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Wage flexibility in Britain: some micro and macro evidence

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Abstract

This paper uses the British New Earnings Survey (NES) to derive both macro and micro measures of wage rigidities. The data set spans the 1975-2000 period, with wage observations covering approximately 1% of the British workforce. Using this data set, we consider whether wages have become more flexible in recent years. Evidence drawn from macroeconomic wage equation estimates suggests that, while the relationship between wages and unemployment seems to have changed, the responsiveness of wages to unemployment rates has declined in the 1990s. In contrast to previous findings, these results suggest that wages have become less flexible. Micro tests mainly reveal that the prevalence of nominal rigidities has not declined, although there is weak evidence for a decline in the most recent years. Micro tests of wage rigidities also include an analysis of real wage rigidities. There is evidence in the NES of spikes in the wage-change distribution at the rate of inflation, beyond what would be anticipated from the rest of the distribution, suggesting the presence of real wage rigidities. This evidence is shown to be statistically significant, and points to an increase in real rigidities over the period of study.

Key words: Wage rigidities, wage flexibility, wage changes.
JEL classification: E24, J31.

Summary

Increased wage flexibility was often cited as the main reason behind weaker inflationary pressures in the 1990s. Wage flexibility can be defined as either a micro or macroeconomic concept; each is quite distinct, although potentially related. For example, in a Phillips curve an increase in macroeconomic wage flexibility is often captured by a larger wage response for a given unemployment rate. On the other hand, increased microeconomic wage flexibility is usually identified by the lack of evidence for wage rigidities, such as limited evidence of a spike in the distribution of wage changes at zero. An abundance of zero wage changes in the data might indicate an inability to adjust wages in a timely manner.

This paper uses data from the British New Earnings Survey from 1975-2000 to derive both macro and micro measures of wage rigidities. Because the data span a twenty-six year period, the behaviour of micro and macro flexibility measures can be compared over time. In addition, we can investigate whether there is any evidence that the behaviour of either measure of flexibility has shifted over time.

To keep the analysis simple, we consider whether there has been a one-off shift in wage flexibility. Regional wage rates do not appear to have been more responsive to regional unemployment levels in the 1990s than in earlier years. Instead, estimates focusing on the 1990s reveal no statistically significant aggregate wage response to regional unemployment levels. The overarching conclusion is that macroeconomic tests leave much to explain, but the estimates have revealed some patterns that are worth trying to reconcile with other sources of evidence.

In individual-level wage data, there is evidence of so-called nominal wage rigidity. In contrast to previous findings, the evidence is generally stronger for the latter half of the sample period. The degree of nominal wage rigidity is somewhat smaller than found in related research using data for the United States over the full sample period, but the estimates over the recent past are similar. The evidence of nominal rigidities is somewhat surprising in the light of the popular view that the UK labour market has become gradually more flexible.

1. Introduction

Increased wage flexibility is often cited as the main reason behind weaker inflation pressures in the 1990s. Wage flexibility can be defined as either a micro or macroeconomic concept, each of which is quite distinct although potentially related. For example, an increase in macroeconomic wage flexibility is often captured by a larger wage response for a given unemployment rate in a Phillips curve. On the other hand, increased microeconomic wage flexibility is usually identified by the lack of evidence for wage rigidities, such as limited evidence of a spike in wage-change distributions at zero. The abundance of zero wage changes in the data presumably reflects some hindrance in the accuracy and speed of wage adjustment.

This paper uses the British New Earnings Survey in order to derive both macro and micro measures of wage rigidities. Because the data span the 1975-2000 period, the behaviour of micro and macro flexibility measures can be compared over time. In addition, we can investigate whether there is any evidence that the behaviour of either measure of flexibility has shifted over time.

To keep the analysis simple we consider whether there has been a one-off shift in the parameters of interest. A permanent change in the wage-setting pattern can be motivated by a variety of factors, although two stand out. First, institutional arrangements for employment in the United Kingdom have clearly changed over the period under investigation.[1] In particular, Cully *et al* (1999) document the decline of private sector unions and substantial shifts in work patterns, such as the

[1] See for example Clegg (1979) and Kessler and Bayliss (1995).

increasing popularity of flexible hours and the rise of the service sector, using the Workplace Employee Relations Survey.[2] Second, monetary policy arrangements have changed markedly since 1992. The form of wage contracts may have changed in response to the low and stable inflation rates of the late 1990s under inflation targeting.

This paper offers a new approach in its exploration of real wage rigidities from a micro perspective. The nominal rigidities literature has typically focused on the observation that spikes are evident in the distribution at points well above zero. A modified Kahn (1997) approach reveals that these spikes and distributional anomalies around the prevailing inflation rate are statistically significant.

Since this paper compares alternative definitions of wage flexibility, the existing literature is reviewed throughout the paper. Many of our results are new for the United Kingdom, while others update existing work. A substantial part of the paper involves showing what these alternative definitions and measures reveal for Great Britain when derived using a consistent data set. As we shall see, such a comparison exposes the importance of some methodological differences. In addition, the 'facts' of the UK labour market are open to interpretation and existing results are challenged.

In the following sections, we will begin by describing the data before moving to the macro and micro evidence on pay rigidities.

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[2] Cully, Woodland, O'Reilly and Dix (1999).

2. The New Earnings Survey panel data set

The New Earnings Survey panel data set [3] is an exceptionally good source of information on wages. The survey covers a random 1% sample of the British workers based on two digits of their National Insurance (NI) number. Because NI numbers do not change and are always used when income is reported to the tax authorities, individuals can be followed if they change employers and residences within Britain. This feature makes the data set preferable to household-based panel data sets, where following individuals who change addresses can be difficult.

The survey is sent to the employers of these individuals, who are instructed to report the identified individuals' current earnings each April. In addition, employers report on several key pay determinants for the relevant pay period: the hours of work on which the pay is based, the number of paid overtime hours, whether pay was reduced due to absence and whether any bonuses were paid. In addition, the firm reports whether the individual's 'job' has changed since the previous year. The individual's job may change in the firm without a new occupation code or a change of firms, so wage changes can be analysed for persons continuing in the same position within a firm. These features enable calculation of an unusually pure measure of base pay, free from several of the common sources of error.

Employers also report an individual's sex, age, occupation, industry, place of work, and whether a collective agreement or a wage council covers their pay. The occupation and industry codes change

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[3] The New Earnings Survey is crown copyright and is used with the permission of the Office for National Statistics. More details on the design and coverage rates in each year can be found in annual result books ONS (1975-2000).

over the sample period. Helpfully, in the year of the largest change to the occupation coding system (1995), individuals' occupations were coded both in the old and new scheme. To construct consistent occupation codes, these codes were matched at the three-digit level and aggregated on both sides until the resulting codes were likely to correspond with both the old and the new codes. In most cases, the vast majority of individuals within a code do not change their constructed 'consistent' code when the new coding scheme is put into effect.

