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Deflation and Downward Nominal Wage Rigidity: Evidence from Japan*

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Abstract

This study empirically analyzed downward nominal wage rigidity using time-series cross-industry data from 1981 to 2002, a period which included deflation. We found that nominal wages remained rigid to downward pressure by expected deflation and labor-market tightness. Estimations according to worker age categories revealed downward wage rigidity with deflationary pressure for most age categories. Wage rigidity during labor-market tightness was greater for younger workers.

JEL Classification Numbers: E24, E31, J30.
Keywords: wage rigidity, nominal wage, deflation, unemployment, Japan

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1. Introduction

The unemployment rate in Japan has remained high since the late 1990s. One possible cause may be downward nominal wage rigidity. Downward rigidity, or "stickiness", can be characterized by the following: (1) nominal wages respond asymmetrically to upward and downward pressures, and (2) downward wage adjustments are slower than upward wage adjustments. Downward rigidity can remarkably obstruct regulating functions in a labor market through nominal wage adjustments and can keep unemployment rates high. Figure 1 illustrates the recent sustained deflation in Japan, which should have placed downward pressure on nominal wages. Such a situation provides important information on the regulating functions of nominal wages.

Most previous studies on downward nominal wage rigidity have used microdata (*e.g.*, McLaughlin, 1994, 1999; Lebow *et al*., 1995; Card and Hyslop, 1996; Kahn, 1997; Altonji and Devereux, 1999; and Kuroda and Yamamoto, 2003a, 2003b). Microdata can help explain downward nominal wage rigidity as related to individual workers. However, such data are insufficient when discussing downward wage rigidity for the entire labor market. For instance, even if an individual's nominal wage is downwardly rigid, he or she can be replaced with another worker of equal ability and lower wage, so that wages are actually downwardly flexible for the entire labor market. This replacement effect would not necessarily be obvious through observations of individual workers. Therefore, we investigated aggregate data to determine the downward flexibility of wages within the overall labor market.

Two main factors cause nominal wage fluctuations in the labor market: (1) the tightness of the labor market and (2) variation of the expected inflation rate, as implied by the wage-version Phillips curve. Using this curve as a framework, we analyzed whether nominal wages have responded sufficiently to pressures presented by the above factors.

We also verified differences in downward rigidity across various age categories. In recent years, Japan has witnessed a disparity in unemployment rates by age. Younger workers in particular have faced higher unemployment rates. This condition can create serious social problems (see figure 2) and may relate to downward nominal wage rigidity. Therefore, we examined the existence of more severe downward rigidity in the young-worker labor market.

This paper is organized as follows: in section 2, we introduce the wage-version Phillips curve and its relevance to this study; in sections 3, 4, and 5, we present estimation equations, the data set, and estimation results, respectively. Conclusions are presented in section 6.

2. Wage-Version Philips Curve

2.1. Phillips Curve and Downward Nominal Wage Rigidity

First, we will briefly introduce the Philips curve, which is a key to this study. The wage-version Philips curve shows the hyperbolic, inverse relationship between the unemployment rate and the rate of nominal wage change. The curve implies that the rate of nominal wage change increases as the labor market tightens and falls as the labor market loosens. Therefore, this curve characterizes adjustment mechanisms of the economy-wide labor market by indicating the trade-off between employment and stable wage changes.

As Friedman (1968) noted, however, the labor market should be adjusted not by nominal wages, but by the real wages, unless workers have monetary illusions. Another viewpoint sees the actual Philips curve as shifting with the expected inflation rate, although basic Philips curve methods ignore expectation effects. In other words, the unemployment rate should be independent of the long-term rate of nominal wage change, which the expected inflation rate equates to its actual value. The unemployment rate will then reach a unique value determined by the real factors of labor market. This rate is referred to as the natural rate of unemployment. By contrast, in the short term, when the expected inflation rate differs from the actual rate, the relationship between the unemployment rate and the rate of nominal wage change will create a downward-sloping curve for each expected inflation rate. This is called the short-run Philips curve.

Second, the Phillips curve form should be distinguished from a straight line, with the focus placed on the curve's non-zero curvature, as well as its negative slope[1]. The Phillips curve is defined by its appearance, *e.g.*, a steeper slope on the left side of the natural unemployment rate and a gentler slope on the opposite side. To understand the ramifications of this shape, suppose that the unemployment rate is greater than its natural value. If the slope on the right side were nearly equal to that on the left, then the rate of change in the nominal-wage would be much smaller than what the actual curve indicates. The labor market would then soon adjust without any wage rigidity, and the unemployment rate would decrease and approach its natural state. By contrast, downward wage rigidity would cause the wage to behave differently once the unemployment rate increased beyond its natural value. For every marginal increase in the unemployment rate, the nominal wage would decrease at a much slower rate, below a certain value. The slope of the Phillips curve would therefore become much gentler, reflecting this relation. In other words, the wage would reflect the nonlinearity of the changing slopes of the left- and right-hand areas bounded by the natural unemployment rate. In summary, the Phillips curve would have non-zero curvature only if the nominal wage bore the downward rigidity.

