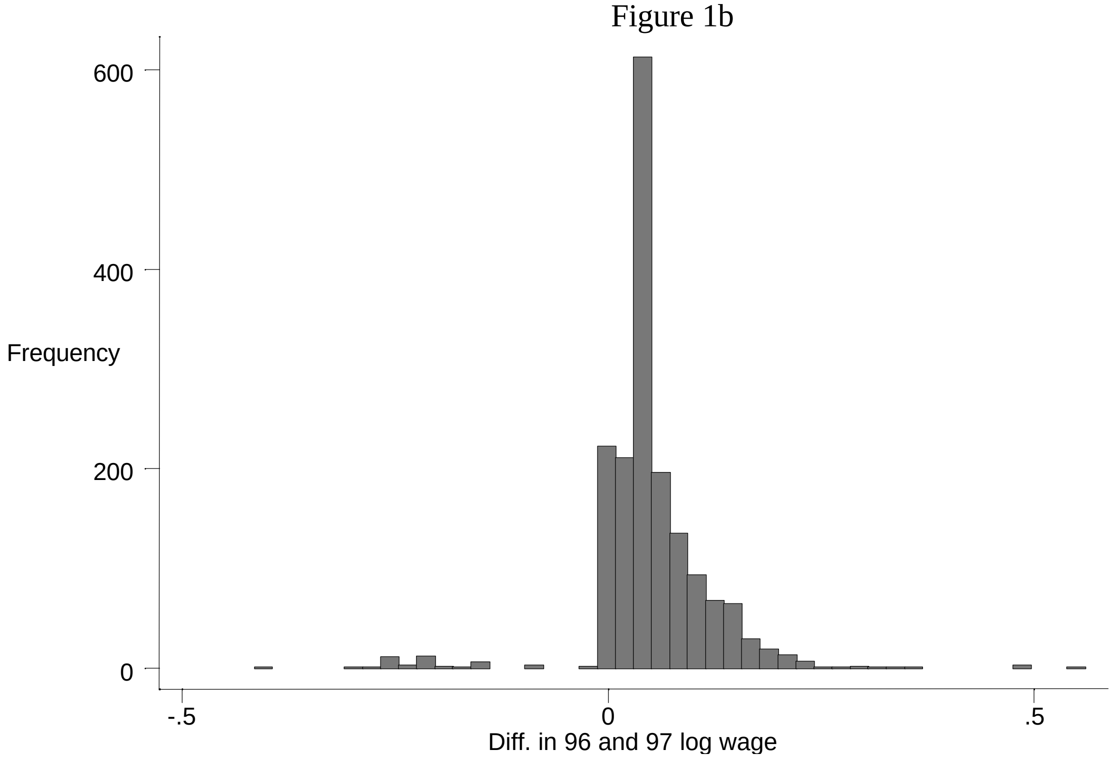

Figure 1b: Frequency of the change in the log hourly wage, Hourly Workers.

Figure 2: The relationship between the Actual wage change and the Notional wage change

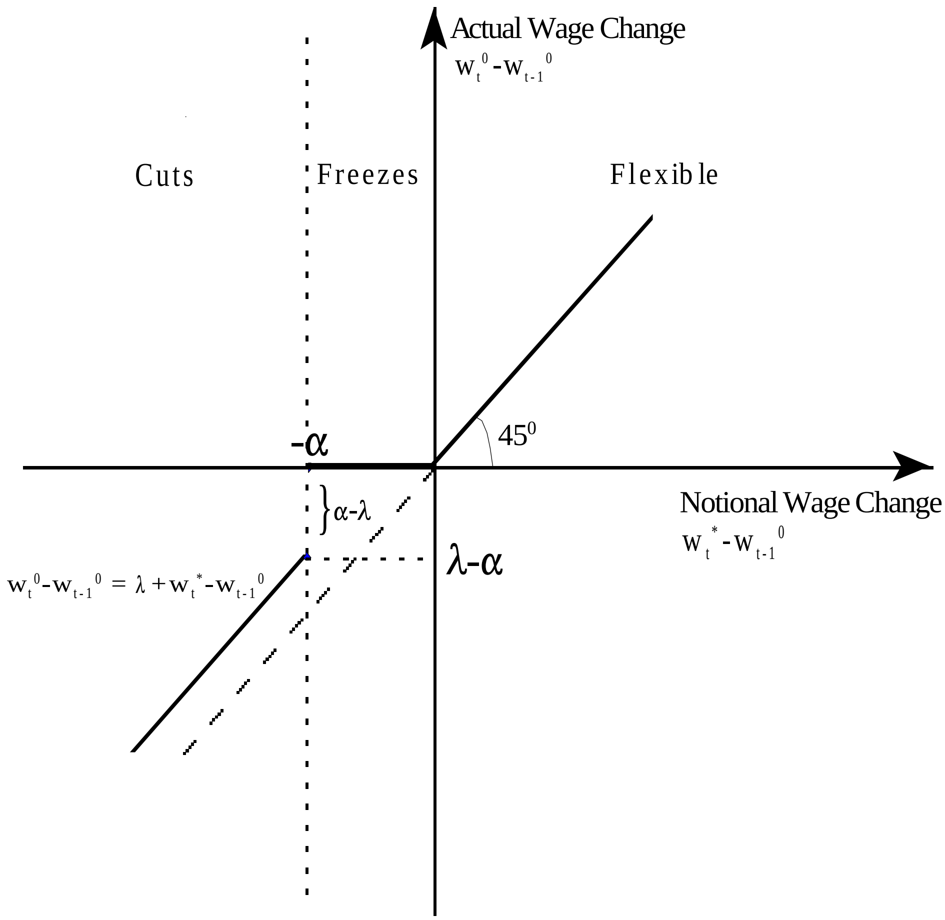

Figure 3: Frequency distribution of nominal wage changes, PSID non-union white males