Despite the straightforward design of the survey, there is a tendency to undercount low-wage workers and workers in smaller firms. This tendency has been documented in Stuttard and Jenkins (2001). Under-sampling occurs in part because individuals whose wages are below a threshold are excluded from the tax system and thus no address is provided for their employer. Also, despite the ability to follow workers using tax records, a sizable fraction of surveys are returned uncompleted because the employee no longer works for that organisation. Overall, about 24% of the potential records are not available, with the largest categories being a change of employer (8.7%), survey not returned (5.8%), and out-of-scope (3.7%).[4]

These data losses are not likely to alter the qualitative results of this analysis. There may be a tendency for low-wage workers to be more exposed to wage rigidities if their industries tend to have smaller wage increases. Alternatively, low-wage workers and those who change employers may be

[4] These figures were drawn from 1996 New Earning Survey notes, but were not unusual. Major 'out of sample' groups are pensioners, non-salaried directors, working for spouse, working outside the UK.

more likely to have substantial wage changes because they are earlier in their careers or are changing contracts. To the extent that wage-setting behaviours are present for groups that are surveyed, the primary loss will be in getting the level of wage rigidities wrong. Trends should be robust because the sample losses do not vary much over time.

3. Macro evidence: wage flexibility

The simplest characterisation of macroeconomic wage-setting behaviour is a wage curve or wage equation. Theoretical models provide insight into the key macroeconomic influences on wage setting, but they are not informative about the particular relationship to be estimated. Most of the literature has therefore been primarily focused on studying empirical relationships. In particular, the literature on whether the best estimate of the wage response to unemployment is a wage curve or a Phillips curve is extensive, albeit inconclusive, in that both functional forms continue to be estimated.[(5)] At a micro level, most tests are directly focused on wage changes, as wage levels are determined by a wide variety of idiosyncratic factors that are largely unchanged from year to year and which are accounted for in the previous year's wage.

Chart 1 shows the relationship between claimant count unemployment rates and wage changes in

[(5)] The macroeconomic estimates most closely follow Bell, Nickell, and Quintini (2000) and Staiger, Stock and Watson (2001). Bell, Nickell and Quintini (2000) estimate both forms but favour a flexible form of the wage curve (with lagged wages entering with a coefficient less than one). Staiger *et al* (2001) favour estimates in changes, although they consider some alternatives that are consistent with wage curves.

Great Britain.[6] Following the approach of Bell, Nickell and Quintini (2000), the wage rate shown controls for changes in the composition of the workforce over time.[7] The graph shows periods, for example 1980 to 1985, when the traditional Phillips curve relationship appears to hold. The figure does not attempt to take account of shifts in expected inflation, of course, which were pronounced over much of the period in question. Nevertheless, the period since 1993, during which inflation was steady or dropping in tandem with falling unemployment and when expectations were plausibly firmly anchored under the new inflation-targeting regime, is more difficult to account for without assuming substantial structural changes in the economy. These changes need not be located in the labour market. For example, firms could have accepted lower product margins in this period, resulting in a higher labour share of income. Alternatively, changes could have occurred within the labour market. An increasing degree of wage flexibility may be a potential explanation for the Phillips curve relationship breaking down.

3.1 'Wage equations, wage curves, and all that'

Wage inflation and the unemployment rate show considerable variation across regions in the United Kingdom, allowing for regional variation to refine the estimation of models of wage inflation. In 'Wage equations, wage curves, and all that', Bell, Nickell and Quintini (2000) (BNQ) measure flexibility in the context of regressions of regional wages (levels or changes) on regional

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(6) Claimant count unemployment rates are seasonally adjusted figures for the month of April provided by the Office for National Statistics.

(7) A first-stage regression of wages on age, age squared, 8 occupation dummy variables, 10 industries, 10 regions, and years dummy variables is run. The coefficients on the time dummies represent the average wage rates after controlling for human capital. This estimate is differenced to yield the reported wage inflation measure.

unemployment rates and panel regression controls. Their paper is focused on comparing alternative specifications of the relationship between wages and unemployment. The BNQ paper is highly relevant for establishing a baseline for the United Kingdom, since their estimates show the value of applying regional estimates over the recent period (eg through 1996) and using the same data source (NES Panel).

The approach to adjusting regional wages throughout this paper follows that of BNQ. A model of wages is estimated using individual-level data on the workers' characteristics, with a set of region cross year effects, which become our adjusted mean wage figures for the region.

$$w_{ijt} = \alpha_{jt} + \boldsymbol{\gamma} \int \int \mathbf{X}_{\mathbf{it}} + \varepsilon_{it} \qquad \textbf{(1)}$$

where w_{ijt} is individual i's in region j wage level at time t, α_{jt} accounts for regional pay differences in each year, $\mathbf{X}_{\mathrm{it}}$ are vectors of individuals' observed characteristics and γ is a vector of coefficients constant across regions.(8) The regional cross time intercepts (α_{jt}) can be interpreted as the average wage rate, controlling for human capital.

Second-stage regressions use the α_{jt} estimated in the first stage as the dependent variable. The control variables are the lagged dependent variable, regional unemployment rates and general panel controls for time and region. In its most detailed form the wage curve regression is as follows:

$$\alpha_{jt} = \beta_1 \alpha_{jt-1} + \beta_2 \ln u_{jt} + \boldsymbol{\delta}_j \mathbf{D}_j + \boldsymbol{\delta}_t \mathbf{D}_t + \boldsymbol{\delta}_{2j} (\mathbf{D}_j \cdot t) + \upsilon_{jt} \qquad \textbf{(2)}$$

(8) This regression is conceptually quite close to that used in Bell, Nickell and Quintini (2000). The first-stage regression becomes wages on age, age squared, 8 occupation dummy variables, 10 industries, and 10 regions interacted with year dummy variables. The differences are mainly due to coding differences in the occupations.

where u_{jt} is the region j claimant count unemployment rate. $\mathbf{D}_j$ and $\mathbf{D}_t$ are vectors of dummy variables to absorb all constant regional and time effects. The last term is a vector of region-specific time trends, to account for steady changes between regions. With these controls, no additional explanatory power is gained by including variables that vary only between regions or across time, as this variation can be fully absorbed by the fixed effects. For example, there would be no difference in the R^2 or coefficients of interest if national inflation rates or productivity growth were added as controls in this regression. Estimating the models on wage change (rather than levels) simply imposes $\beta_1 = -1$ and moves the $\alpha_{jt\text{-}1}$ term over to the left-hand side of the equation.

$$\omega_{jt} \equiv \alpha_{jt} - \alpha_{jt-1} = \beta_2 \ln u_{jt} + \boldsymbol{\delta}_j \mathbf{D}_j + \boldsymbol{\delta}_t \mathbf{D}_t + \boldsymbol{\delta}_{2j} (\mathbf{D}_j \cdot t) + \upsilon_{it} \qquad \textbf{(3)}$$

Our estimates extend the data set through 2000 and include both male and female earners. Not surprisingly, the results are close to BNQ's findings for men. Table A reports the baseline regression for some of the specifications tested in BNQ.(9) In all cases, the standard errors have been adjusted for contemporaneously correlated errors across panels. The log unemployment rate has a negative coefficient in all of these regressions, but it is not always significant. The coefficient on the log of the unemployment rate is somewhat lower in these estimates than in BNQ, but the overall conclusions are similar. Their preferred specification is with a lagged dependent variable and regional trends and this specification shows the clearest relationship between unemployment and wages in our estimates as well.

