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THE CYCLICALITY OF REAL WAGES WITHIN EMPLOYER-EMPLOYEE MATCHES

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Using data from the Panel Study of Income Dynamics, the author examines the cyclicalities of wages within employer-employee matches for the years 1970–91. Recent research on wage cyclicalities has suggested that wages are very procyclical (tending to rise and fall with economic upturns and downturns), even for workers who remain with the same employer. The author finds, however, that the evidence for wage procyclicalities within the matches he examines is rather weak except for the small group of workers who were paid by piece rate or commissions. Despite having acyclical wage rates, men who were paid hourly had earnings movements that were very procyclical. Salaries exhibited little cyclicalities, but salaried workers who had income sources from bonuses, commissions, or overtime had procyclical earnings. The results suggest that the increasing prevalence of incentive-based pay will increase the procyclicalities of wages within matches.

Recent research has concluded that wages and earnings are much more flexible than previously thought. Unlike most of the older literature, which generally used aggregate data, microdata studies have turned up evidence suggesting that wages since 1970 have been strongly procyclical. While real wage cyclicalities have been found to be particularly strong for workers who change employer, researchers have found that a one point increase in the unemployment rate decreases the average hourly earnings of workers who do not change employers (stayers) by about 1%. This evidence casts doubt on the notion

that wages within matches are sticky and slow to change with productivity and outside labor market conditions. The goal of this paper is to examine in detail the nature of wage and earnings cyclicalities within matches.

At least three features of my analysis set it apart from previous studies that have considered the wage cyclicalities of job stayers. First, I focus on the cyclicalities of wages within employer-employee matches rather than the cyclicalities of wages of workers who do not change employers. These concepts differ because workers may hold more than

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A data appendix with additional results and copies of the computer programs used to generate the results presented in the paper are available from the author at Department of Economics, University of California, 405 Hilgard Ave., Los Angeles, CA 90201.

one job at a given time. Second, I estimate separate wage cyclicity parameters for workers who were paid hourly, salaried workers, and workers paid in some other fashion. Third, for hourly and salaried workers, I attempt to differentiate between the cyclicity of the base wage or salary, and the cyclicity of income from overtime and from incentives such as bonuses. Because of data limitations, and because the literature has concentrated on men, I limit my sample to men for the analysis.¹

Theory suggests that cyclical wage elasticities should be at least as large for movers as for stayers. In a spot market model, wages equal marginal product and, since productivity is procyclical, wages should be procyclical. In this model there is no real distinction between the wage changes of workers who remain with the same employer and the wage changes of workers who change employer. Incentive models also suggest that wages will adjust in line with workers' marginal product. A different class of models posits the existence of implicit employer-employee contracts that imply non-equality between wages and marginal product within matches. Risk-averse employees may be shielded from productivity shocks, and the wages of workers who do not change employer may thus not exhibit much cyclicity.² However, in these models, workers may gain from economy-wide increases in productivity by changing employers and getting a higher wage at a new employer. Still other models suggest that job changers should have procyclical wages because the jobs obtained by job changers in recessions are not as good as the jobs they obtain in booms. Thus, on numerous counts it is reasonable to expect

the wages of job changers to be procyclical. However, it is less clear that the wages of stayers should exhibit substantial procyclicality.

Of the many reasons why it is important to understand the nature of wage cyclicity within matches, two particularly stand out. First, there is a growing literature on wage determination within the black box of the firm. The adjustment of wages, salaries, and incentive pay over the business cycle within matches is an important but under-researched aspect of firm behavior. Second, the development of meaningful macroeconomic models depends crucially on an understanding of cyclical behavior in the labor market. Recent findings suggest that wage procyclicality should be a property of all models. This paper provides insight about whether it is important for researchers to distinguish between, on the one hand, wage cyclicity that arises because of matches starting or ending, and, on the other hand, the cyclical behavior of wages within matches.

Data

The data set I use in this study is the Panel Study of Income Dynamics (PSID). For my purposes, this data set is superior in several ways to alternatives such as the Current Population Survey (CPS) and the National Longitudinal Survey of Youth (NLSY). First, unlike the NLSY, it contains observations on people of all ages. This is of particular importance in a study of job stayers because of the high rate of job changing among young people. Second, PSID data enable me to identify workers who remain with the same employer from one year to the next.³ By comparison, in the

¹Solon, Barsky, and Parker (1994) found very small estimates for women in the PSID and discussed other small estimates in the literature. I find similarly small levels of procyclicality for women. (Results available on request.)

²See Azariadis (1975), Baily (1974), Hashimoto (1981), Raisian (1983), and Beaudry and DiNardo (1991) for more details on implicit contract models. Rosen (1985) and Malcomson (1997) have provided surveys of this literature.

³I determine the starting and ending dates of jobs using algorithms developed in Altonji and Williams (1997) and Altonji and Devereux (2000). There are many inconsistencies in the tenure data in the PSID (see Brown and Light (1992)). Any errors in determining whether or not a worker changed jobs will tend to bias my estimates of wage cyclicity for job stayers in a procyclical direction, because job changers have more procyclical wages.