2.2. Downward Nominal Wage Rigidity and Recent Deflation in Japan

To our knowledge, only Kimura and Ueda (2001) have analyzed the recent nominal

[1] Kurosaka and Hamada (1984) and Genda and Kondo (2003) provided similar explanations of this issue.

wage rigidity in Japan using aggregate data from the deflationary periods[2]. They examined the nonlinear relation between the actual rate of change in the nominal wage and the "notional" rate, derived from the wage-version Phillips curve equation. The nonlinearity was thought to result from the disparity between the notional and actual nominal wage change rates if actual rates failed to decrease sufficiently to equal the negative, notional values. Conversely, the actual rates were assumed to equal the notional rates when both rates were positive and wages were going to increase. Consequently, Kimura and Ueda (2001) determined the nominal wage downward rigidity defined by the above disparity, using time-series cross-industry data from 1976 to 1998. By contrast, applying the same method to aggregated time-series data from 1976 through 2000 failed to confirm downward rigidity throughout the period. Integrating those results, they concluded that wages should converge at their equilibrium levels after a time lag, although downward rigidity had been temporarily observed because of labor's monetary illusions under unfamiliar zero or negative inflation rates.

However, two problems from this study must still be solved. First, two different sources of downward pressure on wages were combined: the expected inflation rate and the unemployment rate. The authors' conclusions, however, emphasized only the role of the former factor in the disappearance of wage rigidity in recent years. Although downward pressure exerted by expected deflation began in the late 1990s, they postulated its prevalence before that period and paid little attention to how unemployment rates affected nominal wages. That is, the downward pressure observed from the time-series cross-industry data until the late 1990s should have come solely from an increase in the unemployment rate. Therefore, it is

[2] By contrast, many other studies found a nonlinear relation between the labor supply-and-demand situation and the rate of nominal-wage change. These studies, such as one by Nishizaki and Watanabe (2000), used data from the pre-deflationary period in Japan. However, Nishizaki and Watanabe (2000) has two shortcomings. First, their initial estimates were apparently biased by sample selection, depending on whether the value of the consumer price index inflation rate, adopted as the dependent variable, was three percent or more. Second, subsequent analysis emphasized a slope decrease in linear Phillips curves for a few consecutive periods as evidence of the nonlinear relation, although those separately estimated curves could have lacked mutual coincidence and thus give only indirect evidence of nonlinearity over the period.

problematic to have concluded that a greater familiarity with zero or negative inflation rates from 1999 on suddenly enabled the nominal wage to begin adapting flexibly to the expected deflation. In principle, it seems misleading to have assumed that the expected inflation rate and the unemployment rate exerted equal influence on the actual nominal wages with regard to downward rigidity. Our analysis, by contrast, contained no such assumptions.

Second, as Ohtake (2001) pointed out, Kimura and Ueda (2001) failed to consider the effect of substituting part-time workers earning lower wages for better-paid regular employees, although that replacement could have hidden persistent rigidity in the nominal wages earned by regular workers after 1999. Since Kimura and Ueda's macro-level aggregated data averaged the earnings of regular employees and part-time workers, the nominal wage could have appeared flexible in the economy-wide labor market. The examination of downward rigidity requires data that identify the wages of only regular employees since 1999.

To solve these problems, we used time-series cross-industry data to examine the behavior of nominal wages in Japan; similar data were also used by Kimura and Ueda (2001), but only until 1998, during the period 1981-2002. Our analyses were based primarily on direct estimations of the Phillips curve. We were able to identify sources of downward nominal wage rigidity using the estimated relation between each explanatory variable and the rate of change in the nominal wage. Using this method, nonlinearity of the nominal wage with regard to the unemployment rate, one of the explanatory variables, would prove the non-zero curvature of the Phillips curve. Similarly, the nonlinearity of nominal wage in regard to the expected inflation rate would show downward rigidity in the nominal wage with the expectation that such a condition would cause deflationary pressure. Different reactions to inflationary and deflationary pressure, implied by the above expectation concerning nonlinearity, would suggest asymmetric shifts of the Phillips curve, depending on whether the directions were upward or downward.

Note that our analysis contains no strong assumption regarding the perfect adjustment of nominal wages faced with upward pressure, as was postulated by Kimura and Ueda (2001). The estimation, detailed in the next section, had no need for such an assumption because it tested nonlinear influences directly (*i.e.*, the expected inflation rate and unemployment rate) on the nominal wage. Note also that the remaining part of this paper adopts the following criterion for determining the existence of nominal wage rigidity: the response of the nominal wage to downward pressure was significantly smaller than its reaction to the upward pressure. The next section will describe the specific estimation methods, which were based on the arguments presented in the first two sections.

3. Specifications

First, we postulated a basic Phillips-curve-type equation (1) with the rate of change in the nominal wage as the dependent variable. The explanatory variables π_t^e, U_t, and $y_{i,t}$ denote the expected inflation rate, unemployment rate, and industry-specific shock measured by the rate of change in the industrial real GDP, respectively:

$$w_{i,t} = \alpha_1 \pi^e{}_t + \beta_1 U_t + \gamma\, y_{i,t} + c_i + u_{i,t}\,. \qquad (1)$$

Kimura and Ueda (2001) also used these specifications to explain the rate of nominal wage change in the absence of nominal-wage downward rigidity. In the above equation, the expected inflation coefficient α_1 should be positive, and the unemployment rate coefficient β_1 should be negative. Furthermore, α_1 is thought to be unity if the nominal wage has no downward rigidity with deflationary pressure.