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(9) Results were also calculated for a male-only sample, like BNQ, and were similar enough to make them repetitive.

The wage curve (levels) specification rejects (at the 5% significance level) a coefficient of 1 on the lagged dependent variable, favouring the wage curve specification. We are not focusing on this issue, we favour the wage curve specification only because the relationship between regional wages and unemployment rates seems to be statistically more precise. Nonetheless, wage equations (the specification with wage changes as the dependent variable) are also estimated to make the difference between BNQ and Stock, Staiger and Watson's (2001) approach, more clearly delineated.

Both the regional and time dummy variables are doing a substantial amount of work in these regressions. Both are generally significant at the 1% level, indicating the importance of general controls in this relationship. The controls for regional trends sometimes reduce the significance of the fixed regional dummies, but it is logical to include the intercept shifters as a baseline when trend variation is seen. In any case, regional trends help make the unemployment/wage relationship statistically significant. This could be due to substantial regional productivity trends that are evident in the UK data, which will be more directly explored in the next section.

In the wage curve (or wage equation) framework, wages are said to have become more flexible if their response to unemployment increases. BNQ test whether the responsiveness of wages to unemployment is larger in the second half of their sample (1988 to 1996) and find evidence that wages are more flexible in the later portion of the sample.

Given the minor differences between our results and BNQ's results reported in Table A, the results of these wage flexibility tests might appear to be a foregone conclusion. This is not the case, as the flexibility estimators put a lot of emphasis on the time variation arising from the additional four years of data in our sample: a period of low inflation rates and unemployment declines. Table B shows the results of break-point regressions on wage levels and changes including regional trends.

The indicators of a change in flexibility are generally statistically significant, but the sign of the interaction indicates that the labour market has become less flexible. Wage responses to unemployment are unambiguously lower in any of the suggested Phillips curves specifications shown earlier when the later 1992-2000 period is included.

While Chart 1 suggests that 1992 (the beginning of the expansion) is a logical point to break the sample, the choice remains arbitrary and may be the source of the difference between our results and those reported in BNQ. Chart 2 shows the full set of results of moving the potential break point in the data from 1981 to 1996. This chart shows the parameters estimated in a sequence of regressions following Table B (column 1) with alternative dates for the change in wage-setting patterns. The left-hand panel shows the expected response to unemployment rates in the earlier portion of the data, defined by the year of the break shown on the x-axis. The right-hand panel shows the response after the date of the break. The standard error bands shown are for 95% confidence ranges. Choosing an earlier break point in the data yields coefficients that are more negative in the later period than in the early data. This result is reversed for any break point after 1988.

One possible explanation for this result is that productivity gains in some regions have offset the anticipated wage response not picked up by the regional trend terms. Naturally, it is also possible that there is not enough data available for the current expansion to identify the response of detailed wages to unemployment. Tentatively, 1991 stands out as the point where wage-setting changes might have had an impact on the wage curve; but this change is not consistent with the standard macroeconomic definition of greater wage flexibility.

3.2 Building in productivity growth and inflation

The general panel controls used in BNQ suggest some additional factors that could be controlled for, such as productivity and inflation. Simply adding these controls to the regressions often makes the results more confusing, possibly because they are more likely to have substantial errors-in-variable problems. Smoothing the series can help, which is part of the approach followed by Staiger, Stock and Watson (2001) (SSW).

SSW use panel regression on US states to estimate a US Phillips curve.[10] Their approach is also suitable for evaluating when the labour market might be more flexible. Their implicit measure of the NAIRU should respond to the wage structure becoming more flexible. Compared to BNQ, the underlying motivation is similar, but these authors impose more restrictions on the functional form of the relationship, by starting with real unit labour costs. Their primary analysis is based on the following form:

$$\omega_{jt} - \rho_{jt} - \pi_t = \beta u_{jt-1} + \boldsymbol{\delta}_j \mathbf{D}_j + \boldsymbol{\delta}_t \mathbf{D}_t + \upsilon_{it} \qquad \textbf{(4)}$$

where ω is the rate of change of the nominal wage, less current productivity growth (ρ) and inflation (π).[11] Note that the regional trend is dropped from the specification, since the primary reason for including a trend—regional productivity patterns—is now directly accounted for.

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[10] SSW use these estimates to produce a time-varying estimate of the NAIRU, which would also be possible here. Greenslade, Pierse and Saleheen (2001) have shown that aggregate Kalman-filter estimates for the UK yield evidence of a declining NAIRU.

[11] Productivity growth figures are derived from published output per worker figures for the closest regional parallel available. Inflation is the annual percentage change in the retail prices index between NES surveys.

Table C shows panel estimates for SSW-style wage equations across ten British regions. There are several ways in which these estimates might differ from US estimates other than the substantive differences between the US and UK labour markets. Because we have detailed claimant count information by region, the unemployment rate is not affected by sampling variation in the Current Population Survey-based measures, which are highly variable in small states. We do not smooth these figures, as we are confident that they accurately reflect conditions within each region. On the other hand, productivity figures are difficult to calculate for UK regions and do vary substantially from period to period. We smooth these estimates using a Hodrick-Prescott filter, which yields a smooth, time-varying estimate similar to the Kalman filter used in SSW.

As a check on the importance of regional productivity figures, these regressions are also run using wages adjusted for the smoothed, national trend in productivity growth. These calculations are the best parallel with BNQ, because the time dummies in BNQ would absorb national productivity figures even when they are excluded. The coefficient on unemployment will be different primarily because (as in SSW) the unemployment rate is not entered in logs. Finally, subtracting these time-varying aspects of the model from the dependent variable, rather than including an estimate of them in the explained variation, lowers the R^2.