CPS one must compare industry and occupation codes in order to determine if a person has changed employers. Third, the PSID covers a long time period—my sample runs from 1970 to 1992 and so is well suited to business cycle analysis. The only restrictions I place on observations are that persons must be aged between 18 and 64 and must work at least 100 hours in the calendar year. The cyclical variable I use is the national unemployment rate (the unemployed as a percentage of the total civilian labor force). This variable has been widely used in studies of wage cyclicalities.⁴

The PSID contains information on labor earnings, hours worked, and average hourly earnings for the previous calendar year. It also has the hourly wage rate at the interview date for hourly workers. For workers paid weekly, monthly, or annually, the PSID reports a wage rate that is calculated by dividing weekly, monthly, or annual salary by a fixed number of hours.⁵ For example, if a worker is paid weekly, the weekly payment is divided by 40 hours irrespective of how many hours he or she works each week. Thus the cyclicalities of this measure represents the cyclicalities of salaries. There is no point in time wage measure for workers, such as piece rate and commission workers, who are neither hourly nor salaried. Thus, for these workers, earnings divided by hours is the only source of wage information. The point-in-time wage measures have the advantage of being specific to the main job the employee works on. On the other hand, the earnings measure includes earnings from extra jobs and other sources of income. Thus, from this perspective, the point in time wage measures are superior to the average hourly earnings measure.

⁴I have also used the unemployment rate for white men aged 20 years and older. The results with this measure are virtually identical to the results using the national civilian unemployment rate.

⁵I use the unemployment rate in the interview year for wage rates that relate to the interview week. I match the unemployment rate for the previous year to average hourly earnings because it relates to earnings in the previous calendar year.

However, the hourly wage measure and the salary variable do not capture income from overtime and income from bonuses.⁶

One problem with average hourly earnings is that any measurement error in earnings or hours is likely to affect the estimated levels of procyclicality. In particular, the clumping of reported hours at 40 hours per week might imply that actual hours are more procyclical than reported hours. Hence the procyclicality of reported average hourly earnings might overstate the procyclicality of true average hourly earnings.⁷

There are topcode problems with both the wage and average hourly earnings measures. Only 0.08% of average hourly earnings observations are affected. However, 2.27% of reported hourly wages are topcoded, and the problem is quite severe for 1976 and 1977.⁸ I impute topcoded values for both of these measures by multiplying the topcoded value by 1.33.⁹ I deflate all nominal amounts using the Personal Consumption Expenditures deflator for Gross Domestic Product (the base year is 1987). The reported hourly wage is available from 1970 onward, and so the sample I use covers wage changes between 1970–71 and 1990–91.

There are some variables in the PSID that facilitate the examination of the cyclicalities of different sources of income. After answering questions about earnings in the previous calendar year, respondents were asked, “Did you earn any income from overtime, tips, or bonuses that is not in-

⁶Unfortunately, the information in the PSID does not allow one to cleanly measure bonus and overtime income.

⁷The results of Bound, Brown, Duncan, and Rodgers (1994) using PSID Validation data suggest that changes in earnings divided by hours are replete with measurement error.

⁸Before 1978 wages are topcoded at \$9.98 per hour; from 1978 onward the topcoded value is \$99.98.

⁹I have also experimented with deleting topcodes, replacing topcoded wages with wages predicted using average hourly earnings, and dropping years in which topcoding is severe. The results are not sensitive to the method used to deal with topcodes.

cluded in that?" About 5% of respondents said they did. Unfortunately, many more respondents are likely to have received income from these sources but to have included it in the initial earnings measure. Salaried workers were asked whether they got paid overtime if they worked more hours than they would usually work. I use both of these variables in the analysis. The data appendix contains more details about the construction of the samples and some descriptive statistics.

The Issues

Capturing Cyclical- ity of Wages within Matches

This paper focuses on the cyclical-ity of wages within employer-employee matches rather than the cyclical-ity of wages of workers who do not change employers. These concepts differ because workers may hold more than one job at any given time. Thus, if workers are more likely to hold extra jobs in expansions, the cyclical-ity of average hourly earnings for these workers may differ from the cyclical-ity of average hourly earnings within their main job. Fortunately, workers in the PSID are asked whether they hold any extra jobs in addition to their main job. One method to ensure that earnings variation is purely within one match is to exclude workers who hold extra jobs in either period. However, a more satisfactory method for hourly and salaried workers is to use the wage and salary measures that are specific to the worker's main job. In this way one can get at the cyclical-ity of wages and salaries within matches.

Overtime

A second issue that arises in measuring wage cyclical-ity is the treatment of overtime. The merits of using a wage measure that includes overtime are unclear. If the overtime premium is a compensating differential for having to work long hours, then working overtime does not increase the utility of workers. If, however, workers and firms agree on a package of weekly

compensation and hours of work, it is irrelevant which hours are labeled overtime hours.¹⁰ In this case, one would want the wage measure to include income from overtime. In any case, whether the base wage or salary is procyclical is an important question in itself in regards to understanding the process of wage formation. As discussed earlier, one way in which the average hourly earnings measure differs from the hourly wage or annual salary is that the former includes overtime while the latter does not. For hourly workers, average hourly earnings should be more procyclical than the reported hourly wage, because overtime is procyclical (Bils 1985). I will investigate the extent to which overtime explains the differences in cyclical-ity of the two measures. For salaried workers, the cyclical-ity of the annual salary should be similar to that of earnings, at least for workers who are not eligible for overtime and do not receive bonuses.

Method of Pay

The information in the PSID allows me to separate workers by payment type into workers paid hourly, salaried workers, and workers paid in other ways (through piece rates and commissions, for example). It is likely that the level of cyclical-ity will differ for these different kinds of workers, and so it is useful to study them separately. Also, by separating them I can use the reported hourly wage for hourly workers, and the reported salary for salaried workers, to complement the average hourly earnings measure.