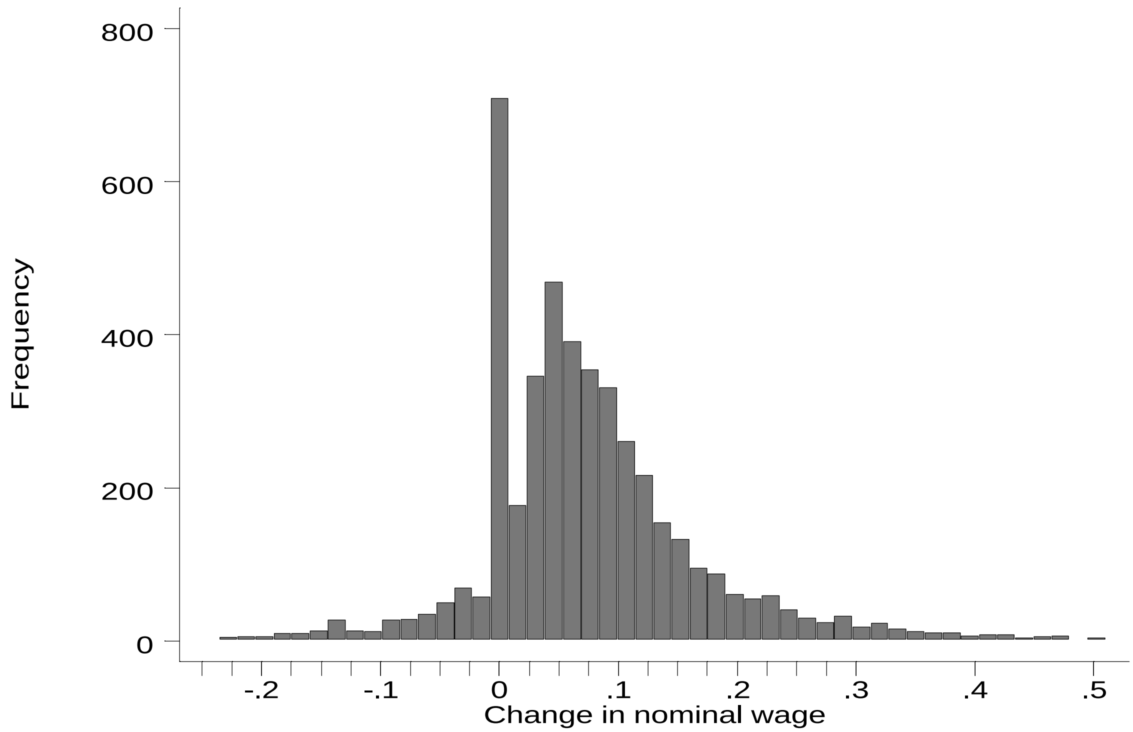

Figure 4(a): Proportion of Cuts and Freezes and Inflation

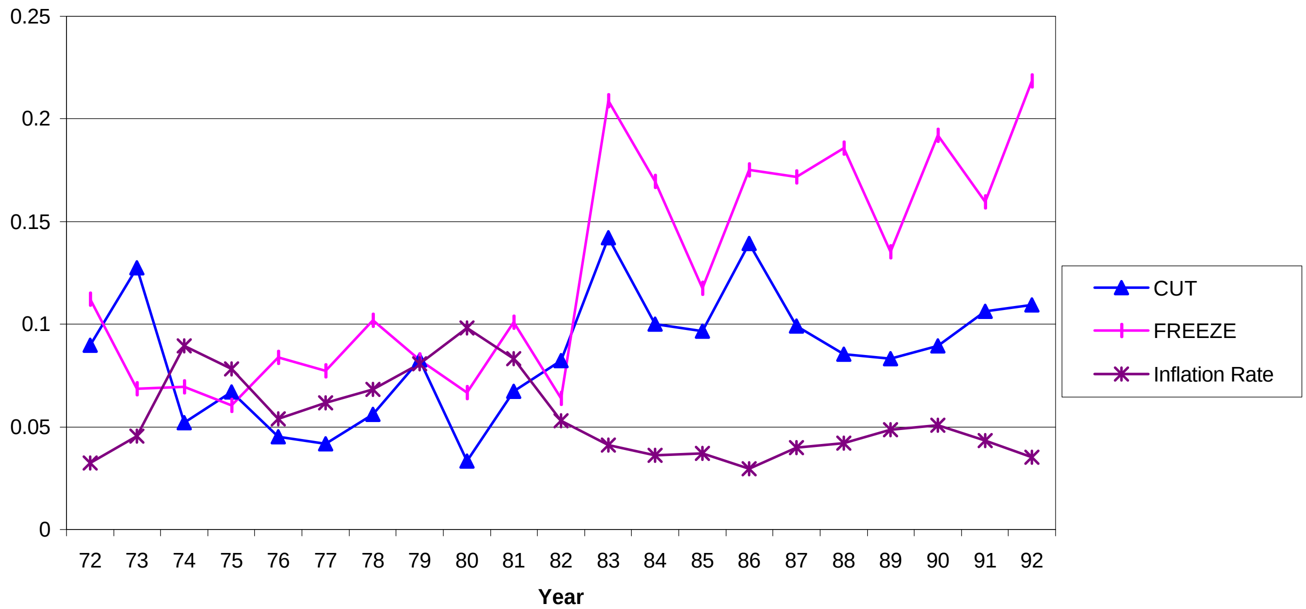

Figure 4(b): Nominal Wage Changes and Inflation