The sign of the industry-specific shock coefficient γ can be either positive or negative. That effect is predictable as long as the nominal wage behavior coincides with the real wage through perfect correspondence with the expected inflation rate. For example, the coefficient representing the real GDP of each industry, γ, has to be positive when real wages are procyclical. For counter-cyclical behavior in the real wage, γ should be negative. Literature on the relationship between the real wage and the business cycle provides insight into the influence of each industry's business cycle. An array of related empirical research has found both signs of the γ coefficient, and both possible effects have been theoretically supported by many other studies[3] .Recent works by Sumner and Silver (1989) and Abraham and Haltiwanger (1995) pointed out the correspondence between the inflation rate and real-wage responses to the business cycle. That is, when the former moved procyclically, the latter changed counter-cyclically and vice versa. As indicated in Figure 3, the inflation rate moved procyclically during the observation period in this study; Therefore, γ was assumed to be negative as long as α_1 equaled unity. However, γ could be positive or negative if α_1 differed from unity due to nominal wage rigidity.

$$w_{i,t} = \alpha_1 \pi^e{}_t + \alpha_2 d_t \pi^e{}_t + \beta_1 U_t + \gamma\, y_{i,t} + c_i + u_{i,t} \qquad (2)$$

Equation (2) also adds an interaction term between the expected inflation π_t^e and the dummy variable d_t, which indicates whether the expected inflation rate is below zero. Adding this variable enabled us to examine the symmetry of the nominal wage response to any downward or upward pressure by the inflation rate. In other words, downward wage rigidity would be confirmed if the coefficient α_2 in the interaction between the deflationary dummy

[3] Abraham and Haltiwanger (1995) reviewed these studies particularly.

d_t and the expected inflation π_t^e was estimated to be significantly negative and sufficiently large in absolute value to eliminate the effect of α_1. In addition, the unemployment rate coefficient β_1 was expected to be negative; the industry-specific shock coefficient γ was deemed unpredictable using the same argument given in specification (1).

$$w_{i,t} = \alpha_1 \pi^e{}_t + \beta_2 \frac{1}{U_t} + \gamma\, y_{i,t} + c_i + u_{i,t} \tag{3}$$

Equation (3) replaced the unemployment rate variable U_t with its inverse. If that specification had a higher goodness-of-fit index score than that from equation (1), the response of the nominal wage to the unemployment rate should be regarded as non-linear rather than linear. The nonlinearity implies the different slopes of the Phillips curve in the left and right areas divided by the natural unemployment rate. Further, nonlinearity provides evidence of wage rigidity against the downward pressure generated by looseness in the economy-wide labor market, as discussed previously. In addition, stronger nonlinearity should lead to a greater value for the inverted unemployment rate coefficient β_2. Using the same argument used for equation (1), α_1 was expected to be positive, and γ estimations were unpredictable *a priori*.

$$w_{i,t} = \alpha_1 \pi^e{}_t + \alpha_2 d_t \pi^e{}_t + \beta_2 \frac{1}{U_t} + \gamma\, y_{i,t} + c_i + u_{i,t} \tag{4}$$

Equation (4) addresses not only the symmetry of effects from both positive and negative expected inflation rates, but also the non-linear influence of the unemployment rate.

4. Data and Expected Inflation Rate

Our analyses primarily used industry-specific time-series data concerning changes in the nominal wage and industry-specific shock measured by the rate of change. Kimura and Ueda (2001) used a similar method. Our estimation, based on time-series cross-industry data, also used macroeconomic data on expected inflation and unemployment rates. We used averaged wage data and wage data classified into five groups based on worker age: 15-24, 25-34, 35-44, 45-54, and 55 and over[4]. Descriptions of the data sources and variables follow.

First, the nominal wages were evaluated using hourly wages including bonuses calculated from data on "contractual cash earnings", "annual special cash earnings", "annual number of scheduled hours worked", and "annual number of overtime worked". These data categories were created from the definitions provided in the *Basic Survey on Wage Structure* or the so-called *The Japanese Wage Census*.

Second, based on the static expectation, the inflation rate of each prior year was adopted as a proxy for the expected inflation, as in Kimura and Ueda (2001). We chose this expectation based on an examination of adaptive expectation. The "short-cut method" was used to obtain the worst fitness for models with larger numbers of lagged variables[5]. The best fitness was marked only when the one-term lag was thought to have explained the expectation. The results should justify the use of static expectation and the exclusion of adaptive expectation. Kitamura *et al*. (2003) and Genda and Kondo (2003) also noted that adaptive expectation poorly

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[4] See the appendix for details on the data for each age category.

[5] The short-cut method regresses lagged explanatory variables with linearly decreasing weights for longer time lags and clarifies the optimal number of such variables with regard to the goodness of fit. Specifically, the weight for the i^{th} term is defined as $w_i = (m+1-i)/[m(m+1)/2]$, where m signifies the maximum degree of lag. Each explanatory variable has the following weighted average $\sum_{i=1}^{m} w_i X_{t-i}$ for a particular m. The simple regression of X_t using the above regressors indicates the optimal degree of m, introducing the model with the highest fitness.

explained the expected inflation rate in their research on a non-accelerating inflation rate of unemployment (NAIRU) Phillips curve in which the inflation rate was the dependent variable. The inflation rate was represented by the GDP deflator in the *Annual Report on National Accounts*, as used by Kimura and Ueda (2001)[6].

Third, each annual unemployment rate used here equaled the "ratio of unemployed in the labor force" as given in the *Monthly Report on the Labor Force Survey*. The industry-specific shock was evaluated using the rate of change reported in the "Gross domestic product classified by economic activities (2001-)" or the "Gross domestic product by kind of economic activity (1980-2000)" publications in the *Annual Report on National Accounts*. Data were collected for 1981 through 2002 for 25 industries[7]. Descriptive statistics are given in the appendix.

5. Response of the Nominal Wage Rate of Change

5.1. Estimated Results using Averaged Data for all Regular Employees

Using the average wage for all age groups, pooled ordinary least squares (OLS), random effect (RE), and fixed effect (FE) estimators were obtained from regression equations (1) through (4). The Breusch-Pagan test rejected the null hypothesis of zero variance for fixed effects for all the equations at the 5% significance level. Consequently, the pooled OLS model could be rejected for all of the models. The Hausman tests did not reject the null hypothesis (*i.e.*, no correlation between the fixed effect and explanatory variables) for any of the regression

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[6] The use of the CPI, instead of the GDP deflator, did not affect our conclusions. Since the absolute values of the CPI during the deflationary period were smaller than those of the GDP deflator, a significant change was observed only in the absolute value of the interaction term between the expected inflation and the deflationary dummy.