Estimation Methodology

The previous literature has demonstrated the importance of individual heterogeneity to estimates of wage cyclical-ity and has generally used first difference estimators to deal with this (see, for example, Solon, Barsky, and Parker 1994). Following this

¹⁰The results of Trejo (1991) suggest that neither of these competing models of overtime is fully supported empirically.

literature, I regress the change in the log wage on the change in the unemployment rate:

$$(1) \quad \Delta \ln w_{it} = \alpha_1 + \alpha_2 \Delta U_t + \alpha_3 x_{it} + \alpha_4 \text{YEAR}_t + v_t + \varepsilon_{it}.$$

By differencing, one nets out unobserved individual and firm heterogeneity that may be correlated with the error term. The control variables in x_{it} are a cubic in labor market experience. When I do the estimation for stayers only, I also include a cubic in tenure with employer. The YEAR variable is a linear time trend and U_t is the national unemployment rate. Using extra controls for worker characteristics has little effect on the results, and so I use this parsimonious specification.

As noted by Moulton (1986), the standard errors from the ordinary least squares (OLS) estimates of equation (1) may be underestimated in the presence of a year-specific error (v_t), because the change in the unemployment rate is the same for all workers in the same year. For this reason I use a two-step estimation method that has been used by others in this literature (Solon, Barsky, and Parker 1994; Shin 1994). In the first step, the estimation equation is as follows:

$$(2) \quad \Delta \ln w_{it} = \beta_1 + \beta_2 x_{it} + \sum_{t=1}^T \phi_t D_t + \varepsilon_{it}.$$

In equation (2), D_t is a dummy variable equal to one if the observation is from year t . In the second step the coefficients on the year dummies, $\hat{\phi}_t$, are regressed on the change in the unemployment rate and a time trend:

$$(3) \quad \hat{\phi}_t = \delta_1 + \delta_2 \Delta U_t + \delta_3 \text{YEAR}_t + v_t.$$

I weight each second step observation by the number of individual observations in that year.¹¹ In all regressions, I multiply the change in log wages by 100 so that the

cyclical coefficient approximates the percentage change in the wage that results from a one point increase in the unemployment rate.

In addition to this least squares strategy, I estimate equation (2) by median regression and carry out the second step estimation exactly as described above. Median regression is more robust than mean regression with respect to extreme observations such as topcoded wage data.

One issue that the first difference framework does not deal with is selection into the labor force and into the match based on any time-varying error. Unfortunately, satisfactorily dealing with this problem would require variables that affect the workers' probability of leaving a match or leaving the labor market but do not affect the workers' wage in the match. I do not have such variables at my disposal. Attempts to deal with selection in the wage cyclical literature have been generally unsatisfactory. See Abraham and Haltiwanger (1995) for a discussion of this issue.

Some General Results

In Table 1, I present results from the estimation of equation (3) using different samples of men. The wage measure is average hourly earnings. For each sample, I present the coefficient on the change in the unemployment rate for the 1970–71 to 1990–91 period. In column (1), the results are for the full sample. In column (2), I exclude the self-employed. In column (3), I include only employees who did not change their main employer, and in column (4) I include employees who did not switch employer and had no extra jobs in either year. The estimates in column (1) for the full sample indicate that over the full 22-year period, there was substantial procyclicality of average hourly earnings. The coefficient on the change in the unemployment rate is -1.16 (0.21). This suggests that a one point increase in the unemployment rate reduces average hourly earnings by about 1.2%. Figure 1a plots the estimated coefficients on the year dummies against the change in the unemployment

¹¹Instead of the two-step method, one could estimate consistent parameter values and standard errors using a single-stage Generalized Least Squares Estimator. Amemiya (1978) showed that when GLS is used in the second stage of the two-step model, the two methods are equivalent.

Table 1. Effect of Change in Unemployment Rate on Change in Log Average Hourly Earnings, 1970–71 to 1990–91.^a
(Standard Errors in Parentheses)

Variable	(1)	(2)	(3)	(4)
		<i>Full Sample Less Self-Employed</i>	<i>Job Stayers</i>	<i>Job Stayers with No Extra Jobs</i>
	(N: 68,463)	(N: 52,525)	(N: 42,164)	(N: 31,957)
<i>Mean Regression</i>				
$U_t - U_{t-1}$ ^b	-1.16** (0.21)	-1.01** (0.21)	-0.81** (0.20)	-0.54** (0.17)
<i>Median Regression</i>				
$U_t - U_{t-1}$	-0.82** (0.15)	-0.70** (0.15)	-0.62** (0.16)	-0.54** (0.15)

^aAll specifications are estimated using the two-step method described in the text. In all specifications there are 22 observations in the second-stage regression. In addition to the change in the unemployment rate, each specification includes a linear time trend and a cubic in labor market experience. The specifications that include only stayers also contain a cubic in tenure with the employer.

^b U_t refers to the national civilian unemployment rate in year t .

**Statistically significant at the 0.05 level.

rate. The procyclicality of average hourly earnings is very clear, with high values of the year dummy coefficients tending to accompany falls in the unemployment rate. When the self-employed are excluded, the estimated coefficient falls somewhat (Table 1, column 2).

In column (3) of Table 1, we see that the average hourly earnings of workers who did not change employer were procyclical. The coefficient on the change in the unemployment rate is -0.81 . This procyclicality can also be seen graphically in Figure 1b. This result is consistent with others in the literature. Solon, Barskey, and Parker (1994) reported a point estimate of -1.2 for stayers in the PSID. This compares to my estimate of -0.8 (Table 1, column 2). However, when I include the self-employed (as they did) and estimate over their sample period (1969–87), I get a coefficient of -1.1 , which

is very similar to their estimate of -1.2 . This coefficient is also very similar to the findings of Shin (1994) in the National Longitudinal Survey (NLS).¹² Similarly, Bowlus (1993) estimated that the effect of the national unemployment rate on wages for stayers is -1.2 in the NLSY. Thus, my wage cyclicity estimates for average hourly earnings of stayers are similar to estimates from other research using the PSID, the NLS, and the NLSY.