The coefficient on the unemployment rate (regional claimant count) is negative and significant: -0.25 with regional productivity adjustment and -0.17 with national productivity adjustment. The coefficient is larger primarily due to the unemployment rate being entered (untransformed) in levels, but these estimates also have relatively smaller standard errors compared to earlier estimates. The reasonably tight standard errors for the specification with regional productivity are encouraging, given the noisiness of the underlying regional productivity data. On the basis of the strength and reliability of the wage response to unemployment, this is the favoured specification, with a *t*-statistic of 3.8. Compared to the earlier BNQ estimates, this is quite high. This is primarily due to the change in functional form, because the estimates based on UK productivity growth should be (other than the functional form) similar to the BNQ estimates.

Including regional productivity figures in the SSW estimation reveals a major break in the relationship between wages and unemployment—after 1992, the estimated relationship is essentially zero. The estimated coefficient prior to 1992 is slightly smaller than the full-sample result, rather than larger as we might have anticipated if the full-sample results were simply an average of the pre and post-1992 results. This result is economically and statistically insignificant when national productivity figures are applied, implying that regional productivity patterns are important to this result.

The robustness of this result to alternative dates is assessed in Chart 3. The coefficient on unemployment for the first half of the period is always negative and statistically significant, but there is a clear trend toward a less negative relationship as more recent data is added to the early sample. The later-period estimates also rise until about 1990, when the coefficient is essentially zero. Compared to Chart 2, based on the BNQ approach, the BNQ estimates are more likely to find a statistically insignificant early-period coefficient on unemployment, while the SSW estimates find an even more negative estimate using only the first few years of data. The SSW specification generally results in smaller standard errors relative to the coefficients, although having specific regional productivity trends clearly matters. This break is not evident in Table C without the regional productivity trends. This lack of break point when national productivity figures are used is true regardless of the dating of the potential break point.

Both of the results based on BNQ and SSW are consistent with a changed labour market in recent years, but these changes are inconsistent with the increased responsiveness of wages to unemployment rates. Overall, the macro-econometric evidence, while seemingly indicating that something has changed, is not particularly informative about the nature of the change.

4. Micro evidence: wage rigidities

The micro perspective begins with a *notional wage change* – that is the change an employer would desire for a particular job within their firm in the absence of wage rigidities. This model follows Altonji and Deveraux (1999). These notional wages respond to the macroeconomic factors highlighted in the model above, but would in addition be expected to vary quite substantially from person to person within the firm. Defining a notional wage change helps to clarify how wages might be less than fully flexible at the individual level. Let notional wages be

$$w_{it}^{*} = x_{it}\beta + \alpha_t + \varepsilon_{it} \qquad \textbf{(5)}$$

where x_{it} refers generally to controls particular to the individual, and α_t is included to account for aggregate factors (like the unemployment rate and productivity growth) in the notional wage contract. The error term, ε_{it}, accounts for individual heterogeneity in every time period. Conceptually, the notional wage change can be thought of as desired wage variability, and not a recording or firm error. [12]

With the notional wage defined, the precise form of the restricted wages follows from the nature of the restriction. Altonji and Deveraux (1999) consider nominal rigidities of particular form. If the

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[12] Groshen and Schweitzer (1999) estimate the size of firm errors and treat a rising relationship with inflation as evidence of a 'sand' effect of inflation in labour markets. This would complicate this model as inflation rose.

notional wage change is less then zero but more than -χ, then a constrained wage change of zero is realised. If the notional wage change is less than -χ, then the firm actually pays according to the schedule:

$$\begin{array}{ll} w_{it}^{0}-w_{it-1}^{0}=x_{it}\beta+\alpha_{t}-w_{it-1}^{0}+\varepsilon_{it} & if \quad 0\leq x_{it}\beta+\alpha_{t}+\varepsilon_{it}-w_{it-1}^{0} \\ w_{it}^{0}-w_{it-1}^{0}=0 & if \quad -\chi<x_{it}\beta+\alpha_{t}+\varepsilon_{it}-w_{it-1}^{0}<0 \\ w_{it}^{0}-w_{it-1}^{0}=x_{it}\beta+\alpha_{t}-w_{it-1}^{0}+\varepsilon_{it} & if \quad x_{it}\beta+\alpha_{t}+\varepsilon_{it}-w_{it-1}^{0}\leq-\chi \end{array}$$

where w_{it}^{o} is the realised wage based on the notional wage of w_{it}^{*}. The parameter χ in this equation is a free parameter to account for workers and firms agreeing that an extreme event requiring a reduction in wages has occurred. It could be well below zero. Unemployment might occur if the wage is constrained—depending on how the firm is modelled—but should not occur when the notional wage is paid.

Altonji and Deveraux (1999) further augment this specification with an error term for the measurement of wages. This is an important issue, since there is the potential for error in most micro-based data sources and the identification of wage rigidities ultimately depends on the existence of spikes or reduced observations below zero, both of which are sensitive to errors. Reporting errors will typically reduce the size of the spike and increase the fraction below zero, unless they are correlated across years.(13) Fehr and Goette (2000) use the Altonji and Deveraux (1999) estimator on Swiss social security data (which is similar to NES data) to ascribe most of the

(13) Smith (2001) finds that the spike is potentially boosted by measurement error in a household survey where many observations are rounded and thus more likely to be unchanged between years.

negative wage change observations to errors. The NES data ought to be fairly reliable, but it is certainly worth analysing some of the identifiable sources of error and being careful to use techniques that are not overly sensitive to errors.

Altonji and Deveraux (1999) focus solely on downward nominal rigidities, but it is not difficult to extend their set-up to study real wage rigidities. A lower and an upper boundary needs to be established. These are identified by the parameters ϕ and γ.

$$\begin{array}{lll} w_{it}^{0} - w_{it-1}^{0} = x_{it}\beta + \alpha_t - w_{it-1}^{0} + \varepsilon_{it} & if & \gamma \le x_{it}\beta + \alpha_t + \varepsilon_{it} - w_{it-1}^{0} \\ w_{it}^{0} - w_{it-1}^{0} = r, \quad -\phi \le r \le \gamma & if & -\phi < x_{it}\beta + \alpha_t + \varepsilon_{it} - w_{it-1}^{0} < \gamma \\ w_{it}^{0} - w_{it-1}^{0} = x_{it}\beta + \alpha_t - w_{it-1}^{0} + \varepsilon_{it} & if & x_{it}\beta + \alpha_t + \varepsilon_{it} - w_{it-1}^{0} \le -\phi \end{array} \qquad \textbf{(6)}$$

An optimising model of why firms might deviate from the notional wage change can be found in MacLeod and Malcomson (1993). The firm's acceptance of restrictions on wage changes is supported by information and renegotiation costs, which could allow either nominal or potentially real rigidities to exist. Furthermore, there is no reason a more complicated model allowing for two types of contracts (for example, nominal and real rigidities) could not be patterned on the MacLeod and Malcomson model.