[7] See the appendix for industry classifications. The study period (1981 to 2002) was limited due to the

equations, even at the 10% significance level. Therefore, the RE model was adopted. Its estimation results are briefly explained below and summarized in Table 1.

Model (1) produced a positive value for the expected inflation coefficient α_1; however, its level differed substantially from unity, indicating that wages did not fully behave as real wages. The effects of unemployment (β_1) and industry-specific shock (γ) were estimated to be negative. The latter, showing the negative but insignificant influences of industrial GDP, were also confirmed by specifications (2), (3), and (4).

Second, the interaction term between the deflationary dummy and expected inflation produced a significantly negative estimated α_2 by equation (2). Moreover, the marginal effect of expected inflation (α_1) approached unity more closely than α_1 in model (1). The absolute value of estimated α_2 was large enough to eliminate the effect of expected inflation (α_1); this result showed that wages had not responded to deflation pressure. In other words, while positive inflation was expected, the wages behaved as a real variable to a certain degree in response to real economic factors. Conversely, wage behavior showed downward rigidity with expected deflation and therefore played an insufficient role as a labor-market adjustment factor.

Third, the estimation results from models (1) and (3) indicated a higher significance of β_2 in (3) than β_1 in (1) and a higher coefficient of determination from (3) than from (1). Model (3) estimated coefficient β_2 from the inverse of the unemployment rate as an explanatory variable, while model (1) evaluated β_1 based on the unemployment rate itself. Hence, the wage change should have reacted non-linearly rather than linearly to the unemployment rate. There should have been downward rigidity in the nominal wage change with a real factor, *i.e.*, the labor supply-and-demand situation. The same point can be drawn from comparing the results of (2) and (4). From regression (4), however, the estimated expected inflation coefficient (α_1) was far from unity, and the counterpart of the interaction term (α_2) had a smaller absolute value. This

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availability of real GDP data for each industry at the time of writing.

result might reflect a possible correlation between expected inflation and the unemployment rate (*U*). Introducing the inverse of the latter could weaken the nonlinear influences of those two variables.

A main finding was that the nominal wage made no response to deflationary expectations; such rigidity thereby raised the real wage. In other words, the Phillips curve shifted upward under such an expectation, although it theoretically should have moved downward. Now, it can be shown how the unemployment rate would have behaved if nominal wage adjustment had worked under deflation in the same way it changed under inflation. That is, the following simple calculation based on the result of (4) can show the hypothetical value of *U* when $\alpha_2 = 0$. We based this simulation on unemployment rates from 1999-2002 because the calculation requires a one-period lagged variable of expectation. Recall that the consecutive deflationary period began in 1998 and was indexed by the GDP deflator. The average inflation from 1998 to 2001 equaled –1.265%; the coefficient of interaction between the deflationary dummy and expected inflation (α_2) was estimated to be –1.476. Therefore, the Phillips curve should have, on average, shifted downward by 1.867% (= 1.265–1.476) during the period. Any downward shift of a negatively sloped Phillips curve would reduce the unemployment if the identical rate of nominal wage change were then realized. Second, the mean unemployment rate (*U*) for 1999-2002 was 4.95%, while the estimated coefficient of the inverse of *U* equaled 24.723. The U^* denoted the unemployment rate without wage rigidity. Then, the following equality could be derived from model (4): $(1/4.95) \times 24.723 + 1.867 = (1/U^*) \times 24.723$. The equation revealed that $U^* = 3.60\%$. Since the above procedure was based roughly on the wage-unemployment correlation underpinning the Phillips curve, room was left for improvement, such as dynamic modeling of the process.

5.2. Estimation Results for Each Age Group

This section examines the estimations for each age group. The Breusch-Pagan and Hausman tests were applied to equations (1) through (4) for every age category, as in the previous estimations. The former test rejected the null hypothesis of zero fixed-effect variance about all age groups, but model (4) revealed a 5% significance level for the 55-and-over age category. Nevertheless, the *p*-value in that exceptional case did not exceed 0.0501. By contrast, the latter test never rejected the hypothesis that the fixed effect is uncorrelated with the explanatory variables even at the 10% significance level. The RE estimation was then applied to every age class; tables 2 through 6 show the outcomes.

The model (2) results for each age category resulted in a significantly negative coefficient for the interaction between the deflation dummy and expected inflation α_2; the wage showed downward rigidity with deflation. Except for the 25-34 age category, the results also showed that the influence of expected inflation α_1 tended to approach unity much more closely than the α_1 estimated for all age groups. The coefficients for the 45-54 and 55-and-older age categories were especially close to unity: 0.877 and 0.893, respectively. Testing the null hypothesis that estimates would equal one revealed *p*-values of 0.324 and 0.506, respectively. This result implies that the nominal wage responded somewhat flexibly to expected prices when positive inflation was predicted. The data were separated by age groups to clarify the theoretical relationship between the nominal wage and expected inflation for each age group.

By model (4) the coefficient representing the interaction between the deflation dummy and expected inflation, $d_t \pi^e{}_t$, was significantly negative for all age categories except the 55-and-older category; this indicates downward wage rigidity with deflation. The deflation rate coefficient α_1 was far from unity with weaker non-linearity than that estimated using the averaged total for all age groups. Here, the insignificant estimate for the 55-and-older category could indicate flexible adjustment of the nominal wage during the deflationary period, as opposed to the contradictory result obtained by model (2) and described above. Considering these two opposite effects, no decisive argument can be derived concerning this age group.