As discussed earlier, the interest in this paper is in the cyclicity of wages within employer-employee matches rather than the cyclicity of average hourly earnings of workers who do not change employers. These concepts differ because workers may hold more than one job at any particular time. To increase the likelihood that earnings variation is purely within one match, I have excluded workers who held extra jobs in either period and re-estimated the cyclicity equations. The results are in column (4) of Table 1. The estimate of the coefficient on the change in the unemployment rate is -0.54 , indicating mild procyclicality of average hourly earnings (the coefficient value is about half that reported by previous studies in the literature for stayers). This suggests that the wage cyclicity of stayers overstates the cyclicity of wages within matches because earnings from extra jobs exert a procyclical bias. The median regression results are generally quite similar to the OLS results.

Wage and Earnings Cyclicity within Matches

I now turn to a more detailed analysis of the wage cyclicity of stayers. I investigate cyclicity separately for workers paid hourly, salaried workers, and workers paid in some other fashion. I also investigate how the cyclicity of earnings differs across these three groups.¹³

¹²Bils (1985) estimated coefficients of -0.64 for whites and -0.44 for blacks in the NLS. Shin (1994) re-analyzed the NLS data and estimated a coefficient of -0.95 for stayers.

¹³There are no important systematic changes in the composition of employment by mode of payment

Figure 1. Coefficients on Year Dummies and Change in Unemployment Rate.^a

Figure 1a. Full Sample, Average Hourly Earnings.

Figure 1b. Stayers, Average Hourly Earnings.

Figure 1c. Hourly Workers, Hourly Wage.

Figure 1d. Salaried Workers, Annual Salary.

Figure 1e. Other Workers, Average Hourly Earnings.

^aIn all figures, the left axis is for the change in the unemployment rate and the right axis is for the coefficients on the year dummies.

over the business cycle. I estimated a multinomial model with mode of payment (hourly, salaried, or other) as the dependent variable and the change in unemployment rate and a time trend as explanatory variables. I use non-salaried, non-hourly as the base category. A one point increase in the unemployment rate has no significant effect on the relative risk of

being hourly versus being other or the relative risk of being salaried. Perhaps more interestingly, the time trend results show that each year the relative risk of being hourly declined by 3% and the relative risk of being salaried declined by 2%. This is consistent with other evidence that incentive-based forms of compensation are becoming more widely used.

Table 2. Wage and Earnings Cyclicity of Hourly Workers Who Did Not Change Employer, 1970–71 to 1990–91.^a
(Standard Errors in Parentheses)

Sample (Sample Size)	Mean Regression			Median Regression		
	Hourly Wage	Earnings Divided by Hours	Earnings	Hourly Wage	Earnings Divided by Hours	Earnings
All Job Stayers (19,993)	-0.10 (0.20)	-1.09** (0.26)	-2.34** (0.44)	-0.04 (0.21)	-0.79** (0.19)	-1.31** (0.29)
Stayers with No Extra Jobs (15,544)	-0.01 (0.20)	-0.74** (0.22)	-1.87** (0.42)	-0.02 (0.21)	-0.67** (0.18)	-1.15** (0.26)
Stayers with No Extra Jobs and No Income from Bonus and Overtime (13,533)	0.02 (0.22)	-0.67** (0.21)	-2.07** (0.43)	-0.04 (0.22)	-0.67** (0.21)	-1.24** (0.27)

^aAll specifications are estimated using the two-step method described in the text. In all specifications there are 22 observations in the second-stage regression. In addition to the change in the unemployment rate, each specification includes a linear time trend, a cubic in labor market experience, and a cubic in tenure with the employer.

**Statistically significant at the 0.05 level.

Hourly Workers

In addition to the average hourly earnings variable, the PSID records the hourly wage rate at the interview date for hourly workers. This variable is likely to be reported much more accurately than average hourly earnings, because it is not affected by measurement error in hours. Also, since respondents report their current wage rather than earnings and hours for the previous year, there is less recall bias. Finally, the hourly wage is specific to the worker's current job; average hourly earnings may include earnings from several jobs. For these reasons, the reported hourly wage is the cleanest wage measure available for this sort of analysis. Workers who do not change employer are very unlikely to change their method of pay. Only 8% of stayers report being paid hourly in one period and being paid in a different manner in the next period. Thus, restricting the sample to workers who report being paid hourly in both periods does not constitute a major restriction.¹⁴ Using this sample of hourly

workers, I can compare the cyclicity of earnings divided by hours to the cyclicity of the reported hourly wage.

The estimated coefficients from equation (3) for stayers who are paid hourly are in Table 2. The first results I present are the coefficients on the change in the unemployment rate in regressions of the reported hourly wage and average hourly earnings, respectively. The estimates imply that average hourly earnings were much more cyclical than the reported hourly wage—the absolute value of the coefficient on the change in the unemployment rate is 10 times as large in the average hourly earnings regression. The reported hourly wage appears to be almost acyclical. Figure 1c plots the relationship between the year dummy estimates for the hourly wage and

¹⁴I have investigated the wage cyclicity of workers who changed from being paid hourly to being paid a salary. There is some evidence that they were being

promoted, as they had higher wage changes than workers who remained as hourly workers. However, their wages were no more cyclical than the wages of other stayers, and including them in the hourly sample made no difference to the results. Solon, Whitley, and Stevens (1997) found that promotions increased wage cyclicity during the Great Depression era. Analyses by Wilson (1997) and Devereux (2000) using recent data found no such effect.

the change in the unemployment rate. As the estimates suggest, there is no apparent relationship between the two series.

