REAL WAGE CYCLICALITY OF JOB STAYERS, WITHIN-COMPANY JOB MOVERS, AND BETWEEN-COMPANY JOB MOVERS

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Abstract: Using the British New Earnings Survey Panel Data (NESPD) for the period 1975 to 2001 we estimate the wage cyclicality of job stayers (those remaining within single jobs in a given company), within-company job movers, and between-company job movers. We also examine how the proportion of movers varies over the business cycle. We find that the wages of internal movers are slightly more procyclical, and wages of external movers considerably more procyclical, than those of stayers. The greater cyclicality of movers is particularly apparent for private sector workers and persons uncovered by collective agreements. Notwithstanding, a decomposition shows that in Britain, wage cyclicality arises predominantly from the procyclicality of wages for job stayers, with across- and within-firm mobility playing a lesser role. Thus, there is little evidence for rigid wage models that imply that employers use changes in job titles as a means of adjusting wages to the business cycle.

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Increased flexibility is a prime objective of government labor market policy in many economies. Against a recent background of relatively sluggish economic performance, it has been especially emphasized by policy makers of member countries within the European Union. Central to the goal of flexibility has been the attainment of overall wage cyclicality as it determines the extent to which adverse shocks result in wage adjustment rather than in the loss of jobs. Total wage adjustment derives from three primary sources. These are the wage changes of workers within single jobs, workers moving between jobs within the same company, and workers moving between different companies.

Evidence from panel microdata shows that real wage changes of between-company movers are more procyclical than wages of within-company stayers (Bils 1985; Shin 1994). Also, real wages of all job movers – that is within- and between- company job movers combined – have been found to be more procyclical than wages of job stayers (Hart 2005). However, these studies fall short of providing detailed evidence concerning the process of internal real wage cyclicality. Is within-company wage cyclicality mainly the result of internal promotions and demotions, with wage stickiness prevailing within individual jobs? Or does product and labor market competition require that within-job wages also respond to prevailing market conditions?

We investigate the relative importance of wage adjustments of job stayers, internal movers, and external movers within the British economy. Britain is generally regarded as enjoying the most flexible labor market among the main European economies. We make use of a rich panel data set, the New Earnings Survey Panel Data (NESPD). It contains a random 1% sample of British workers in employment. The data provide highly accurate individual wage and hours statistics taken from employers' company payroll records. Our period of analysis is 1975 to 2001.

This paper adds to the wage cyclicality literature in several ways. First, the previous literature that distinguishes between internal and external moves is largely based on case studies and the extent to which their results generalize is in some doubt. By using a nationally representative sample of workers, we get results that apply to more than just individual companies. Second, our use of modern data from Britain complements the existing literature that is based on U.S. data. Third, because we observe employer changers in addition to job changers, we can decompose overall levels of wage cyclicality into within-job, within-employer across job, and across-employer components, something that has not previously been implemented in the literature.

Wage Cyclicality of Stayers and Movers

There is little previous literature on wage cyclicality that distinguishes between internal and external mobility. Using data from the Ford and Byers companies from the 1920s and 1930s, Solon, Whatley, and Stevens (1997) find that the bulk of wage cyclicality in these two companies was a result of workers changing job titles rather than changing wages within a job title. In contrast, Wilson (1997), using recent data from two U.S. companies, finds no evidence that the wages of position changers are more cyclical than the wages of position stayers. She finds mixed evidence for the hypothesis that the rate of position changing is procyclical.[1] As pointed out by Solon et al., the within-job/between-job dichotomy within companies is a potentially important dimension for research into wage cyclicality and one that would benefit from more up to date and comprehensive data.

Why might wage cyclicality differ between job stayers, job movers, and employer movers? In a competitive spot market for labor in which human capital is general and wages of all workers are fully flexible and adjust in line with marginal revenue product, there are no clear reasons to expect differences in the wage cyclicality of these groups.

However, wages may be governed by implicit contracts rather than a spot market. Malcomson (1999) summarizes this literature and describes how both risk-sharing and human capital investment motives may lead to wages being less flexible than a spot market. In risk-sharing models, risk-averse workers may be insured by employers against fluctuations in their wage income.[2] Similarly, implicit contracts may be used to reduce

[1] Devereux (2000) carries out a similar exercise to Wilson using the Panel Study of Income Dynamics (PSID). This exercise is limited by a short sample period (1981-1992) and the noisy information on internal job mobility in the PSID.

[2] Malcomson discusses three types of insurance contract all of which serve to a greater or lesser extent to constrain wage responsiveness to current market conditions. The parties to a fully binding contract agree that, over the contractual spell of employment, the real wage will be set to reflect the market conditions that prevailed when the contract was initially drawn up. Alternatively, if the contract is non-binding on the worker the real wage remains constant unless the firm believes that prevailing market conditions may induce a job quit. If the contract is non-binding on the company the real wage will remain constant unless the firm believes that market conditions are such that it will be cost-effective to layoff the worker. Empirical work in North America has found that contracts that are non-binding on the worker are especially influential on wage behavior (Beaudry and DiNardo 1991; McDonald and Worswick 1999; Grant 2003).

transaction costs or avoid holdup problems when specific or general human capital acquisition is important (Malcomson, 1999; Hashimoto 1979; Aoki 1984). Since implicit contracts imply some detachment of the wages of job stayers from current labor market conditions, the wages of company changers may be more procyclical than those of company stayers.

Less attention has been paid to why the wages of job stayers might be less procyclical than the wages of job movers *who remain in the same company*.[3] If human capital is job-specific then rent sharing, and its associated effect of blunting wage responsiveness to market conditions (Hashimoto 1979), may also be primarily job-specific. Within company job moves would then involve losses of specific capital, and wage changes associated with internal job changes may more directly reflect marginal revenue product and, hence, current business cycle conditions. This would imply that the wages of job changers (even within companies) are more procyclical than the wages of job stayers.[4]

Recent empirical work on the importance of implicit contracts in Britain (Devereux and Hart 2005) has shown that the spot market plays the dominant role in real wage determination but also finds evidence consistent with wage contracts that are non-binding on the worker.