The replacement of explanatory variable *U*, the unemployment rate, with its inverse value improved the performance of the estimation models. Comparing the R^2uared and t-value on each regressor between the results from models (3) and (1), the estimates using (3) were higher than those using (1) for most age groups. Similar results were obtained on comparing specifications (4) and (2). While equations (1) and (2) included the unemployment rate *U* itself as an explanatory variable, (3) and (4) used its inverse value. The same findings were obtained for all age groups. However, for the 55-and-older category, the performance of (3) was poorer than that of (1); similarly, model (4) performed worse than (2). Therefore, the possibility of downward rigidity with unemployment rate could almost be ignored for that age category.

Table 7 contrasts the estimated inverse unemployment rate (β_2) coefficients for all age categories. Greater values indicate stronger nonlinearity of the nominal wage with the unemployment rate. The results from both (3) and (4) show large estimated values for the 15-24 age bracket, but small values for the 35-44 and 45-54 age categories. These results suggest stronger downward rigidity for workers aged 15 to 24, which can be considered a reason for the recent rise in unemployment among young workers. The effects of downward rigidity on unemployment in the 55-and-over category can be regarded as linear, as mentioned above, although the corresponding estimates were greater than those for the 35-44 and 45-54 age categories.

Strong wage rigidity for those aged 15-24 was confirmed by the higher inverted unemployment rate coefficient; this result might be explained by a greater proportion of workers in their early 20s, as opposed to those under age 19. In other words, the increasing sample weight for the early 20s of relatively higher wages could have raised the mean wage for the 15-24 age groups and caused the mean wage to appear rigid. To examine this possibility, we analyzed younger-worker subcategories as follows. First, the 15-24 age category was divided into the three age subcategories: 15-16, 17-18, and 19-24[8]. Second, regression analyses for the three categories were carried out using the processes described above. Model (3) estimated the inverse unemployment rates as 57.759, 28.781, and 41.796, respectively, for each group; model (4) provided estimates of 57.920, 34.397, and 41.839. These estimates were greater than those for the same coefficient estimated for groups aged 25 and over (Figure 7). This result proves the stronger downward rigidity for all three under-24 subgroups. Thus, the greater numbers of workers in their early 20s in the Japanese labor market is not the reason for the nominal wage rigidity in the wages of the 15-24 age groups, contrary to the suggested alternative.

6. Conclusions

This study examined downward nominal wage rigidity using data from Japan observed during a period that included deflationary periods. Estimates were based on time-series cross-industry data from 1981 to 2002. We found that nominal wages were rigid to downward pressure by both expected deflation and the unemployment rate. Estimates for each age category confirmed the existence of downward rigidity to deflationary pressure in all the groups, although rigidity for the 55-and-older category was ambiguous. Considering those effects, it can be concluded that mild inflation is preferable from the standpoint of labor market adjustments. Moreover, the nominal wages for workers in younger age categories were found to be more downwardly rigid to the real pressure of the unemployment rate. Therefore, downward rigidity likely contributed to rising youth unemployment in recent years. Further study using dynamic analysis should examine the extent to which downward rigidity has raised the

[8] This classification follows that adopted by the *Basic Survey on Wage Structure*.

unemployment rate.

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Figure1 Trends in the Inflation Rate