Several possible factors could explain the difference between estimates for the hourly wage and average hourly earnings. One possibility is that the estimates differ because the average hourly earnings measure includes income from extra jobs. Excluding the cases in which the worker held extra jobs in either period reduces the cyclicalities of average hourly earnings somewhat, but average hourly earnings remain more cyclical than the hourly wage. Another possibility is that the coefficients differ because income from bonuses and overtime is included in average hourly earnings. However, when I restrict the sample to people who said they earned no bonus or income from overtime in either year, the coefficients do not change much.¹⁵ The differing cyclicalities of the two wage measures may arise because of clumping in hours worked that leads to an understatement of the cyclicalities of hours and hence an overstatement of the cyclicalities of average hourly earnings.¹⁶ Because the reported hourly wage is specific to the worker's main job, it is the preferred wage measure. This wage measure exhibits wage procyclicality that is of neither economic nor statistical significance.

¹⁵This restriction is not fully satisfactory because, as discussed earlier, there is reason to believe that some workers who report zero for this variable actually do have income from overtime. From 1977 onward, respondents were asked whether they worked any overtime in the previous year. I have tried restricting the sample of hourly workers to the 1977–91 period and dividing the sample into two groups. The first group is composed of workers who did not work overtime in either year. The second group includes workers who worked overtime in either or both years. The results indicate that overtime explains some but far from all of the divergence in the coefficient on the unemployment rate in the reported hourly wage and average hourly earnings regressions. This is consistent with the conclusions of Bils (1985).

¹⁶I investigated the possibility that the timing of the unemployment rate is inappropriate for the hourly wage. However, I get almost identical results when I replace the annual unemployment rate with the unemployment rate at the interview month.

Despite the lack of cyclicalities in wage rates, hourly workers had very procyclical earnings (Table 2, column 3). The evidence is consistent with a response to the business cycle that involves adjustments in hours worked at relatively constant hourly wages.

Salaried Workers

In addition to the average hourly earnings variable, the PSID records an hourly wage rate at the interview date for salaried workers. As discussed earlier, the PSID coders divide the salary of the worker by some fixed number of hours so the cyclicalities of this variable reflect the cyclicalities of salaries. In the PSID, it is only possible to identify salaried workers from 1976 onward. Therefore, the sample period for the analysis is restricted to changes during the 1976–77 to 1990–91 period. The results are presented in Table 3.

The estimates in the first row are for the full sample of salaried stayers. The point estimates from both mean and median regression suggest that salaries were slightly procyclical (a one point increase in the unemployment rate reduces salaries by 0.3%). However, these estimates are very noisy and are not statistically significantly different from zero. The lack of evidence for strong procyclicality is also apparent from the lack of a clear relationship between the salary change series and the change in the unemployment rate series in Figure 1d. Earnings appear to be somewhat more procyclical than salaries and the estimates are statistically significant in the median regressions. This is as expected, because non-salary elements of earnings such as bonuses, commissions, and overtime are likely to be procyclical. Because hours are also procyclical, average hourly earnings are close to acyclical. Of course, the estimated cyclicalities of average hourly earnings may be biased by measurement error in hours. In the second row, I omit workers who had extra jobs but find that it has little effect on the point estimates.

In the third and fourth rows of Table 3, I take the sample of workers who had no

Table 3. Wage and Earnings Cyclicalities of Salaried Workers Who Did Not Change Employer, 1976–77 to 1990–91.^a
(Standard Errors in Parentheses)

<i>Sample (Sample Size)</i>	<i>Mean Regression</i>			<i>Median Regression</i>		
	<i>Hourly Wage</i>	<i>Earnings Divided by Hours</i>	<i>Earnings</i>	<i>Hourly Wage</i>	<i>Earnings Divided by Hours</i>	<i>Earnings</i>
All Job Stayers (12,260)	–0.29 (0.59)	–0.12 (0.38)	–0.51 (0.32)	–0.15 (0.51)	–0.25 (0.31)	–0.59** (0.23)
Stayers with No Extra Jobs (9,078)	–0.28 (0.67)	–0.08 (0.39)	–0.41 (0.30)	–0.14 (0.55)	–0.13 (0.30)	–0.66** (0.24)
Stayers with No Extra Jobs with Non-Salary Income (4,102)	–0.37 (0.70)	–0.76** (0.33)	–0.95** (0.33)	–0.24 (0.54)	–0.51 (0.37)	–1.12** (0.26)
Stayers with No Extra Jobs without Non-Salary Income (4,976)	–0.18 (0.72)	0.47 (0.47)	0.03 (0.34)	–0.13 (0.61)	0.09 (0.30)	–0.29 (0.27)

^aAll specifications are estimated using the two-step method described in the text. In all specifications there are 16 observations in the second-stage regression. In addition to the change in the unemployment rate, each specification includes a linear time trend, a cubic in labor market experience, and a cubic in tenure with the employer.

