Wage differentials associated with flextime

Analysis of the Current Population Survey indicates positive wage differentials overall for women on flextime in 1989 and for both men and women in 1997; significant differentials emerge for selected motivations, industries, and occupations

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This article presents an empirical test of wage differentials associated with flextime, by gender, stated motivation for using flextime, industry, and major occupation. The test implicitly compares the relative strengths of two opposing effects: a negative compensating wage differential resulting from workers' preferences for flextime and a positive wage differential associated with higher productivity of workers on flextime attributed to what economists call the "efficiency wage hypothesis." Although previous studies have found evidence that flextime increases both productivity[1] and workers' satisfaction,[2] scant evidence has emerged thus far regarding the net quantitative or qualitative impact of these factors on equilibrium wages.

One exception is an article by Nancy Johnson and Keith Provan,[3] who applied a similar test to a much smaller data set and found flextime to be positively associated with wages for professional women, negatively associated with wages for nonprofessional women, and not significantly associated with wages for men. Johnson and Provan's sample totaled 258, obtained by survey from within a single State. The study reported in the current article, by contrast, uses nationwide samples of more than 5,000 workers, obtained from the U.S. Current Population Survey (CPS) supplement, "Multiple Job Holding, Flexitime, and Volunteer Work," for 1989 and 1997. In addition to estimating aggregate wage effects by gender in each year, the article estimates the flextime wage differential associated with specific reasons each worker reportedly preferred flextime in 1989. (Reasons for choosing flextime were not reported in 1997, preventing a comparison with that year.) Also estimated is the flextime wage differential associated with specific industries and specific major occupations for 1997. (Again, in 1989, the number of workers on flextime in particular occupations and industries was too small to draw a meaningful comparison with the later year.)

Results of the study indicate that flextime is associated with significantly higher wages overall. The size of the flextime wage differential for women is stable across the years 1989 and 1997 and is similar to the 1997 estimate for men. However, the 1989 flextime wage differential for men is much smaller than in 1997 and is not significantly different from zero. This finding suggests that the pattern of compensation has evolved in a similar direction for both male and female workers, but it evolved later for men.

The more detailed regressions for 1989 find that the only stated reason for desiring flextime associated with a significant wage differential among women is transportation. Among men, flextime taken for personal reasons is associated

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with a positive wage differential at the 0.01 level. Only a small number of industries exhibit significant flextime wage differentials for either men or women in 1997, and all of those differentials are positive. Two major industries (automotive and repair services; and social services, other professional services, and forestry and fisheries—grouped collectively as "professional" industries (see p. 5)—exhibit significant wage differentials for both men and women. Significantly positive flextime wage differentials emerge for men in all major occupations except operators, movers, and handlers, while women exhibit significantly positive flextime wage differentials only for sales and administrative occupations.

The article continues by presenting a brief overview of the history of flextime, describing the empirical and conceptual framework of the analysis, and characterizing the sample data. The article concludes with a discussion of the results and some suggestions for future research.

Background

Flextime is generally defined as a worker's ability to alter the starting and quitting time of a workday. It was introduced in Germany in 1967, spread quickly to other parts of Europe, and has been adopted by some U.S. employers during the past 20 years.[4] One of the first groups in the United States to experiment with a system of flexible working hours was the Federal Government's agencies. Over time, other firms have begun to adopt some form of flexible working hours as a means of attracting employees of higher quality or from a larger pool of applicants.

As of 1992, more than 13 percent of the U.S. workforce was covered by flextime arrangements, with a higher incidence among part-time than full-time workers.[5] Many of the firms offering flextime have found that it confers benefits on the employer, besides fostering employee morale. Flextime has been reported to reduce absenteeism and turnover, increase lines of communication, reduce stress in the workplace, and, in some cases, even increase productivity.[6] Increasing flexibility in the work schedule can reduce the uncertainty of conflicts between market work, nonmarket work, and leisure, as well as enabling workers to devote themselves more fully to their job responsibilities.