4.1 Distributions

Before moving on to statistical tests, it is useful to see what the data suggest regarding key restriction points. Chart 4 shows the distribution of hourly wage changes for men continuing in the same job, excluding overtime pay and unusual hours. The bars shown in Chart 4 are 1 percentage point wide.[14] The prominence of the spike at zero varies from year to year, but is always present and is clearly larger in low-inflation periods. In 1994, the fraction of the workforce with no wage change reached over 6%. Interestingly, the spike at no wage change is just as evident in the 1970s and it is quite high relative to the bars to the right and left of zero. This type of information can be useful in identifying nominal rigidities. It may indicate that wage rigidities were just as common – or even more common – in the 1970s, but that the high inflation rates made this form of rigidity less relevant.[15]

There may be similar tendencies for wage changes to be primarily above the inflation rate, resulting in either a spike or a clustering of observations just to the right of the relevant inflation rate. However, defining the relevant inflation rate when wage changes may have been agreed upon at any time in the last year makes this form of rigidity particularly hard to identify. The vertical lines in these charts are drawn for the inflation rate from December to December and February to February, in the previous year. The variety of possible inflation references points, and the complexity of response will make this hypothesis more difficult to evaluate. In any case, the spike just above the inflation rate in 1979 or 1985 in Chart 4 is potentially indicative of this form of rigidity. The chart for 1982 offers an interesting contrast, in that wages seem to be concentrated just below the inflation rate.

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(14) The microeconomic model of wage rigidities is fundamentally ambivalent on what rate of pay is relevant, yet this can matter a great deal to the measurement of wage rigidities. When we compared three different definitions of wages available in the NES we found that the qualitative features of the wage-change distributions, including spikes at zero and the inflation rate were largely unchanged. An appendix describing these results is available from the author.

(15) While not shown, 1999 does not stand out from other nearby years. It is true that the distribution has more mass just to the right of the inflation than either 1998 or 2000, but the difference is small and unlikely to disrupt any of the tests proposed below.

The other aspect of a rigidity, whether nominal or real, is that there ought to be less mass to the left of zero relative to the right, once the natural decline in frequencies away from the centre is accounted for. The evidence for this seems stronger for real rigidities than for nominal rigidities. The distribution in 1979 shows a substantial drop in mass left of the December and February inflation rates. On the other hand, the 1970s and early-1980s distribution has many bars between the inflation rate and zero, and seems to be quite smooth through zero once the spike has been excluded. It is, however, difficult to visually evaluate the fraction below zero or the inflation rate in these histograms, because it is difficult to evaluate how much the frequency should have declined due to being further from the median wage change.

Nonetheless, the existence of the spike at zero is *prima facie* evidence of wage rigidities; but the nature and full extent of the rigidity is unclear. We would not expect to see rounding (as an error) in this data set large enough to generate this spike. It could be that firms do adjust pay rates less frequently than once a year, but this would still be interpreted as a wage rigidity on an annual basis.

4.2 Kahn's measures

Kahn (1997) offers an alternative approach to measuring wage rigidities that does not require detailed functional form restrictions.[16] Using Kahn's method, estimates are based on deviations from the expected fraction of observations of a given amount left of the median associated with

[16] There is a literature of tracking moments of the distribution of wage changes, beginning with McLaughlin (1994). Summary measures of wage flexibility, for example skewness or the more sophisticated measures offered by Lebow, Stockton and Wascher (1995) and McLaughlin (1999) were also explored. Unfortunately, these measures failed to

being at or near zero wage change in the histograms shown in Chart 4. This is a general approach that relies on few distributional assumptions and is open to extension, but it may be susceptible to shifts in the form of the underlying notional wage change distribution. This weakness applies quite generally to micro-econometric measures as this underlying distribution – no matter how it is approximated – forms the basis of identifying the alternative in any micro approach. Kahn’s method handles reporting errors in essentially the same way, by assuming that the distribution of errors change little between years.

Specifically, Kahn’s test takes the other years’ empirical distributions as the implicit comparison. Computing a histogram with a given window width (1 percentage point in this and Kahn’s original work), the fraction in the *n*th percentile to the left of the median is identified by r_1 to r_n. Using these values as the dependent variables, Kahn estimates a panel regression of these frequencies on n, and whether the observation is 1% or 2% left of zero, zero wage change, or just above zero.

$$r_n = f\left(n, r_n \subset \Delta w < 0, r_n \subset \Delta w = 1, r_n \subset \Delta w = 2, r_n \subset \Delta w = -1, r_n \subset \Delta w = 0\right) \quad \textbf{(7)}$$

offer a consistent interpretation of the UK data.

To simplify the specification let $D0$ be a dummy variable that is 1 when the histogram range r_n includes zero wage change, ie $r_n \subset \Delta w = 0$. Similar dummy variables are defined for percentage changes including 1, 2, and -1 and denoted $D1$, $D2$, and $DN1$ respectively. Finally, fixed-effects for being a given amount less than the median provide the estimates of typical wage change frequencies.

$$prop_{rt} = \alpha_r + \beta_1 DNEG_{rt} + \beta_2 D1_{rt} + \beta_3 D2_{rt} + \beta_4 DN1_{rt} + \gamma D0_{rt} + \mu_{rt} \quad \textbf{(8)}$$

All of the useful variation in this regression comes from these dummy variables moving through the distribution, so only r_n – where at one time one of these dummy variables is positive – is included in the regression.(17)

Because the Kahn approach uses variation in other years to define the baseline, we cannot estimate a degree of wage rigidity for each year, so instead we split the sample into two periods. This will indicate whether rigidities are becoming more or less pronounced relative to the underlying distribution, although the timing of the change is somewhat ambiguous. Ideally, we would focus on the degree to which wage cuts are rare, but this measure may be prone to being falsely low if the distribution of reporting errors is symmetric. This follows because there will always be more mass in the true distribution above zero than below with or without wage rigidities, since the median wage change is always likely to be above zero in the United Kingdom due to inflation and productivity

(17) Kahn's (1997) proportional estimator accounts for the non-linear restrictions implied by estimation of a density, but significantly complicates estimation and the extension of the model. The results for the proportional model of nominal rigidities were very similar to the linear model.

growth. The size of the spike is also susceptible to being reduced by errors, but if only some reports are in error, the spike should still be recognisable.

This approach reveals statistically significant nominal wage rigidities in the hourly wage change distributions. The test results shown in Table D are based on the data revealed in Chart 4. The level of the nominal rigidity is lower than Kahn finds in the Panel Study of Income Dynamics (PSID), measured by the coefficient on the dummy variable for including a nominal zero wage change (*D*0). She typically obtains coefficients well above 1 for her all-worker estimates. This may be due to the nature of either the wage data and how it is collected, or the institutions in place.