[3] In a life cycle context, we know from the work of McCue (1996) that position moves within companies are potentially important; they are estimated to account for 15% of male wage growth in the U.S.

[4] A similar argument suggests that internal movers will have less cyclical wages than external movers. Rents associated with specific human capital may derive from both job- *and* company- level knowledge acquisition. As emphasised by Aoki (1984) collective, or company-specific, skills may lead to organizational rents that are shared between workers and entrepreneurs. In this event, moving between jobs within the company would lead to a partial loss of specific capital while changing companies would lead to total loss.

An alternative model developed by Reynolds (1951), Reder (1955), and Hall (1974) assumes that wage levels within job titles are unresponsive to the demand conditions faced by firms. Therefore, employers respond to the business cycle by transferring workers between job titles so as to adjust labor costs appropriately. For example, in expansions firms lower promotion and hiring standards and hence lower the average quality of worker in each job title. Consequently, real wages per quality unit of labor rise even if real wages within job titles are rigid. Similarly, in a recession, firms increase promotion and hiring standards and thus reduce the wage per unit quality. The model predicts that a significant proportion of overall wage cyclicality results from workers changing job titles rather than wage changes within job titles. This arises either because the rate of job title changing is procyclical or because the wage changes of internal movers are more procyclical than the wage changes of job stayers.

Data

The New Earnings Survey Panel Data (NESPD) is comprised of a random sample of all individuals whose National Insurance numbers end in a given pair of digits. Each year a questionnaire is directed to employers, who complete it on the basis of payroll records for relevant employees. The questions relate to a specific week in April. Since the same individuals are in the sample each year, the NESPD is a panel data set that runs from 1975 to the present. Because National Insurance numbers are issued to all individuals who reach the minimum school leaving age, the sampling frame of the survey is a random sample of the population. Employers are legally required to complete the survey questionnaire so the response rate is very high. Also, individuals can be tracked from region to region and employer to employer through time using their National Insurance numbers.

The questions in the NESPD refer primarily to earnings and hours of work. Since the data are taken directly from the employer's payroll records, the earnings and hours information are considered to be very accurate. The wage measure we use is "gross weekly earnings excluding overtime divided by normal basic hours for employees whose pay for the survey period was not affected by absence."[5] We deflate wages using the British Retail Price Index as it is the UK's most widely used price index and is similar to the U.S. Consumer Price Index (CPI). The NESPD also includes information on age, sex, occupation, industry, and geographic location of individuals (but not education or race). We confine attention to full time workers holding single jobs. Our samples cover 177 thousand males and 112 thousand females. We cannot calculate experience for each individual and so we use age as a regressor in its place.

One concern is that employers may report hours worked inaccurately and this would bias our estimates of wage cyclicality. In particular, employers may report contract hours rather than actual hours worked and this would cause reported hours to remain constant from year to year. The result would be a countercyclical bias for hours and a procyclical bias in hourly wages. While there is no way to validate the hours' reporting in the NESPD, we have examined the stickiness and the cyclicality of reported hours. We found that, on average, 80% of stayers had the same reported weekly hours in two adjacent

[5] We also estimated wage specifications in which hourly earnings (including overtime) replaced hourly standard rates. These produced no substantive changes and so we confine attention to standard rates throughout the paper.

periods (the proportion varied between 61% in 1981/82 and 89% in 1978/79). Thus, there seems to be significant variation in reported hours from year to year, especially given that true hours are likely to remain constant for most stayers. Furthermore, reported hours are significantly procyclical for both men and women over our sample period, once again suggesting that any biases from misreporting of hours may not be very large. Overall, we believe our hours data are at least as good as the self-reported data from individual surveys but a thorough examination of this issue will require a new system of collecting hours information.

Our business cycle proxy is the national claimant count unemployment rate produced by the British Office for National Statistics.[6] Wage agreements in Britain typically cover a 12 month period and so the wage measures in the NESPD generally refer to wage settlements negotiated between April, when the samples are taken, and May of the previous year. Accordingly, we use as our unemployment rate measure the average of the 12 monthly unemployment rates between May of the previous year and the survey month of April.

Between one April census and the next, the NESPD provides a very clear distinction between job stayers and job movers. A question in the Survey records whether an employee has remained in a given single job within the company for more than 12 months or less than 12 months. This information allows us accurately to identify job movers, defined as individuals who have *either* changed jobs within the same company *or*

[6] Our main reason for choosing claimant count data is that they allow us to obtain consistent monthly data back to 1975.

changed companies. [7] For two consecutive years of NESPD data we have complete information that allows us definitively to separate internal and external movers. Before describing our method of determining this mover dichotomy for the remaining years, it is useful to report key information for these two years.

For 1996 and 1997, we know precisely whether each job move has taken place within companies or between companies. We consider the sample of individuals who are employed at the survey date in both periods (the unemployment rate is about 8% in both years). Between the two years, 92 percent of male workers and 91 percent of female workers remained in the same job. Of the movers, 50 percent of males and 55 percent of females changed job within the same company. Thus, internal mobility is quantitatively as important a phenomenon as the much more heavily studied external mobility. Let us define 'no wage change' in real basic hourly wage rates between the two years as a wage in 1997 that remained within the bounds of the 1996 wage by +/- 1 percent. Then, for both genders, the modal groups of job stayers experienced a wage increase -- 51 percent of males and 53 percent of females. But wage reductions also occurred for significant

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[7] The questions used to determine this variable are as follows [bold type as used in questionnaire]:

(a) What, if any, is the employee's full job title **and** rank or grade? [BOX TO FILL IN]

(b) Give a short description of the work this employee does. For engineers and accountants state professional qualifications, if any. [BIGGER BOX TO FILL IN]

(c) Has the employee worked in this **same job in your organisation** for one year or more?