Still, not everyone embraces the flexible work schedule. Unions have opposed the idea of flexible work hours because it makes labor laws more difficult to enforce and may create an opportunity for firms to abuse the system. Also, some have argued that flextime is a hindrance to the effectiveness of the workplace because a worker must be present and visible in order to contribute fully to the job. Thus, empirical research into the net effects of flextime continues to be useful. With this in mind, the objective of the present study is to quantify whether, on average, employees find that flextime is associated with productivity gains that are not only positive, but also great enough to more than offset any compensating wage differentials that would be expected when workers prefer flextime to traditional work schedules. The analysis that follows is based on equilibrium wage theories.

It seems clear why women, at least, desire flextime benefits as they pursue careers and families. Even women who are employed full time spend 20 to 30 hours per week on housework; employed men spend at most half that time.[7] Traditionally, flexible schedule arrangements were sometimes offered to women who needed to take care of their children. Recently, however, because of a shortage of qualified labor, growing numbers of working mothers in the labor force, unacceptable levels of career progress for women, and work schedules for women that constrained their productivity, more employers have begun to offer family-related benefits. (Some of these changes in the roles of women and men are explored by Francine D. Blau and Marianne A. Ferber.[8])

Empirical framework and sample

Both the compensating wage differential theory and the efficiency wage hypothesis predict that wage rates are affected by pecuniary and nonpecuniary attributes. The compensating wage differential refers to a worker's willingness to pay (or forgo income) for desirable job attributes.[9] In contrast, according to the efficiency wage hypothesis, in a competitive labor market an employer will be forced to pay higher wages for more productive workers.[10] Thus, any given job attribute may have two types of effect on the overall wage: one reflecting the worker's direct preference for the attribute, the other reflecting any impact of the attribute on the worker's productivity (or, in this case, any possible selection of more productive workers into the attribute). In the case of flextime, the two effects may be intertwined to the extent that improved employee morale associated with a flexible work schedule may contribute to improved productivity through lower absenteeism, lower turnover, and greater effort expended on the job. Also, flextime may be able to contribute to higher productivity by reducing any interference from employees' outside obligations, and employers may selectively offer flextime only to their more productive workers.

It is the objective of this section to isolate and measure the impact of flextime on wages. To the extent that flextime is desired by workers, the compensating wage theory alone would predict a negative association between flextime and wages, controlling for a vector of other job attributes. If, however, flextime is associated with higher productivity among workers, the predicted impact on wages is slightly more complex. One might question why an employer should pay more for the added productivity of employees who are working in an improved environment. One answer would involve competition

among employers, as in conventional applications of the efficiency wage theory, plus an element of asymmetric information in that only the worker knows his or her personal (hedonic) value of flextime. As long as more than one employer offers flextime for a particular category of worker, employers may be forced to bid up their wages—possibly as high as the marginal value of the worker's product. Whether such a positive wage differential exists is an empirical question. If one is found, it would represent a *lower bound* on the value of actual differences in productivity, bearing in mind that some offsetting compensating wage differential may also be reflected in the observations.

The sample used in the analysis was collected from the CPS of May 1989 and May 1997.[11] The supplement titled "Multiple Job Holding, Flextime, and Volunteer Work" contains data on the usual number of hours worked daily and weekly, usual number of days worked weekly, specific days worked weekly, starting and ending times of an individual's workday, whether the starting and ending times could be varied, and—for 1989—the primary reasons each individual desired the flextime benefit in his or her workplace. The sample is drawn from all persons aged 18 to 65 in the civilian noninstitutional population of the United States living in households.

The 1989 sample size of full-time workers totaled 5,385 observations, of which 2,324 (43.2 percent) were women and 3,061 (56.8 percent) were men. The average hourly wage rate was $9.23: $10.35 for men and $7.74 for women.[12] The 1997 sample comprised 8,358 observations, including 3,800 women (45.5 percent) and 4,558 men (54.5 percent). A minimum hourly wage of $2.00 was imposed to reduce the impact of miscoded responses.[13] Table 1 presents descriptive statistics. Because of small samples in certain industries and occupations, several categories are grouped together: social services, other professional services, and forestry and fisheries are collectively denoted as "professional," and operators, movers, and handlers are collectively denoted as "operators." These groupings resulted in a minimum of 15 flextime observations, plus larger numbers of nonflextime observations, per industry or occupation in 1997, as needed to obtain statistically meaningful estimates in table 4. As shown in that table, of the 40 parameter cells (representing 20 industry or occupation categories times two genders), only 4 comprised fewer than 20 observations, while another 8 cells represented between 20 and 40 observations each. The 1989 data, representing a smaller sample and drawn from a period in which flextime was less common, contained fewer than 15 observations in each of 28 cells and between 15 and 17 observations in each of 6 more cells; those data were therefore not subjected to further decomposition. Smaller samples reported certain reasons for desiring flextime in 1989 (see table 3), but no natural groupings of those disparate reasons suggested themselves.