When the sample is split, we find much stronger evidence of nominal wage rigidities in the 1992-2000 period: the coefficients on the dummy variable for including zero wage change in the observation (*D*0) is more than doubled in the later period. This may not be surprising in light of the fact that zero spikes are far larger after 1992 as seen in Chart 4, although much of the added height of those spikes is due to them being closer to the median. Nevertheless, it still suggests an increase, rather than a decrease, in wage rigidities.

The regressions also reveal that there are statistically significant reductions in the frequency of wage cuts (based on the coefficients on *DNEG*). Bars less than zero are about a half percentage point lower than would be predicted based on their distance from the median. When the sample is split, this effect is again more than twice as strong in the later period.

Extending the early and late period break analysis to the Kahn measures results in Chart 5. The conclusion that nominal rigidities (at zero) have not declined in the later period is robust up to any split in the sample up to 1993. After that the results change dramatically to reveal dramatically reduced evidence of nominal rigidity. Unfortunately, this could be an artefact of the limited data available to identify the underlying wage distribution.

This result seems counter to the conclusions of both Smith (2001) or Nickell and Quintini (2000) as each paper concludes that nominal rigidities are always insignificant in UK data. One critical difference in this conclusion is the use of a tighter testing strategy. Previous studies were more focused on how large the spike was relative to the prevailing view of wage setting that argues that wage cuts are uncommon. Thus these papers focused more on demonstrating some flexibility in UK wage-setting patterns. Since our focus is on establishing whether patterns have changed, we need a tighter standard. That said, it remains to be shown using the Kahn approach what level of rigidities are *economically* significant. ‘Small’ nominal rigidities may or may not be important.

4.3 Real wage rigidities

An interesting feature in Chart 4 is the evidence of spikes, or at least a collection of observations, around the inflation rate. While real rigidities have been discussed in the literature as one of the key differences between the US and European labour markets, the nominal wage rigidities literature has avoided exploring the distribution above zero. This is in part because most papers in the nominal rigidities literature aim to connect with the theoretical literature on sources of monetary non-neutrality, but it is also because these features are more difficult to define.

Extending Kahn’s methodology to real wage rigidities involves two major complications. First, the reference value of inflation is not known. Second, wage responses around the inflation rate might be more diffuse, and thus harder to identify. Indeed, spikes and bumps in wage-change distribution around the inflation rate have not been as evident as those at nominal zero. To account for the first issue, we estimate the model for both December and February inflation rates. The data shown in Chart 4 suggest the second complication may not be a problem in this data set, although the effects do vary from year to year.

We add dummy variables for wage changes including the inflation rate, negative real wage changes, and being 1 or 2 percentages up from these values or 1 percentage point below the inflation rate to the Kahn nominal wage rigidity measures. This specification parallels the earlier implementation and is run simultaneously with the nominal controls, since in a low-inflation period more than one effect could occur at wages a given distance from the median.

Specifically, the linear estimation equation becomes

$$\begin{aligned} prop_{rt} &= \alpha_r + \beta_1 DNEG_{rt} + \beta_2 D1_{rt} + \beta_3 D2_{rt} + \beta_4 DN1_{rt} + \gamma D0_{rt} \\ &+ \beta_5 DRNEG_{rt} + \beta_6 DR1_{rt} + \beta_7 DR2_{rt} + \beta_8 DRN1_{rt} + \gamma RD0_{rt} + \mu_{rt} \end{aligned} \qquad \textbf{(9)}$$

It is not obvious a priority that this specification will cover most of the cases where wages bunch near inflation rates as seen in Chart 4. In particular, the fact that in some years the apparent concentrations around inflation have moved above or below the rate suggests that this technique will struggle to find a real wage-setting pattern if these moves occur too frequently.

The full-sample estimates show statistically significant evidence of real rigidities, based on the coefficients on *DR*0. Table E shows estimates using the February inflation rate, but results using the December inflation rate were similar. Interestingly, the scale of these estimates is similar to those accounting only for nominal rigidities. This is also true for the coefficients on negative real wage changes (*DRNEG*), and these coefficients will apply to a larger fraction of wages, particularly in the high-inflation years. The coefficients around the inflation rate do not fit with the menu cost story (collecting small deviations at round figures in the nominal estimates). Instead, these nearby wage-change proportions (*DR*1, *DR*2, and *DRN*1) tend to show increases (when statistically significant), which is consistent with using different time frames for inflation, as these deviations will tend to be small in most years.

Splitting the sample yields a similar picture for the early and late periods. The evidence does not support a reduction in wage rigidities, as evidenced by the rising coefficient on wage changes, *DR*0. Comparing across the full range of possible splitting points (in Chart 6) shows that these estimates are generally higher in the later period. The coefficients for bars near zero real wage change are also typically negative, along with the coefficient on negative real wage changes. In the linear specification, the other boosted categories are nominal zero, and just below the inflation rate. Nominal zero frequencies are actually little changed from the Table D results, suggesting that the apparent rise in nominal rigidities in the later period was associated with a reduction of real wage rigidities. Generally, including real wage controls does not substantially alter the nominal-only estimates. Together, these tables show little movement toward a more flexible labour market, while nominal rigidities may have increased.

Real wage rigidities potentially represent a large wedge between notional wages and actual outcomes, when inflation is positive. This characterisation is weakened in part because the evidence for reduced real wage cuts, as reported in Table E, is sometimes small or insignificant in when the estimates of the effect *at* the inflation rate are large. This may not be critical however, because the indicator variables for nominal wage cuts and real wage cuts are largely overlapping and may not be identified well enough to distinguish these effects. These estimates are suggestive of some real wage rigidities, although a more structural approach is required to better untangle the evidence.

5. Conclusion

This paper has sought to collect both macroeconomic and microeconomic evidence on pay flexibility from a consistent data source, the New Earnings Survey Panel. In the 1990s, regional wage rates do not appear to be more responsive to regional unemployment levels. This is true both for estimates based on the approach of Bell, Nickell and Quintini (2000) and that of Staiger, Stock and Watson (2001). Nonetheless, these estimates do seem to reveal some important changes in wage adjustment over the period. Estimates focusing on the 1990s reveal no statistically significant aggregate wage response to regional unemployment levels. The overarching conclusion is that the macroeconomic tests leave much to explain, but the estimates reveal some patterns that are worth trying to reconcile with other sources of evidence. Both sets of estimates suggest focusing on the beginning of the 1990s.