(If the employee has changed to a different job or been promoted within the last 12 months then 'Under one year' is appropriate.) Circle 1 or 2.

One year or more 1

Under one year 2.

numbers of stayers -- 29 percent of males and 27 percent of females. In the case of between-company job movers, wage reductions affected 34 percent of males and 22 percent of females. [8]

Apart from 1996 and 1997 a direct breakdown of individuals into within- and between-company job moves is not possible. We need, therefore, to identify such moves indirectly. Let M_t denote a binary variable indicating that a job move has taken place at time t. We can obtain M_t from the NESPD. Let $M_t = M_{Wt} + M_{Bt}$ where M_{Wt} denotes a within-company job move and M_{Bt} denotes a between-company job move. In order to identify M_{Wt} and M_{Bt} we adopted the following decision rules:

(1) $M_{Wt} = 1$ if M_t does not intersect with M_{Bt}

$= 0$ otherwise.

(2) $M_{Bt} = 1$ if M_t involves a change in geographical area and/or industry and/or sector

$= 0$ otherwise.

We chose three sets of combinations of area, industry, and sector to identify M_{Bt} in (2):

(A) 10 standard British regions, 1-digit industries and public/private sector;

(B) 97 geographical areas, 1-digit industries and public/private sector;

(C) 97 geographical areas, 3-digit industries and public/private sectors.

Table 1 shows actual and estimated job moves and real wage changes for the years 1996 and 1997 using (A), (B) and (C). All three correctly identify about 75 percent of all moves. The public/private sector split is common to all choices.[9] By moving from (A) to

[8] The prevalence of downward wages in Britain is well known. For example, Nickell and Quintini (2003) find that significantly larger proportions of British workers experience nominal wage cuts or unchanged nominal wages compared to their U.S. counterparts.

[9] The public sector covers workers in central government, local government, and public corporations.

(C), one classifies more of the moves as being external and fewer as being internal. Choice (A) correctly picks out over 80% of within company movers but incorrectly classifies 35-40% of external moves as internal. Disaggregating regions into 97 sub-areas and industries to a three digit breakdown – i.e. choice (C) – reverses the relative predictive balance in favor of between company movers. Choice (B), consisting of 97 areas combined with 1-digit industries, produces a reasonably even balance and correctly classifies about 75% of moves. These findings are very similar for both males and females.[10]

In the lower part of Table 1 we compare actual and estimated real basic hourly wage changes (i.e. excluding overtime) between 1996 and 1997. Real wages rose by an average of 6.5 percent for male stayers. Of course, actual and estimated real wage changes coincide in the case of stayers. Mean real wage changes among both types of movers are over twice as large, albeit accompanied by considerably larger standard deviations. Both first and second moments are well estimated by each of our three mover identifiers although choice (A) appears to provide marginally the best estimates of the actual means.

Table 2 presents summary statistics, based on our complete data set, for the key variables underlying the subsequent analysis. Note that 90 percent of males and 88 percent of females are job stayers. Thus females are slightly more likely to undertake

[10] Consider classification scheme (B) for males. Here, 447 internal movers are wrongly classified as external movers. 50% of these moved region, 38% moved industry, 2% changed sector, 9% changed region and industry, 0.2% changed region and sector, and 1% changed industry and sector.

internal and external job moves compared to males. This may reflect less contractual security in female compared to male jobs.[11]

The table also shows how the proportions of movers and stayers vary depending on whether (A), (B), or (C) is used. In line with the reported findings in Table 1, Table 2 also shows that the mean real wage changes (expressed in logarithms to conform with our estimating equations) are greater for both types of movers compared to stayers.

Estimation

Decomposition of Overall Wage Cyclicality

Here, we extend the analysis of Solon, Whatley and Stevens (1997) to the case where there is information on across-company mobility in addition to within-company mobility. Let P_W and P_B denote the proportion of workers changing jobs within and between firms, respectively. Let $E(\Delta \ln W_S)$, $E(\Delta \ln W_W)$, and $E(\Delta \ln W_B)$ be the expected wage growths of job stayers, within-company movers, and between-company movers, respectively. Overall expected wage growth is given by

[11] The similarity in rates of external mobility between men and women is consistent with prior literature. For the UK, on the basis of the work-history data from the British Household Panel, Booth et al. (1999) investigate job tenure and job mobility of men and women from 1915 to 1990. They find that men are more likely to leave a job involuntarily and are likely to display a higher propensity to quit their jobs in order to take up alternative employment than women. Women show higher propensities to leave their jobs for other reasons. In general, they find that job insecurity was greater for men than women. However, and most relevant to the data period of the present study, these differences narrow appreciably in the later cohorts of these data. For the US, Blau and Kahn (1981) find in the National Longitudinal Surveys of young men and women that the genders display similar tendencies to quit their jobs.

(3) $$E(\Delta \ln W) = (1 - P_W - P_B) E(\Delta \ln W_S) + P_W E(\Delta \ln W_W) + P_B E(\Delta \ln W_B)$$
$$= E(\Delta \ln W_S) + P_W E(\Delta \ln W_W - \Delta \ln W_S) + P_B E(\Delta \ln W_B - \Delta \ln W_S).$$

Differentiating (3) with respect to the change in the unemployment rate, ΔU, provides a decomposition of total wage cyclicality, that is

(4) $$\partial E(\Delta \ln W)/ \partial(\Delta U) = \partial E(\Delta \ln W_S)/ \partial(\Delta U)$$
$$+ P_W[\partial E(\Delta \ln W_W - \Delta \ln W_S)/\partial(\Delta U)]$$
$$+ P_B[\partial E(\Delta \ln W_B - \Delta \ln W_S)/\partial(\Delta U)]$$
$$+ [E(\Delta \ln W_W - \Delta \ln W_S)]\partial P_W /\partial(\Delta U)$$
$$+ [E(\Delta \ln W_B - \Delta \ln W_S)]\partial P_B /\partial(\Delta U).$$

The first term is the wage response of job stayers (individuals who remain in the same job in the same company). The second term defines the incremental effect on wage cyclicality of internal movers relative to job stayers. Likewise, term three defines the incremental wage cyclicality of external movers relative to job stayers. The last two terms represent, respectively, the cyclicality of internal and external job changes. So, three terms comprise wage responses and two job move probabilities. We deal with wage and job effects in turn.