**Statistically significant at the 0.05 level.

extra jobs and subdivide them into one group that had no source of non-salary income in either year, and a second group that had access to non-salary income. I define workers as having access to non-salary income if they reported having income from bonuses, commissions, or overtime or if they reported that they got paid for overtime hours. The results are illuminating: while the cyclicalities of salaries varies little between these two groups, the non-salary income group has very procyclical earnings and the salary-only group has a small statistically insignificant cyclicalities coefficient. The point estimates suggest that a one point increase in the unemployment rate reduces earnings by about 1% for the first group, and has little effect on earnings for the second. Thus, it appears that salaried workers with non-salary earnings sources did have procyclical earnings. The evidence that these people had procyclical average hourly earnings is weaker, because while the mean regression estimate is statistically significant, the median regression estimate is not. In conclusion, the implication is that a major source of procyclicalities in the earnings of salaried workers was variation in incentive pay and

overtime over the business cycle rather than variation in salaries themselves.

My results for hourly and salaried workers are consistent with those of McLaughlin and Bils (2001), who provided estimates for the cyclicalities of the point-in-time wage measures (hourly wage for hourly workers, salary divided by predicted hours for salaried workers).¹⁷ They split the data between industry stayers and industry movers but did not break down the data by hourly/salaried. Their cyclicalities estimates for industry stayers (a group that includes both firm stayers and some firm movers) are slightly larger than my estimates when I pool all stayers.¹⁸ My investigation suggests that the estimates are similar because their use of a more procyclical wage measure (salary divided by predicted hours rather than salary) is offset by the fact that they

¹⁷They predict hours from a regression that includes individual fixed effects, industry effects, and year effects.

¹⁸I do not report results for pooled hourly and salaried workers in this paper because the wage measures are not strictly comparable for the two groups.

Table 4. Wage and Earnings Cyclicalities of Workers Who Were Not Salaried or Hourly and Did Not Change Employer, 1976–77 to 1990–91.^a
(Standard Errors in Parentheses)

<i>Sample (Sample Size)</i>	<i>Mean Regression</i>			<i>Median Regression</i>		
	<i>Hourly Wage</i>	<i>Earnings Divided by Hours</i>	<i>Earnings</i>	<i>Hourly Wage</i>	<i>Earnings Divided by Hours</i>	<i>Earnings</i>
All Job Stayers (1,539)	N/A	–2.17** (0.97)	–2.39** (1.06)	N/A	–3.11** (0.76)	–1.77** (0.62)
Stayers with No Extra Jobs (1,244)	N/A	–1.12 (0.96)	–2.15** (1.12)	N/A	–2.79** (0.81)	–1.68** (0.75)

^aAll specifications are estimated using the two-step method described in the text. In all specifications there are 16 observations in the second-stage regression. In addition to the change in the unemployment rate, each specification includes a linear time trend, a cubic in labor market experience, and a cubic in tenure with the employer.

**Statistically significant at the 0.05 level.

included women, whose wages are slightly less procyclical than men's.

Other Workers

The residual group is composed of persons who were neither paid hourly nor salaried. These are people who were paid piece rates, commissions, tips, and in other ways. One might expect such people to have had more cyclical wages because their compensation was tied directly to their productivity or to the demand for the product. Both productivity and product demand are likely to be procyclical. Once again, these workers can only be identified from 1976 onward, and so the sample is restricted to changes during the 1976–77 to 1990–91 period. For this group of workers there is no point-in-time wage measure; the only wage measure available is average hourly earnings.

The results in Table 4 indicate that average hourly earnings were much more procyclical for these workers than for hourly or salaried workers. This strong cyclicalities is obvious in Figure 1e, where the year dummy estimates from the median regression of average hourly earnings are plotted against the change in the unemployment rate. The median regression estimates suggest that a one point increase in the unemployment rate reduces average hourly earn-

ings by 3%; the mean regression estimates suggest somewhat less cyclicalities. However, the earnings procyclicalities of these workers is similar to that of hourly workers. The difference lies in the cyclicalities of hours: hours were less procyclical for this group of workers than they were for hourly workers. Of course, the cyclicalities of hours may be understated because of measurement error in hours.

These results suggest that a decline in the prevalence of traditional hourly and salaried methods of payment would lead to an increase in wage cyclicalities. For example, suppose the proportion of workers who were neither paid a salary nor paid by the hour increased four-fold from the 6% in the sample to 25%. The estimates from the median regressions in Tables 2–4 suggest that this would have approximately doubled the cyclicalities of average hourly earnings of stayers from –0.54 to –1.02.

Specification Checks

Phillips Curve

Equation (1) is the standard specification used in the wage cyclicalities literature and is similar to the specification used in wage curve estimation (Blanchflower and Oswald 1994). However, it is inconsistent with a Phillips curve specification that re-

Table 5. The Effects of Current and Lagged Unemployment Rates on Changes in Log Average Hourly Earnings, 1970–71 to 1990–91.^a
(Standard Errors in Parentheses)

Variable	(1) Full Sample (N: 68,463)	(2) Full Sample Less Self- Employed (N: 52,525)	(3) Job Stayers (N: 42,164)	(4) Job Stayers with No Extra Jobs (N: 31,957)
<i>Mean Regression</i>				
U_t^b	-1.28** (0.23)	-1.02** (0.24)	-0.74** (0.23)	-0.50** (0.20)
U_{t-1}	1.07** (0.23)	1.00** (0.23)	0.86** (0.22)	0.58** (0.19)
<i>Median Regression</i>				
U_t	-0.83** (0.17)	-0.67** (0.17)	-0.54** (0.18)	-0.46** (0.17)
U_{t-1}	0.81** (0.16)	0.73** (0.16)	0.67** (0.17)	0.60** (0.16)

^aAll specifications are estimated using the two-step method described in the text. In all specifications there are 22 observations in the second-stage regression. In addition to the current and lagged unemployment rate, each specification includes a linear time trend and a cubic in labor market experience. The specifications that include only stayers also contain a cubic in tenure with the employer.