The microeconomic evidence reveals clear evidence of nominal wage rigidities. These rigidities are just as evident in the later data, and indeed are found to be stronger up until 1993. The sharp turn in 1993 must be treated with some caution, since the approach used may require more data to define the underlying change distribution. The evidence of nominal rigidities, particularly in the later periods, is a surprise when compared with the popular view that the UK labour market has become gradually more flexible. The nominal wage rigidity coefficients are somewhat smaller than Kahn showed using PSID data for the United States over the full sample period, but the values for current estimates are similar. The issue for any estimate of rigidities is to consider their effect on employment. It is beyond the scope of this paper to say whether these rigidities are economically meaningful, but they are statistically significant. Indeed, one interpretation of the observation of continued rigidities and the recent shift in Phillips curve estimates is that wage rigidities are not the key source of the shift in estimated Phillips curves. At this point, this conclusion is premature, although these results do point to the need for a better reconciliation of micro wage rigidities and unemployment patterns. The NES data are suitable for a parametric analysis of rigidities and the costs they impose along the lines of Altonji and Deveraux (1999), which might help in the pursuit of these issues.

Finally, this paper reveals a significant pattern of *real* wage rigidities. Some noticeable clustering around the inflation rate is evident in the raw wage change distributions. The Kahn (1997) approach is extended to explore parts of the wage change histogram around the inflation rate. The estimated level of real wage rigidities is found to be steady or even rising in the 1990s. Of course, the distortion caused by real wage rigidities is likely to be smaller under low inflation rates.

These results leave the correspondence between alternative wage flexibility measures puzzling, but it does point out the critical gap that exists between micro and macro definitions. The evidence from Phillips curves suggests that changes have occurred that might have helped to keep wage inflation down. Two directions for further work have been suggested by this research. The macro approach pointed out the importance of productivity differences, suggesting that a better model of notional wages in the micro approach might be informative. It also became clear that the micro approach needed to be formulated such that the cost of wage rigidities, whether nominal or real, can be evaluated.

Finally we note that the results and conclusions in this paper are based on data from the New Earnings Survey for the period 1975-2000. Naturally, the results might change if the sample period were updated to the most recent period using data from the successor to the New Earnings Survey, the Annual Survey of Hours and Earnings.

Tables and charts

Table A: Adjusted regional wages and unemployment

	Log Wage Rate		Log Wage Changes	
Lagged regional log wage	0.7942 (0.0617)	0.9150 (0.0317)		
ln(unemployment rate)	-0.0143 (0.0051)	-0.0058 (0.0042)	-0.0079 (0.0050)	-0.0073 (0.0043)
Region Dummies	p=0.00	p=0.01	p=0.46	p=0.39
Region Trends	p=0.15		p=0.00	
Time dummies	p=0.00	p=0.00	p=0.00	p=0.00
N	10	10	10	10
NT	250	250	250	250
R-squared	0.9999	0.9999	0.9847	0.9842

Source: Author's calculations using 2001 NES Panel data set.
Wage rates and changes are regional averages calculated according to the first-stage regression shown in equation (1). Standard errors are shown in parentheses.

Table B: Adjusted regional wages and unemployment

	Log Wage Rate	Log Wage Changes
Lagged regional log wage	0.7707 (0.0573)	
ln(unemployment rate)	-0.0131 (0.0044)	-0.0065 (0.0047)
Post 1992 * ln(unemployment rate)	0.0170 (0.0048)	0.0121 (0.0056)
N	10	10
NT	250	250
R-squared	0.9999	0.9850

Source: Author's calculations using 2001 NES Panel data set.
Wage rates and changes are regional averages calculated according to the first-stage regression shown in equation (1). Standard errors are shown in parentheses.

Table C: Wage growth beyond inflation and productivity growth

	Regional Productivity Growth		UK Productivity Growth	
Unemployment rate	-0.25 (0.07)	-0.21 (0.05)	-0.18 (0.04)	-0.18 (0.04)
Second half of sample * Unemployment rate		0.21 (0.05)		0.02 (0.04)
Region dummies	p=0.00	p=0.02	p=0.01	p=0.06
Time dummies	p=0.00	p=0.00	p=0.00	p=0.00
N	10	10	10	10
NT	240	240	250	250
R-squared	0.93	0.93	0.96	0.96

Source: Author's calculations using 2001 NES Panel data set.
Wage changes are regional averages calculated according to the first-stage regression shown in equation (1). Standard errors are shown in parentheses.

Table D: Kahn measure: frequency of percentage pay changes

	Full Sample	1976-1991	1992-2001
Within zero band	1.26	0.75	2.24
	(0.10)	(0.14)	(0.11)
Nominal cut	-0.49	-0.43	-1.08
	(0.06)	(0.09)	(0.12)
Within +1% band	-0.71	-0.46	-1.21
	(0.10)	(0.14)	(0.11)
Within +2% band	-0.53	-0.51	-0.48
	(0.10)	(0.14)	(0.12)
Within -1% band	-0.17	-0.05	0.03
	(0.10)	(0.15)	(0.08)
N	24	24	23
NT	600	384	207
R-squared (within)	0.408	0.207	0.903

Source: Author's calculations using 2001 NES Panel data set.

Table E: Kahn-style tests for real rigidities

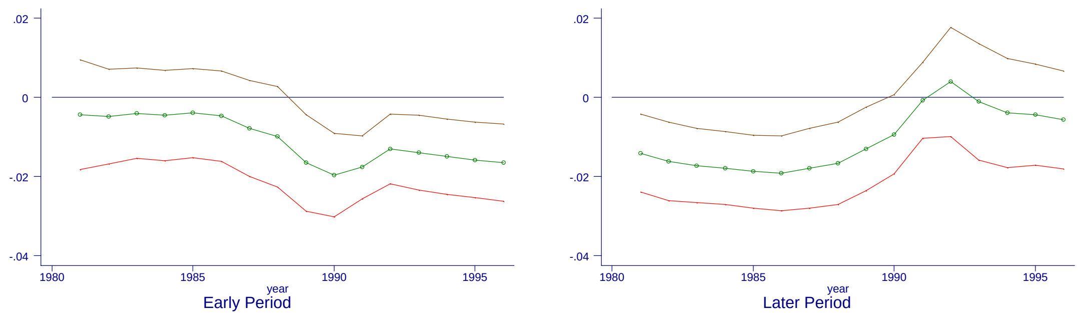
	Full Sample	1976-1991	1992-2001
	Linear	Linear	Linear
Within zero band	1.46	0.79	2.07
	(0.13)	(0.15)	(0.20)
Nominal cut	-0.34	-0.40	-1.33
	(0.08)	(0.09)	(0.22)
Within +1% band	-0.46	-0.40	-1.36
	(0.13)	(0.15)	(0.19)
Within +2% band	-0.29	-0.43	-1.35
	(0.13)	(0.14)	(0.18)
Within -1% band	-0.13	-0.04	0.02
	(0.13)	(0.15)	(0.16)
At inflation rate	1.74	0.78	1.14
	(0.14)	(0.15)	(0.22)
Real cut	0.12	-0.19	-1.58
	(0.12)	(0.12)	(0.25)
1% above inflation rate	1.15	0.42	0.51
	(0.14)	(0.15)	(0.19)
2% above inflation rate	0.59	0.46	-0.18
	(0.13)	(0.15)	(0.16)
1% below inflation rate	-0.03	0.01	0.32
	(0.14)	(0.15)	(0.18)
N	37	37	37
NT	925	592	333
R-squared (within)	0.339	0.214	0.747

Source: Author's calculations using 2001 NES Panel data set.