Estimating wage cyclicality

The empirical work constitutes a simple extension of the approach of Solon, Whatley and Stevens (1997). It incorporates the two-step estimation procedure of Solon, Barsky and Parker (1994) (see also Devereux 2001) designed to get round the problem of using individual wage and other characteristics alongside a national-level cyclical indicator (Moulton 1986); the associated year-specific error is likely to result in OLS overestimating the precision of the unemployment rate coefficient.

In step 1, we estimate the wage change equation for an individual i at time t. This is given by

$$(5)\quad \Delta \ln w_{it} = \alpha_0 + \alpha_1 A_{it} + \sum_{t=1}^{T} \phi_{0t} D_t + \sum_{t=1}^{T} \phi_{1t} M_{Wit} D_t + \sum_{t=1}^{T} \phi_{2t} M_{Bit} D_t + \varepsilon_{it}$$

where w_{it} the real standard hourly wage rate, A_{it} is a cubic in age, D_t denotes a dummy variable equal to 1 if the observation is from year t, and ε_{it} is a random error term. The $M_W D$ and $M_B D$ terms represent interactions between the time dummies and the mover dummies shown in (1) and (2).

In step 2, the three sets of dummy variable estimates $\hat{\phi}_{jt}\ \left(j = 0,1,2\right)$ are regressed on the change in the unemployment rate and a linear time trend (from 1 to 26), or

$$(6)\qquad \hat{\phi}_{jt} = \delta_{j0} + \delta_{j1}\Delta U_t + \delta_{j2} Year_t + \nu_{jt}. \qquad \left(j = 0,1,2\right)$$

Estimation of (5) is undertaken using OLS and the second step regression, equation (6), is estimated by weighted least squares (WLS) where the weights are the number of individuals observed in a given year.[12] In all regressions, the change in the log wage is multiplied by 100. The estimated coefficient on the change in the unemployment rate then approximates the percentage change in the wage for a one-point increase in the unemployment rate.

[12] Instead of WLS, we could do Generalized Least Squares (GLS) in the second stage using the estimated variance-covariance matrix of the year dummies in the first stage. We have verified that this approach gives coefficient estimates and standard errors that are virtually identical to WLS, so we use the simpler approach.

We can link (6) directly to the decomposition of wage cyclicality in (4). Using $\hat{\phi}_{0t}$ in (6), the estimated value of δ_{01} gives the cyclical wage response of job stayers. This is the first term on the right-hand-side of (4). Using $\hat{\phi}_{1t}$ and $\hat{\phi}_{2t}$ in (6), we obtain estimates of δ_{11} and δ_{21}; that is the incremental wage effects of within and between company job movers relative to job stayers. These are reflected in the second and third terms of (4).

Estimating cyclicality of internal and external job moves

We also estimate the cyclicality of internal (job to job within the same company) and external (company to company) moves. These comprise the fourth and fifth terms in (4). We use the same basic two-step approach, replacing Δlnw_{it} in equation (5) with the binary variables in (1) and (2) that indicate, respectively, between and within job changes. Specifically, our estimating equations take the form

$$(7) \quad M_{Kit} = \alpha_0 + \alpha_1 A_{Kit} + \sum_{t=1}^{T} \phi_{Kt} D_t + \nu_{Kit} \qquad (K = W, B)$$

and

$$(8) \quad \hat{\phi}_{Kt} = \delta_{K0} + \delta_{K2} \Delta U_t + \delta_{K2} Year_t + \nu_{Kt}. \qquad (K = W, B)$$

In line with the wage specifications, we estimate (7) using weighted least squares thereby using a linear job change probability model.[13]

[13] An alternative would be to use a probit or logit specification. We use the linear probability model to be consistent with the approach of Solon et. al. (1997), and also because it allows us to take a 2-step approach to deal with the clustering issue that is analogous to our approach with wages. We have verified that we get similar marginal effects if we take a probit approach.

Composition of Movers over the Business Cycle

Our estimates will be biased if there are systematic differences in the types of individuals who move over the business cycle that are not accounted for by the fixed individual effects and age variables in the estimating equations. For example, if movers during a boom are predominantly people whose productivity is increasing, and movers during a recession are predominantly people whose productivity is falling, we would have a procyclical bias for movers, and countercyclical bias for stayers. For both males and females, we have calculated that the average internal and external mover comes from between the 60^{th} and 65^{th} percentile of their respective wage distributions, both during periods of rising and falling unemployment. While this similarity across the cycle is suggestive that composition bias is not a large problem, it is not possible to be definitive.

Results

Results based on our full NESPD male and female data are reported in Table 3. We confine attention to the unemployment rate change coefficients, estimated in step two of our regressions. The table contains two sets of results. The first refer to wages and the unemployment rate (equation (6)), and the second to job moves and the unemployment rate (equation (8)). For both sets, we show estimates based on our three methods of distinguishing between internal and external movers ((A), (B), and (C) in Table 1).