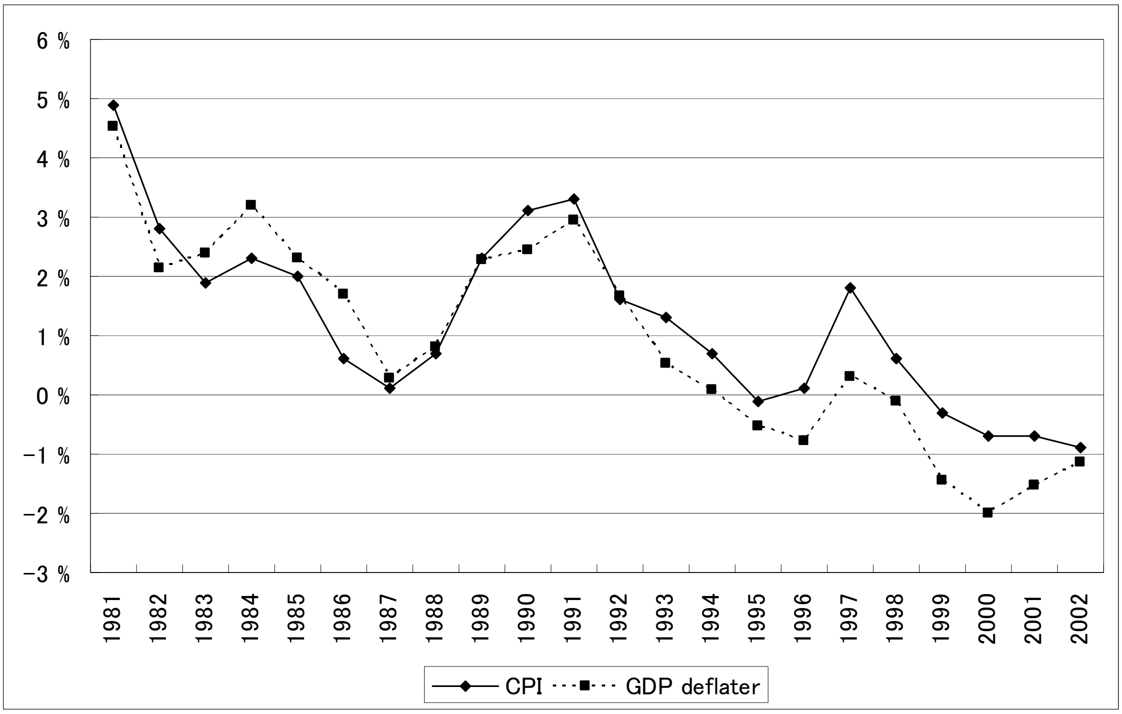

Figure 2 Trends in the Unemployment Rate by Age Categories

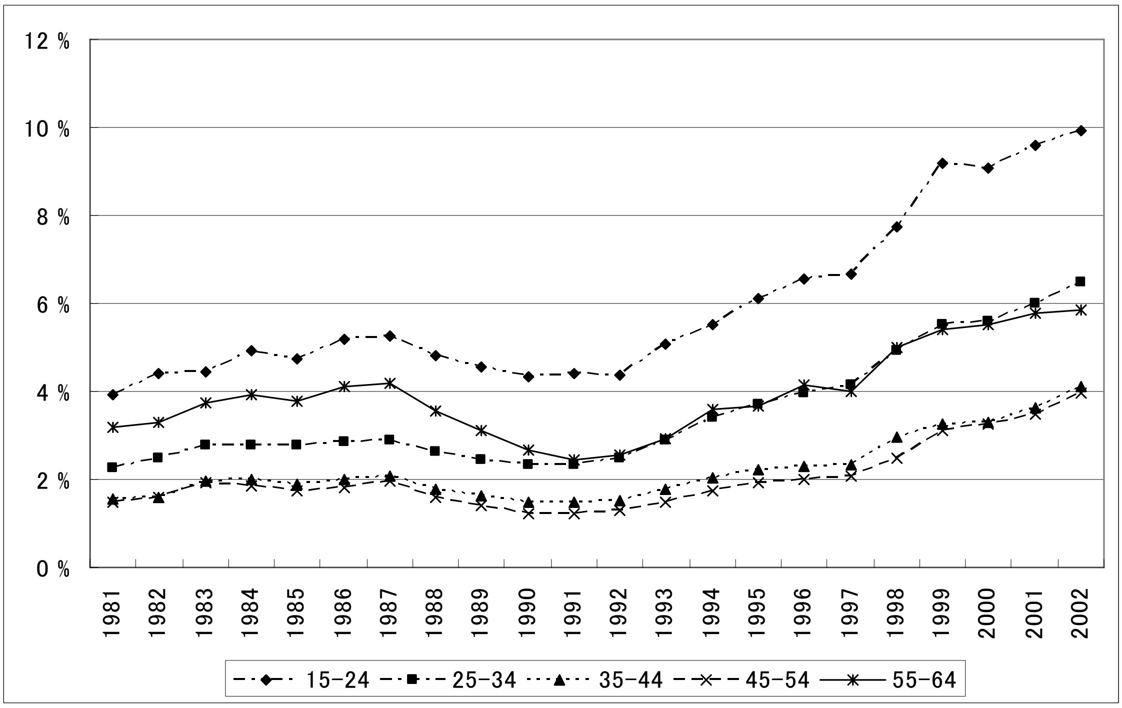

Figure 3 Procyclicality of the Inflation Rate

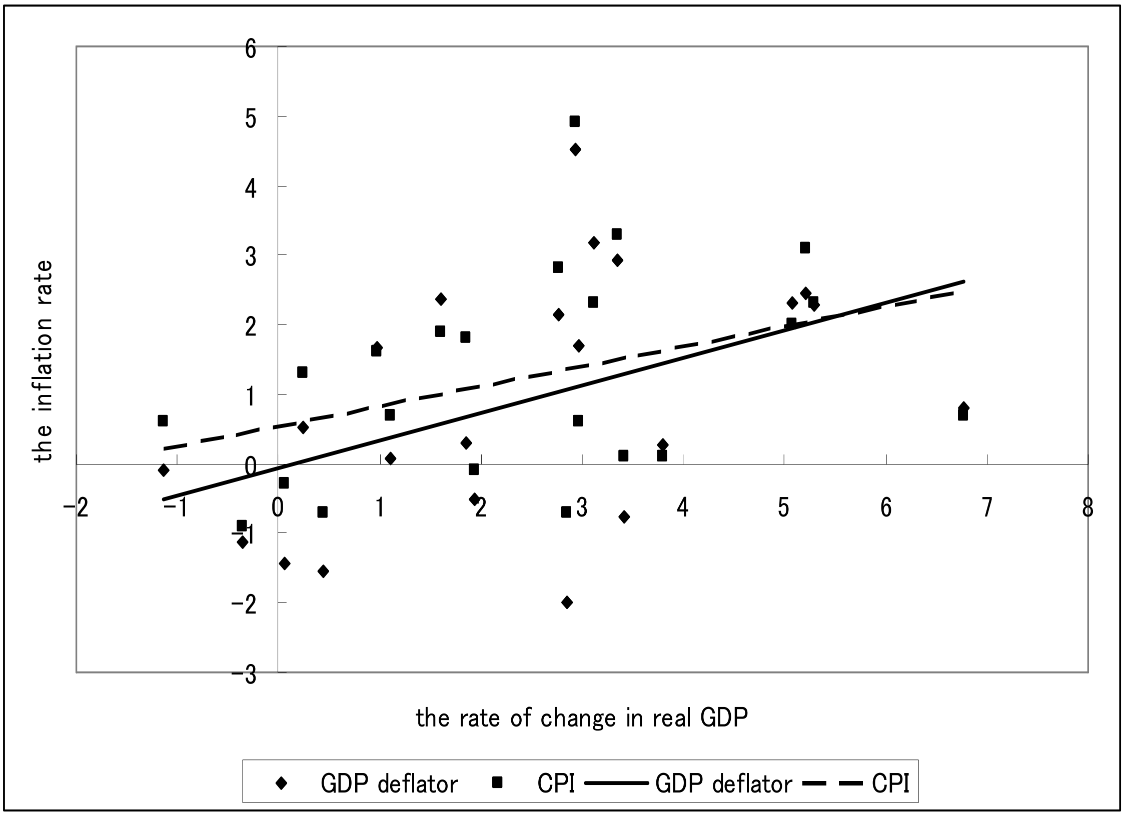

Table 1 Estimation Results using the Average Wage Total for all Age Groups

	(1)		(2)		(3)		(4)	
Expected Inflation	0.376	***	0.585	***	0.115		0.359	***
	(0.109)		(0.110)		(0.105)		(0.111)	
Deflation Dummy *Expected Inflation			-2.004	***			-1.476	***
			(0.311)				(0.272)	
Unemployment	-1.364	***	-2.039	***				
	(0.191)		(0.211)					
1/Unemployment					21.436	***	24.723	***
					(2.017)		(2.055)	
Industry-Specific Shock	-0.302		-0.971		-0.946		-1.151	
	(1.443)		(1.394)		(1.370)		(1.334)	
Constant	6.604	***	7.932	***	-4.797	***	-6.616	***
	(0.707)		(0.711)		(0.617)		(0.688)	
Number of Observations	525		525		525		525	
R^2	0.360		0.408		0.423		0.454	

Note: Standard errors are in parentheses. *** indicates significance at the 1% level, ** indicates significance at the 5% level, and * indicates significance at the 10% level.

Table 2 Estimation Results for the 15–24 Age Category

	(1)		(2)		(3)		(4)	
Expected Inflation	0.543	***	0.661	***	0.193		0.330	**
	(0.117)		(0.119)		(0.124)		(0.129)	
Deflation Dummy *Expected Inflation			-1.293	***			-0.967	***
			(0.318)				(0.282)	
Unemployment	-0.636	***	-0.879	***				
	(0.109)		(0.123)					
1/Unemployment					41.480	***	46.593	***
					(4.802)		(4.982)	
Industry-Specific Shock	1.373		0.914		0.493		0.303	
	(1.453)		(1.436)		(1.407)		(1.394)	
Constant	5.743	***	6.716	***	-5.148	***	-6.490	***
	(0.769)		(0.795)		(0.758)		(0.846)	
Number of Observations	525		525		525		525	
R^2	0.393		0.412		0.434		0.447	

Note: Standard errors are in parentheses. *** indicates significance at the 1% level, ** indicates significance at the 5% level, and * indicates significance at the 10% level.

Table 3 Estimation Results for the 25-34 Age Category

	(1)		(2)		(3)		(4)	
Expected Inflation	0.246	**	0.423	***	-0.172		0.061	
	(0.118)		(0.117)		(0.115)		(0.120)	
Deflation Dummy *Expected Inflation			-2.022	***			-1.452	***
			(0.317)				(0.270)	
Unemployment	-1.082	***	-1.643	***				