^b U_t refers to the national civilian unemployment rate in year t .

**Statistically significant at the 0.05 level.

lates the rate of changes in wages to the *level* of the unemployment rate rather than the *change* in the unemployment rate. Card and Hyslop (1997) suggested a test between these two specifications. Rewrite equation (1) to allow the change in log wage to depend on the unemployment rate and the lag of the unemployment rate:

$$(4) \quad \Delta \ln w_{it} = \alpha_1 + \alpha_2 U_t + \alpha_3 U_{t-1} + \alpha_4 x_{it} + \alpha_5 \text{YEAR}_t + v_t + \varepsilon_{it}.$$

If equation (1) is the correct specification, then current and lagged unemployment rates have equal and opposite signs in equation (4). If the Phillips curve model is appropriate, lagged unemployment will have an insignificant effect on wage growth.

In Table 5, I present two-step results from the estimation of equation (4) using average hourly earnings as the dependent variable and the same samples as in Table 1. Both the OLS and Median regression results tell a consistent story for each of these samples: the coefficient on the lagged unemployment rate is approximately equal to minus the coefficient of the current unemployment rate. In all cases the lagged rate is significantly greater than zero. Thus, the specification in equation (1) appears more appropriate than the Phillips curve specification for these data.¹⁹

Skill Interactions

The results in Tables 2–5 suggest that wage and earnings cyclicity differs by mode of payment. However, as can be seen in Table A1, there are substantial differences in the observable characteristics of different types of stayers. Thus, an alternative explanation is that cyclicity differs by skill level, not by payment method. I investigate this possibility by dividing the sample of stayers into 5 quintiles based on the predicted values from a regression of log average hourly earnings on education, experience, tenure, and race controls. I then pool all stayers from 1976–77 to 1990–91 and allow the effects of the unemployment rate to differ by skill quintile as well as by payment type. As before, the estimation is carried out in two stages. Here the indicator variables used in the first step are Year * Payment Type * Skill Quintile. The coefficients on these 225 dummies are the dependent variable in the second stage.

Table 6 contains coefficient estimates for the variables of interest. Columns (1) and (3) report results where the controls include indicators for mode of payment, indicators for skill quintile, and interactions of mode of payment with the change in the unemployment rate. The estimates

¹⁹Results for hourly wages and for salaries are inconclusive because neither the current nor the lagged unemployment rate is significantly different from zero. This reflects the fact that wages and salaries exhibit no significant cyclicity.

Table 6. Wage Cyclicalities of Workers Who Did Not Change Employer, 1976–77 to 1990–91.^a
(Dependent Variable: Change in Log of Average
Hourly Earnings; Standard Errors in Parentheses)

<i>Description</i>	<i>Mean Regression</i>		<i>Median Regression</i>	
	(1)	(2)	(3)	(4)
Change in Unemployment Rate	–2.09** (0.75)	–2.10** (0.86)	–2.80** (0.56)	–2.51** (0.65)
Change in Unemployment Rate*Hourly	1.06 (0.78)	1.09 (0.79)	2.02** (0.59)	2.08** (0.59)
Change in Unemployment Rate*Salaried	1.91** (0.80)	1.98** (0.81)	2.57** (0.60)	2.49** (0.61)
Change in Unemployment Rate*1 st Quintile		–0.04 (0.62)		–0.51 (0.47)
Change in Unemployment Rate*2 nd Quintile		–0.28 (0.62)		–0.41 (0.46)
Change in Unemployment Rate*3 rd Quintile		–0.58 (0.60)		–0.13 (0.45)
Change in Unemployment Rate*4 th Quintile		–0.43 (0.59)		–0.36 (0.44)

^aAll specifications are estimated using the two-step method described in the text. There are 250 observations in the second-stage regression. The first stage includes a cubic in tenure and a cubic in experience, each interacted with mode of payment (hourly, salaried, or other). As well as the variables in the table, the second stage includes a linear time trend interacted with mode of payment, and indicator variables for mode of payment

reflect the earlier findings that hourly and salaried workers had much less cyclical average hourly earnings than other workers did. In columns (2) and (4), I add interactions of the change in the unemployment rate with the skill quintile indicators. These interactions are all statistically insignificant and the mode of payment interactions is not weakened. The conclusion is that it is method of payment that affects cyclicalities, not level of skill.

I have also carried out this exercise for the earnings of salaried workers. Table 3 showed that the earnings of salaried workers who had access to non-salary earnings were more procyclical than the earnings of other salaried workers. This large and statistically significant difference remains after I include interactions of the change in the unemployment rate with skill quintiles (results available on request).²⁰

State versus National Unemployment Rate

A possible explanation for the weak cyclicalities results for hourly and salaried workers is that shocks are so heterogeneous that the national business cycle is not a very useful measure of the shocks affecting individual firms. When I replace the national unemployment rate with the state unemployment rate, the estimated cyclicalities are generally lower with the state unemployment rate than with the national unemployment rate (results available on request). This suggests that my finding of little cyclicalities is not due to the national level measure being too aggregated. It may also reflect greater measurement error in the state level unemployment rates. Using the state unemployment rate allows one to in-

²⁰I have also tried to account for differences in characteristics across mode of payment by construct-

ing weights as in DiNardo, Fortin, and Lemieux (1996). I found the weighted estimates to be very close to the unweighted ones.

clude fixed year effects. Re-estimating the models including such effects reduces the estimated cyclicity of the point-in-time wage measures for hourly and salaried workers even further.²¹

Nominal Contracting

It is conceivable that real wages do not adjust to labor market conditions because there are nominal wage rigidities that influence the adjustment of real wages. Like McLaughlin (1994), I find that wages do not adjust fully to unexpected inflation. For both hourly and salaried workers, adding controls for the inflation rate to the wage change regression leaves the results essentially unchanged (details available on request). A natural interpretation is that employers do not adjust wages and salaries in line with changes in the unemployment rate. However, they partially adjust both wages and salaries to changes in the price level.