Referring to the top half of the table, there are three main findings in respect of absolute and relative real wage cyclicality. First, both male and female stayers' wages are strongly procyclical. A one point reduction in the unemployment rate among male job stayers is associated with a 1.73 percentage real wage increase. The equivalent wage change for females is 1.66 percent. Second, the real wages of between company job movers display significantly higher cyclicality than those of job stayers. For male and female external movers, a one point reduction in the unemployment rate is associated, respectively, with a 2.9 and 2.5 percent wage increase. Third, male and female within company job movers also exhibit stronger wage procyclicality than job stayers. The increments are decidedly modest when compared with the external mover outcomes. The wage responsiveness to a one point change in unemployment is in the order of 10 percent higher for male internal movers compared to stayers, and about 15 percent for females.[14]

Since wage cyclicality of external movers is greater than that of internal movers, one would expect that mis-classification would cause us to understate the cyclicality of external movers and overstate the cyclicality of internal movers. This has implications for how one might expect the estimates to differ across our (A), (B), and (C) splits. Our designated internal movers are more likely to include external movers in (A) than in (B) and especially (C). Thus, we might expect to find greater cyclicality for internal movers using (A) than using (C). Our designated external movers are more likely to include internal movers in (C) than in (B) and especially (A). Thus, we might also expect to find greater cyclicality for external movers using (A) than using (C). In actuality, we find very little evidence for these types of patterns, suggesting that the bias is not strongly related

[14] The estimate for internal male movers in case (C) is not significant. This may be due in part to our inability to obtain consistent 3-digit industry data across the entire time period. We use three different 3-digit classifications for 1975-81, 1982-95, and 1996-2001 and so the internal/external mover definition is not fully consistent across time. Additionally, movers in 1982 and 1996 are dropped since the previous years contain a non-matching classification. However, we have verified that if we include all years by using 1-digit industry codes for 1981-82 and 1995-96, the point estimates change very little.

to the degree of misclassification, maybe because the classification is working fairly well in all cases.

The bottom half of the table reports job move/unemployment rate associations. Estimated procyclicality is stonger for external compared to internal job movers. What accounts for this difference? Figure 1 plots the estimated time dummies from equation (7) against the change in the national unemployment rate (del U). The graphs are based on the (B) set of results and are not greatly altered if (A) and (C) are chosen. The male and female within-company mover graphs are remarkably similar. They reveal a procyclical pattern in the middle periods, from the early 1980s until the early 1990s. Note, however, that the start and end periods do not exhibit cyclical job movements with internal job changes displaying unbroken year to year declines from the mid 1970s to the mid 1980s. This occurred *despite* a period of falling unemployment in the late 1970s. This may indicate that during the inflationary conditions and economic uncertainty associated with the OPEC supply shocks of the mid- and late- 1970s, medium term pessimistic outlooks among companies detracted from an atmosphere of more short term expansion and job promotion. Additionally, the sharp unemployment rate declines starting in 1993, followed by relatively low unemployment thereafter, do not appear to have stimulated a growth in internal job changes. In contrast, male and female external job moves are procyclical over a longer time period. In particular, they appear to be more cyclically responsive than internal moves in the early years.

One can rewrite equation (4) in a way that expresses total wage cyclicality in terms of the wage cyclicality of stayers and movers, and the cyclicality of moving behavior:

$$(9) \qquad \partial E(\Delta \ln W)/ \partial(\Delta U) = (1- P_W - P_B)[\partial E(\Delta \ln W_S)/ \partial(\Delta U)] + P_W[\partial E(\Delta \ln W_W)/\partial(\Delta U)] + P_B[\partial E(\Delta \ln W_B)/\partial(\Delta U)] + [E(\Delta \ln W_W - \Delta \ln W_S)]\partial P_W /\partial(\Delta U) + [E(\Delta \ln W_B - \Delta \ln W_S)]\partial P_B /\partial(\Delta U) .$$

Combining the results in Table 3, with the summary data in Table 2, we are in a position to evaluate the separate contributions. Results are slightly different across the choice of mover identifiers, but reporting results for choice (B) (see Table 1) is nonetheless highly representative. Our male estimate of overall wage cyclicality (i.e. $E(\Delta \ln W)/ \partial(\Delta U)$ in equation (9)) is -1.83 percent. Of this aggregate figure, 84.3 percent is accounted for by the wages of job stayers, 6.4 percent by the wages of internal movers, 7.5 percent by the wages of external movers, 0.6 percent by internal job moves, and 1.1 percent by external job moves. The overall female estimate of wage cyclicality is -1.78 percent, with respective percentage breakdowns of 81.4, 8.1, 7.5, 1, and 2. Wages of job stayers are highly procyclical and job stayers account for nearly 90 percent of all observations in our data. Unsurprisingly, therefore, their wage contribution dominates overall British wage cyclicality.

As discussed earlier, one model posits that employers may use promotions and demotions to achieve wage flexibility in spite of the stickiness of wages within jobs. It is clear that this is not the dominant influence in contemporary Britain. Wages within jobs seem sufficiently flexible that internal job mobility plays a relatively minor role in moving aggregate wages in line with the business cycle. Thus, it appears that this class of sticky wage models is not particularly important to Britain.

Comparison of Estimates to the Literature

Studies of wage cyclicality in the U.S. have tended to find overall semi-elasticities of between -1 and -2 for the association between real wage changes and the contemporaneous national rate of unemployment (Solon et al. 1994; Bils 1985).[15] Our estimates for Britain are at the high end of this range. The coefficients we find for job stayers are higher than those that have been reported for employer stayers in the U.S. in recent panel data (Solon et al. 1994; Devereux 2001; Shin and Solon 2004). Hart (2005) splits the sample between stayers and movers using the NESPD and finds similar coefficients for stayers to ours.

The main contribution of this paper is to make the distinction between the wage procyclicality of internal and external movers. There are few studies that make this distinction. Solon et al. (1997) used U.S. historical data and found evidence that a large proportion of wage cyclicality was accounted for by internal job mobility, rather than through the cyclicality of wages of job stayers. Wilson (1997) uses recent data from two U.S. companies and finds no evidence that the wages of position changers are more cyclical than the wages of position stayers. Devereux (2000) gets a similar result using state-year variation in the PSID over the decade of the 1980s. While we find that wages of internal movers are more procyclical than job stayers, job stayers in contemporary Britain have very procyclical wages and the process of internal mobility has little net impact on overall wage cyclicality.