	(0.158)		(0.176)					
1/Unemployment					24.944	***	28.192	***
					(2.221)		(2.247)	
Industry-Specific Shock	-0.536		-1.401		-1.937		-2.223	
	(1.476)		(1.429)		(1.387)		(1.352)	
Constant	5.680	***	6.921	***	-5.493	***	-7.169	***
	(0.684)		(0.688)		(0.599)		(0.661)	
Number of Observations	525		525		525		525	
R^2	0.318		0.368		0.402		0.433	

Note: Standard errors are in parentheses. *** indicates significance at the 1% level, ** indicates significance at the 5% level, and * indicates significance at the 10% level.

Table 4 Estimation Results for the 35-44 Age Category

	(1)		(2)		(3)		(4)	
Expected Inflation	0.511	***	0.740	***	0.172		0.450	***
	(0.119)		(0.122)		(0.116)		(0.125)	
Deflation Dummy *Expected Inflation			-2.077	***			-1.599	***
			(0.357)				(0.308)	
Unemployment	-1.369	***	-2.301	***				
	(0.280)		(0.315)					
1/Unemployment					13.447	***	15.714	***
					(1.554)		(1.578)	
Industry-Specific Shock	-2.172		-2.871	*	-2.899	*	-3.065	**
	(1.631)		(1.586)		(1.555)		(1.518)	
Constant	4.992	***	6.245	***	-4.246	***	-6.106	***
	(0.743)		(0.752)		(0.671)		(0.746)	
Number of Observations	525		525		525		525	
R^2	0.268		0.313		0.331		0.364	

Note: Standard errors are in parentheses. *** indicates significance at the 1% level, ** indicates significance at the 5% level, and * indicates significance at the 10% level.

Table 5 Estimation Results for the 45-54 Age Category

	(1)		(2)		(3)		(4)	
Expected Inflation	0.600	***	0.877	***	0.417	***	0.713	***
	(0.120)		(0.124)		(0.110)		(0.124)	
Deflation Dummy *Expected Inflation			-2.528	***			-1.662	***
			(0.408)				(0.343)	
Unemployment	-1.760	***	-2.919	***				
	(0.273)		(0.323)					
1/Unemployment					11.278	***	13.251	***
					(1.160)		(1.206)	
Industry-Specific Shock	0.041		-0.617		-0.289		-0.403	
	(1.760)		(1.703)		(1.678)		(1.643)	
Constant	5.808	***	7.155	***	-3.738	***	-5.610	***
	(0.672)		(0.685)		(0.569)		(0.678)	
Number of Observations	525		525		525		525	
R^2	0.333		0.379		0.390		0.417	

Note: Standard errors are in parentheses. *** indicates significance at the 1% level, ** indicates significance at the 5% level, and * indicates significance at the 10% level.