Recent research has examined whether nominal wage rigidity is primarily in a downward direction.²² When the inflation rate is low, downward nominal rigidity would inhibit real wage cuts; thus, one would expect the unemployment rate to have a larger effect on the wage in years when the inflation rate is high. When I allow for an interaction between the change in the unemployment rate and the inflation rate in the wage equations, I get coefficients on this interaction term that are negative, but they are never statistically significant (results available on request). Thus, I find no strong evidence that the effect of the unemployment rate on wages is greater during periods of higher inflation.

Conclusions

Like others in the literature, I find procyclical estimates for the average hourly earnings of job stayers. The contribution of this paper is to examine in detail wage cyclicity of different types of workers and of different sources of payment. For male hourly workers who did not change employer, there is little evidence that wage rates were procyclical. However, there is strong evidence of procyclical earnings movements for hourly workers. This suggests that, for hourly workers, the primary form of adjustment to the business cycle is in hours worked, and that changes in hours occur at fairly stable wage rates (Abowd and Card [1989] reached a similar conclusion).

The earnings of salaried workers who received income from bonuses, commissions, or overtime were procyclical. However, there is little evidence of earnings cyclicity for the majority of salaried workers who did not have access to these sources of income. Salaries themselves show no evidence of systematic variation over the business cycle. It appears that a major source of procyclicity in the earnings of salaried workers was variation in incentive pay and overtime over the business cycle rather than variation in salaries themselves.

The average hourly earnings of workers who were paid primarily on the basis of commissions, bonuses, or tips rather than being hourly or salaried appear to have been strongly procyclical. However, the number of these workers in the sample is small, so further research would be useful to determine the robustness of this result. Taken as a whole, the results of this paper suggest that incentive pay is much more sensitive to the business cycle than are wage rates or salaries. Survey results from the American Compensation Association suggest that 63% of 2,800 major companies used bonuses or other incentives to adjust the pay of some or all of their employees to performance in 1999. This compares with their estimate of 15% at the start of the decade. The results in this paper suggest that the increasing prevalence of such incentive pay in the U.S. economy will likely make the earnings of workers within matches more procyclical over time.

²¹When year effects are included, the state unemployment rate captures differences in the cycle across states. Thus, it is not surprising that the inclusion of year effects reduces the estimated cyclicity. I thank a referee for making this point.

²²The recent literature on downward nominal rigidities was initiated by McLaughlin (1994). Akerlof, Dickens, and Perry (1996), Card and Hyslop (1997), Kahn (1997), and Altonji and Devereux (2000), among others, have contributed to this literature.

Data Appendix

The full sample consists of 68,463 observations on wage changes and uses data from the years 1970–92. Included in this sample are all year to year changes in average hourly earnings in which the individual worked at least 100 hours in each year and was aged between 18 and 64.

The sample of All Stayers uses the same criteria plus the additional requirements that the individual could not be self-employed in either year, and had to be employed by the same employer in both years. Due to inconsistencies in the employer tenure data, there are many observations for which it is difficult to ascertain tenure. I have taken the conservative approach of omitting 7,865 such observations from the sample of Stayers.

The Hourly Sample is the subset of the Stayers

sample that satisfy two additional criteria: (1) the individual was paid hourly in both periods, and (2) the individual had non-missing values for the hourly wage and average hourly earnings in both periods.

The Salaried Sample is the subset of the Stayers sample that satisfy the following additional criteria: (1) the individual was paid a weekly, monthly, or annual salary in both years, and (2) the individual had non-missing values for salary and for average hourly earnings in both periods.

The Other Sample is the subset of the Stayers sample that satisfy the following additional criteria: (1) the individual was not paid a salary or an hourly wage in either year, and (2) the individual had non-missing values for average hourly earnings in both periods.

Table A1
Descriptive Statistics for the Samples

<i>Variable</i>	<i>All 1971–91</i>	<i>All Stayers, 1971–91</i>	<i>Hourly, 1971–91</i>	<i>Salary, 1977–91</i>	<i>Other, 1977–91</i>
Log of Real Avg. Hourly Earnings	2.3278	2.4579	2.3502	2.6612	2.4659
Change in Log Real Avg. Hourly Earnings	0.0198	0.0234	0.0202	0.0289	0.0129
Age	37.4687	38.1174	37.5370	38.9084	37.3086
Education	12.3519	12.3940	11.1637	14.2714	12.6590
Black	0.2768	0.2917	0.3798	0.1597	0.1741
Other Non-White	0.0226	0.0204	0.0206	0.0174	0.0279
Union Member	0.2590	0.3032	0.4660	0.1213	0.1533
Years of Experience	18.4306	19.1595	19.0722	19.2505	18.5343
Years of Tenure	8.3946	10.1065	9.7400	11.1757	7.6738
Year	81.5763	82.1832	81.9502	84.2883	84.6751
Nominal Earnings	20,995.23	23,363.09	19,165.53	32,479.72	29,815.66
Annual Hours	2,156.42	2,174.05	2,090.53	2,257.16	2,386.19
Observations	68,463	42,164	19,993	12,260	1,539

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