[15] Although, using six cohorts of the National Longitudinal Surveys, Grant (2003) obtains real wage/current unemployment rate semi-elasticities for youth-female, youth-male and young men of between -2 and -2.6.

There is also a related literature on nominal wage rigidity. Most recent studies from the U.S. have concluded that, once measurement error is accounted for, nominal wage changes of stayers are downwardly rigid (for example, Altonji and Devereux 2000; Akerlof et al. 1996). On the other hand, the British evidence suggests that nominal wage cuts are prevalent both in the BHPS (Smith 2000) and in the NESPD (Nickell and Quintini 2003). Measurement error appears unlikely to be the full explanation as Smith finds many cuts even for individuals who report having their pay stub in hand while answering the earnings questions. Our findings that stayers in Britain have greater wage cyclicality than in the U.S. is consistent with these findings from the nominal wage rigidity literature.

Results by Public/Private Sector and by Collective Bargaining Status

In some organizations, promotions and other job changes may be largely based on agreed rules and laid-down formulas. In these cases, the move from one job description to another may not be marked by significant wage increments but merely involve an individual transferring from the top rungs of one ladder across to the bottom rungs of the next higher ladder. Further, such moves may not correlate especially well with market conditions. Other organizations may take much more *laissez faire* approach to job change. Productivity-based promotions may be especially important. Big upward movements for high fliers and demotions for under- performers are likely to be more prevalent in these cases with productivity effects reflecting market conditions.

A priori, two highly interrelated divisions of the data may be expected to capture these general differences in approaches to internal job mobility. The first is the public/private sector split and the second is the division between workers covered and uncovered by collective bargaining agreements. Over all observations in our data, 87 percent of males and 88 percent of females in public sector jobs are covered by collective bargaining agreements. This contrasts with coverage of 28 percent for males and 21 percent for females in the private sector. In general, the terms and conditions of work and pay in the public sector are relatively regularized. First, the size and complexity of large governmental departments and public corporations produce a greater recourse to the use of explicitly defined rules and regulations concerning pay scales. Second, occupational pay and employment conditions are standardized across geographical areas. Third, the prevalence of formal collective bargaining in the public sector reduces the likelihood of *ad hoc* decision making over pay and jobs. The private sector is more heterogenous and less regulated with fewer impediments to the achievement of localized implicit and explicit agreements.

The estimates are in Table 4. In the private sector the wages of both male internal and external job movers are significantly more procyclical than male stayers.[16] This contrasts with males in the public sector where both types of movers exhibit no significantly greater wage effects compared with stayers. Consistent with greater flexibility in the private sector, the difference between private and public sectors is statistically significant for both internal and external movers. The relative picture is similar for females, although the internal mover coefficient for the private sector is not

[16] Note that we only include observations in which the individual is in the same sector at t and t-1. Thus, the results for external moves should be treated with some caution as the group of external movers included are those who chose to move to a different company in the same sector. This is, of course, a selected sample of external movers.

significant at the 5 percent level and, unlike the external mover coefficient, is not statistically different from the equivalent public sector estimate.

The wages of internal and external job movers of both sexes who are not covered by a collective bargaining agreement are also significantly more procyclical than equivalent stayers. By contrast, among workers covered by a collective bargaining agreement, only the wages of male external movers display more cyclicality than stayers. Consistent with uncovered workers having more flexible work arrangements, the wages of uncovered workers appear more procyclical than those of covered workers but the difference is only statistically significant for internal movers (for both men and women).[17]

Unsurprisingly, the intersection of private sector and uncovered reveals very similar patterns to those discussed above for the private sector and the uncovered sector with both kinds of movers having more cyclical wages than job stayers. In contrast, the intersection of public sector and covered shows no significant differences between movers and stayers.[18] Also, the only statistically significant difference between the two groups (private/uncovered versus public/covered) is for male internal movers.

Results by Age

In the final two rows of Table 4, we split the sample at the median age (40 for men, 35 for women) in order to examine how the estimates differ by age. We find that

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[17] Beaudry and DiNardo (1991) and Grant (2001) also find union workers to have lower wage procyclicality than non-union workers in their studies using U.S. data.

[18] The other two intersections – i.e. private sector ∩ covered and public sector ∩ uncovered – are not shown because there are unreliably small numbers of movers in these cases.

there are strong similarities across the two age groups for both men and women. The sole exception is that wages for male internal movers are statistically significantly more procyclical for younger rather than older men. This is consistent with promotions and other internal moves being more likely to occur when young, as would be implied by matching models of the labor market.[19]

Conclusions

In line with earlier studies, our British data demonstrate the value of distinguishing between job stayers and job movers in the study of real wage cyclicality (Hart 2005). Additionally, our work underlines the potential importance of separating movers who change jobs within companies and those who move between companies. In our full samples, external movers exhibit considerably higher wage cyclicality than job stayers – in fact, between 30 and 40 percent higher – while wage cyclicality among internal movers is less markedly higher, at around 10 to 15 percent. When we disaggregate the data into private and public sectors and into workers covered and not covered by collective bargaining then the value added of making the mover distinctions becomes even more apparent. We find that wage cyclicality of both internal and external movers is considerably higher than stayers among private sector workers and those workers uncovered by collective agreements. Thus, it appears that employers who are

[19] We also attempted to see whether the degree of wage cyclicality has changed over time by splitting the sample period into two periods of equal length. We found no significant differences but this may be in part due to the lack of variation in the unemployment rate, particularly in the second of these shorter periods.

less constrained by formal agreements and pay rules are more likely to adjust the wages of internal movers in line with outside economic conditions.