Table 6 Estimation Results for the 55-and-older Age Category

	(1)		(2)		(3)		(4)	
Expected Inflation	0.669	***	0.893	***	0.764	***	0.914	***
	(0.143)		(0.161)		(0.134)		(0.163)	
Deflation Dummy *Expected Inflation			-1.392	***			-0.699	
			(0.462)				(0.436)	
Unemployment	-1.440	***	-1.761	***				
	(0.243)		(0.264)					
1/Unemployment					19.345	***	20.507	***
					(3.340)		(3.413)	
Industry-Specific Shock	-1.483		-1.547		-0.820		-0.758	
	(2.164)		(2.148)		(2.161)		(2.158)	
Constant	8.497	***	9.124	***	-2.519	***	-3.195	***
	(1.087)		(1.099)		(0.831)		(0.931)	
Number of Observations	525		525		525		525	
R^2	0.276		0.288		0.274		0.277	

Note: Standard errors are in parentheses. *** indicates significance at the 1% level, ** indicates significance at the 5% level, and * indicates significance at the 10% level.

Table 7 Comparison of the Inverse Unemployment Rate Coefficients by Age Categories as Determined by Equations (3) and (4)

	15–24	25–34	35–44	45–54	55-and-older
(3)	41.480	24.944	13.447	11.278	19.345
(4)	46.593	28.192	15.714	13.251	20.507

Data Appendix

1. Derivation of Average Wage for Each Age Bracket

NIKKEI NEEDS data were available in electronic form and provided the unemployment rate via a *Monthly Report on the Labor Force Survey* that included figures classified by five age brackets: (1) 15-24, (2) 25-34, (3) 35-44, (4) 45-54, and (5) 55 and older. For comparison, the available data on nominal wages from the *Basic Survey on Wage Structure* were averaged for 12 age groups: (1) 17 and younger, (2) 18-19, (3) 20-24, (4) 25-29, (5) 30-34, (6) 35-39, (7) 40-44, (8) 45-49, (9) 50-54, (10) 55-59, (11) 60-64, and (12) 65 and older. To make these two sources of data correspond, the mean wages for the former five groups were calculated, as shown in the example below.

(Example: average wage of the 25-34 age category)

$$\overline{w}_{25-34} = \frac{\overline{w}_{25-29} \times l_{25-29} + \overline{w}_{30-34} \times l_{30-34}}{l_{25-29} + l_{30-34}}$$

where $\overline{w}_{i-j}$ denotes the average wage over the *i-j* age bracket, and l_{i-j} equals the number of workers aged between *i* and *j* ($i < j$).

2. Industry Classifications

Following the criteria adopted in Kimura and Ueda (2001), Japanese domestic industries were categorized according to “major groups” for non-manufacturing and “medium groups” for manufacturing as given in the *Standard Industrial Classification for Japan*. Of the categories, those in both the *Basic Survey on Wage Structure* and *Annual Report on National Accounts* were chosen for our analyses. Those 25 industries were as follows: (1) mining, (2) manufacture of food products and beverages (including tobacco and feed), (3) manufacture of textile mill products, except apparel and other finished products made from fabrics and similar materials, (4) manufacture of pulp, paper, and paper products, (5) manufacture of chemicals and allied products, (6) manufacture of ceramic, stone, and clay products, (7) manufacture of iron and steel, (8) manufacture of non-ferrous metals and products, (9) manufacture of fabricated metal products, (10) manufacture of general machinery, (11) manufacture of electrical machinery, equipment, and supplies, (12) manufacture of transportation equipment, (13) manufacture of precision instruments and machinery, (14) manufacture of apparel and other finished products made from fabrics and similar materials, (15) manufacture of lumber and wood products except furniture, (16) manufacture of furniture and fixtures, (17) publishing, printing, and allied industries, (18) manufacture of rubber products, (19) construction, (20) electricity, gas, heat supply, and water, (21) wholesale and retail trade, eating and drinking establishments, (22) finance and insurance, (23) real estate, (24) transport and communications, and (25) services.

3. Descriptive Statistics

	Obs.	mean	S.D.	min	max
Rate of change in the nominal wage (overall)	525	2.664	3.167	-10.566	15.755
Rate of change in the nominal wage (15-24)	525	2.473	3.266	-6.545	13.657
Rate of change in the nominal wage (25-34)	525	2.033	3.114	-6.035	13.863
Rate of change in the nominal wage (35-44)	525	2.385	3.347	-8.770	16.552
Rate of change in the nominal wage (45-54)	525	2.801	3.809	-15.593	22.220
Rate of change in the nominal wage (55 and older)	525	3.433	4.520	-13.887	22.791
Expected inflation rate	21	1.009	1.741	-1.990	4.525
Unemployment rate (overall)	21	3.162	1.017	2.100	5.400
Unemployment rate (15-24)	21	6.047	1.907	4.350	9.925
Unemployment rate (25-34)	21	3.592	1.330	2.333	6.467
Unemployment rate (35-44)	21	2.252	0.751	1.475	4.108
Unemployment rate (45-54)	21	2.053	0.773	1.217	3.967
Unemployment rate (55 and older)	21	3.965	1.027	2.442	5.842
Rate of change in the industrial real GDP	525	0.018	0.080	-0.196	0.535