Chart 1: Adjusted wages and unemployment: Great Britain

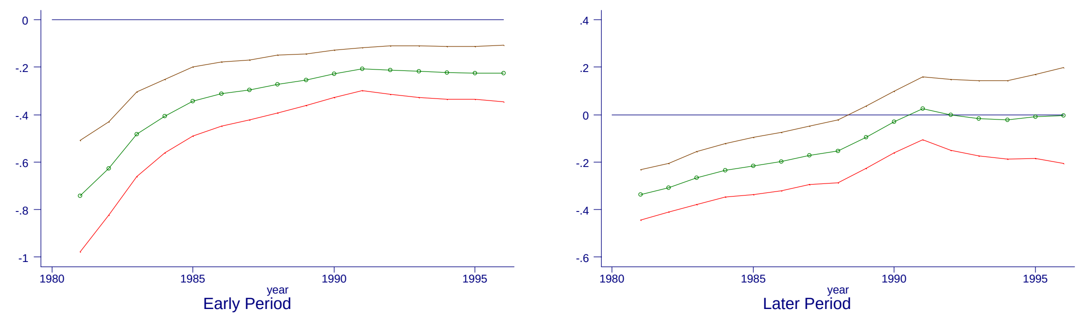
Source: Author's calculations from 2001 NES Panel data set. Wage changes are averages calculated after controlling for individual characteristics.

Chart 2: Coefficient stability in BNQ model: alternative break points



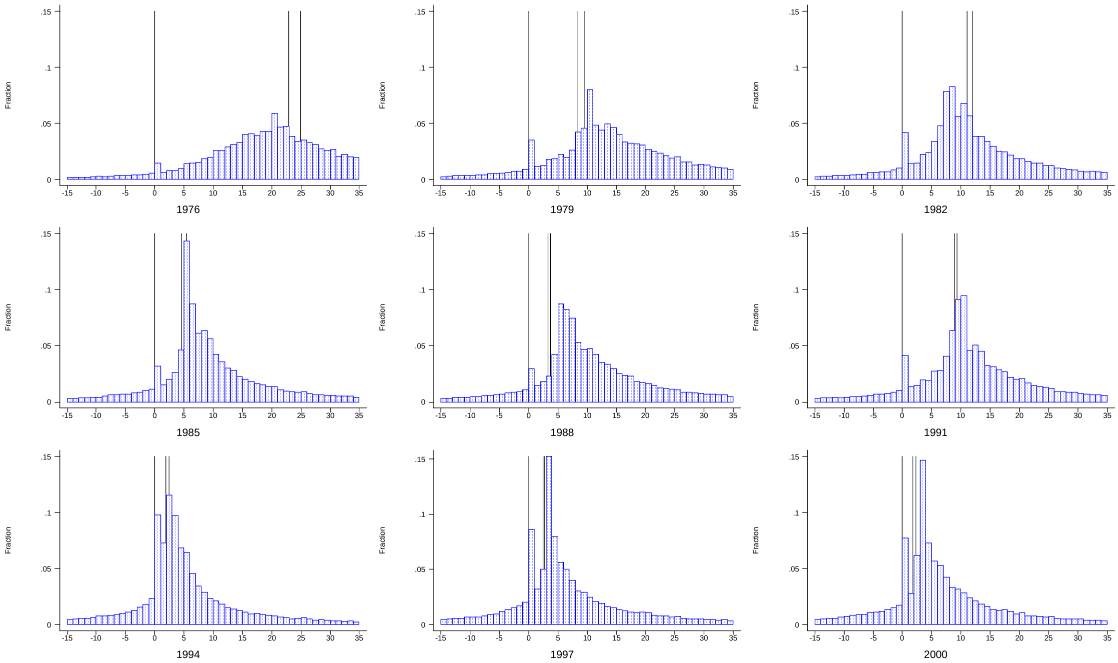
Source: Author's calculations using 2001 NES Panel data set.

Chart 3: Coefficient stability in SSW model: alternative break points



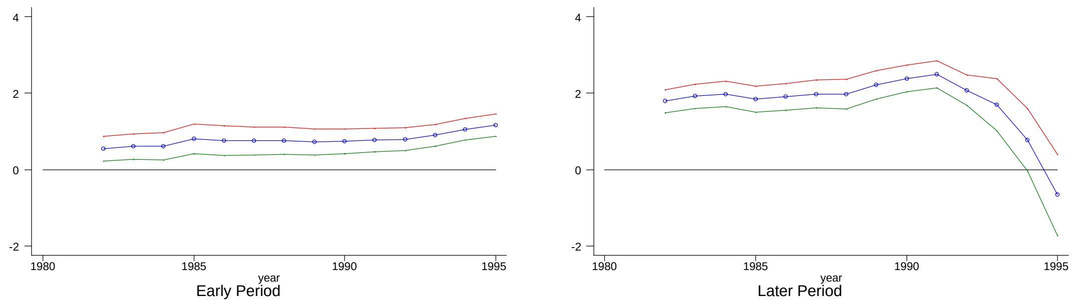
Source: Author's calculations using 2001 NES Panel data set.

Chart 4: Wage change distributions



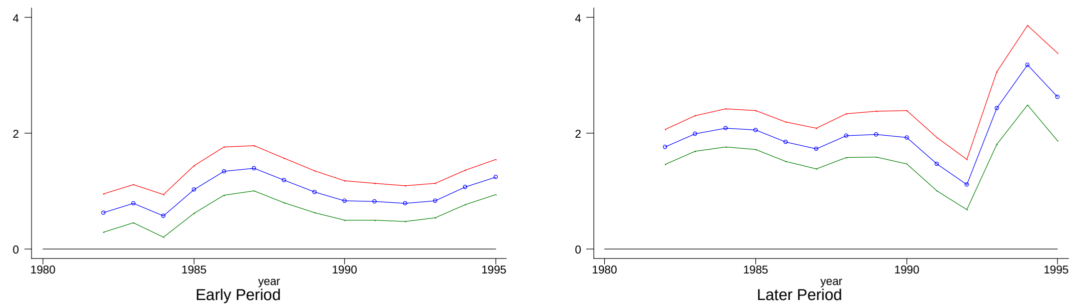
Source: Author's calculations using 2001 NES Panel data set.

Chart 5: Coefficient stability at nominal zero: alternative break points



Source: Author's calculations using 2001 NES Panel data set.

Chart 6: Coefficient stability at real zero: alternative break points



Source: Author's calculations using 2001 NES Panel data set.

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