However, these findings should not detract from the overwhelming importance of job stayers in determining total British wage cyclicality. While the relative wage cyclicality of job movers is higher than stayers, the absolute wage procyclicality of *both* stayers and movers is high. Combining this latter observation with the fact that job stayers comprise about 90 percent of all wage observations, we find that about 80 to 85 percent of overall real wage cyclicality in Britain is accounted for by job stayers. These results suggest that sticky wage models that stress the role of job mobility in enabling wages to adjust to economic conditions are not particularly relevant to contemporary Britain.

Table 1 Job moves and real wage changes between 1996 and 1997

Actual and estimated within and between company job moves					
		Estimated job movers			
		Correctly estimated internal movers/total actual internal movers (percent correct)		Correctly estimated external movers/total actual external movers (percent correct)	
Mover identifiers		Males	Females	Males	Females
(A)	10 regions , 1-digit industries, public/private sector	1592/1924 (82.7)	1152/1366 (84.3)	1272/1979 (64.3)	714/1151 (62.0)
(B)	97 areas, 1-digit industries, public/private sector	1477/1924 (76.8)	1070/1366 (78.3)	1457/1979 (73.6)	812/1151 (70.5)
(C)	97 areas, 3-digit industries, public/private sector	1215/1924 (63.1)	843/1366 (61.7)	1723/1979 (87.1)	980/1151 (85.1)

Actual and estimated values of percentage real wage changes (standard deviations)						
	Stayers		Internal movers		External movers	
	Males	Females	Males	Females	Males	Females
Actual	6.5 (24.8)	6.1 (19.4)	14.6 (33.6)	16.5 (31.1)	12.6 (43.2)	15.5 (47.8)
Using (A)	6.5 (24.8)	6.1 (19.4)	14.1 (38.8)	16.2 (35.3)	13.0 (38.8)	15.8 (46.0)
Using (B)	6.5 (24.8)	6.1 (19.4)	13.9 (39.7)	15.5 (31.1)	13.3 (37.9)	16.8 (48.3)
Using (C)	6.5 (24.8)	6.1 (19.4)	14.7 (41.5)	16.4 (25.8)	12.9 (37.0)	15.8 (46.7)

Table 2 Descriptive statistics, 1975 – 2001

	Males			Females		
Mover identifiers	(A)	(B) **Total**	(C)	(A)	(B) **Total**	(C)
Number of individuals (Number of observations)		**177498 (1346612)**			**112502 (644608)**	
Job stayers as proportion of total observations		**0.896**			**0.879**	
Internal movers as proportion of total observations	0.068	0.059	0.050	0.083	0.072	0.06
External movers as proportion of total observations	0.042	0.051	0.067	0.046	0.056	0.076
Mean age (Median age)		**40 (40)**			**37 (35)**	
Mean $\Delta \ln W_S$ (standard deviation)		**0.021 (0.171)**			**0.030 (0.144)**	
Mean $\Delta \ln W_W$ (standard deviation)	0.059 (0.231)	0.057 (0.224)	0.056 (0.213)	0.077 (0.196)	0.075 (0.192)	0.076 (0.182)
Mean $\Delta \ln W_B$ (standard deviation)	0.061 (0.324)	0.062 (0.316)	0.062 (0.303)	0.081 (0.283)	0.082 (0.272)	0.080 (0.260)
Private sector as proportion of total observations		**0.697**			**0.586**	
Public sector as proportion of total observations		**0.304**			**0.414**	
Bargaining coverage as proportion of total observations		**0.462**			**0.497**	

Table 3 Real wage and unemployment rate changes, 1975 – 2001

Mover identifiers	MALES Coefficient on $(U_t - U_{t-1})$			FEMALES Coefficient on $(U_t - U_{t-1})$		
Wage change [equation (6)]	**Job stayers**	**Incremental wage effect for internal movers**	**Incremental wage effect for external movers**	**Job stayers**	**Incremental wage effect for internal movers**	**Incremental wage effect for external movers**
(A)	-1.73**	-0.24*	-1.19**	-1.66**	-0.35*	-0.83**
	(0.45)	(0.12)	(0.18)	(0.46)	(0.15)	(0.16)
(B)	-1.73**	-0.24*	-0.99**	-1.66**	-0.32*	-0.76**
	(0.45)	(0.10)	(0.17)	(0.46)	(0.17)	(0.16)
(C)	-1.73**	-0.03	-1.11**	-1.66**	-0.37*	-0.76**
	(0.45)	(0.12)	(0.16)	(0.46)	(0.14)	(0.16)
Job move [equation (8)]		**Internal job movers**	**External job movers**		**Internal job movers**	**External job movers**
(A)		-0.004*	-0.005**		-0.005*	-0.006**
		(0.002)	(0.001)		(0.003)	(0.001)
(B)		-0.003	-0.005**		-0.004	-0.007**
		(0.002)	(0.001)		(0.002)	(0.001)
(C)		-0.004	-0.009*		-0.005	-0.011**
		(0.004)	(0.004)		(0.004)	(0.004)

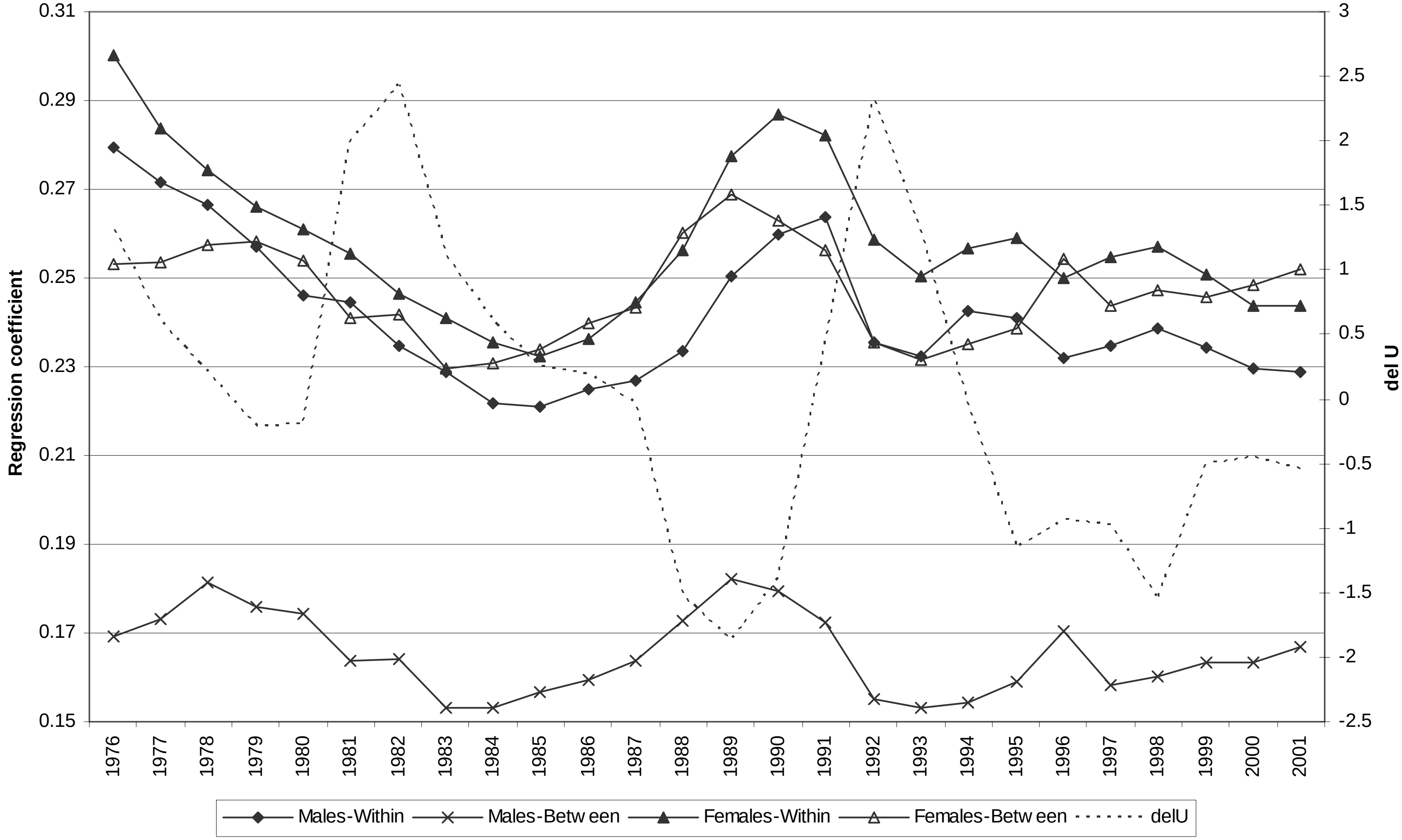
Notes: Standard errors in parentheses. ** (*) denotes significant at 0.01 (0.05) level for two-tail test. Results shown refer to step two of the two-stage estimation procedure. There are 26 observations at this stage. The three-digit industry classification used as part of identifier (C) cannot be obtained on a consistent basis over the entire period. The results are obtained using three different 3-digit classifications for 1975-81, 1982-95, and 1996-2001. Accordingly, movers in 1982 and 1996 are dropped since the previous years contain a non-matching classification.

Table 4 Real wage changes in relation to unemployment changes by sector, collective bargaining coverage, and age 1975 – 2001

	MALES Coefficient on ($U_t – U_{t-1}$)			FEMALES Coefficient on ($U_t – U_{t-1}$)		
	Job stayers	**Incremental wage effect for internal movers**	**Incremental wage effect for external movers**	**Job stayers**	**Incremental wage effect for internal movers**	**Incremental wage effect for external movers**
Private sector	-1.93** (0.37)	-0.35** (0.12)	-1.03** (0.19)	-1.93** (0.33)	-0.38 (0.21)	-0.91** (0.20)
Public sector	-1.39* (0.73)	0.17 (0.23)	-0.12 (0.31)	-1.38 (0.73)	0.03 (0.22)	-0.21 (0.25)
Covered by agreement	-1.50* (0.67)	0.05 (0.20)	-0.45* (0.21)	-1.41* (0.70)	0.16 (0.24)	-0.21 (0.24)
Uncovered by agreement	-1.94** (0.32)	-0.44** (0.16)	-0.84** (0.20)	-1.91** (0.33)	-0.53** (0.18)	-0.68** (0.20)
Private sector and uncovered	-1.98** (0.31)	-0.41** (0.13)	-0.88** (0.22)	-1.98** (0.32)	-0.52** (0.21)	-0.70** (0.21)
Public sector and covered	-1.48* (0.76)	0.29 (0.27)	-0.32 (0.24)	-1.47* (0.76)	0.07 (0.27)	-0.22 (0.25)
Young (below median age)	-1.84** (0.41)	-0.40** (0.12)	-0.91** (0.18)	-1.86** (0.42)	-0.25 (0.20)	-0.65** (0.17)
Old (above median age)	-1.59** (0.50)	0.05 (0.12)	-0.89** (0.17)	-1.44** (0.50)	-0.35 (0.19)	-0.65** (0.22)

Notes: See notes to Table 2. Reported results consist of movements determined by identifier (B) (see Table 1). For job movers, 'private sector and uncovered' means that an individual was in the private sector and uncovered by a collective bargaining agreement in the new job at time t and the old job at time t-1. This matching between the two periods also applies to 'public sector and covered'.

Figure 1 Coefficients of estimated job move dummies against the rate of change of national unemployment: between and within company moves



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