Besides observing the statistics in table 1, note that the 1989 mean wage rate was $8.97 for women on flextime, $7.66 for women not on flextime, $10.98 for men on flextime, and $10.31 for men not on flextime. These raw averages suggest an overall dominance of the efficiency wage hypothesis (reflecting higher productivity of flexing workers) over the compensating wage differential effect. The regressions that follow test this casual impression more formally.

The wage equation was estimated by gender, using the natural logarithm of wages as the dependent variable. Two versions were fitted, one with a simple FLEXTIME dummy variable, the other with a vector of FLEXREASONS described shortly:

$$\ln W_i = \alpha + X_{1i}\beta_1 + \beta_2 \text{FLEXTIME}_i + \varepsilon_i; \quad (1)$$

$$\ln W_i = \alpha + X_{1i}\beta_1 + \beta_2 \text{FLEXREASONS}_i + \varepsilon_i. \quad (2)$$

Here, X_{1i} is a vector of measurable characteristics that are expected to affect wages, such as potential work experience,[14] potential work experience squared, education, marital status, and race. These variables are commonly included in studies of compensating wage differentials.[15] Other included job characteristics that may affect earnings are union status, type of industry, occupation, and flextime. Nonpecuniary binary control variables include metropolitan area, the white race, and the southern geographic region. Also in X_{1i} is a vector of binary variables denoting each respondent's major occupation and major industry, as listed in table 1. Thus, the model that is being fit is a fixed-effects model that controls for both industry and occupation. To avoid a singularity in the presence of the intercept, the analysis omitted utilities as a major industry and farming as a major occupation. The stochastic error term is ε_i . Each equation was fitted by ordinary least squares.

In equation (1), FLEXTIME is a binary variable equal to unity for workers whose schedule allows them to vary the time they begin and end their workday, and equal to zero otherwise. In equation (2), FLEXREASON is a vector of binary variables indicating the primary reason workers on flextime reported for altering their schedules. The choices are as follows:

1. family and child responsibilities;
2. transportation;
3. helps to build up leave;
4. personal reasons;
5. enjoy flextime;
6. nature of the job.

Previous work by Johnson and Provan[16] yielded mixed results that failed to suggest any a priori hypothesis on the sign of FLEXTIME. However, one would expect that the average strength of workers' preferences for flextime might vary by reason, whereas the magnitude of any productivity effect of flextime might be relatively less sensitive to the reason. Thus, unequal coefficients across the reasons may primarily reflect unequal preferences, with the most preferred reasons possi-

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Table 1. Sample statistics

Variable	Women				Men			
	1989 *n* = 2,324		1997 *n* = 3,800		1989 *n* = 3,061		1997 *n* = 4,558	
	Mean	Standard deviation	Mean	Standard deviation	Mean	Standard deviation	Mean	Standard deviation
ln(wage)	1.965	0.405	2.410	0.489	2.245	0.433	2.606	0.500
Potential experience	19.930	11.213	19.140	9.481	19.856	11.202	19.092	9.196
Flextime	.071	.257	.229	.420	.0595	.2365	.2667	.4423
South	.360	.480	.334	.472	.322	.467	.303	.460
Metro	.715	.452	.786	.410	.711	.454	.799	.401
Married	.588	.492	.605	.489	.684	.465	.689	.463
Education	12.558	2.094	13.632	2.292	12.216	2.153	13.515	2.442
White	.832	.374	.835	.371	.865	.342	.882	.323
Unions	.171	.377	.159	.366	.364	.481	.203	.403
Major industry:								
Mining	.017	.128	.016	.124	.166	.372	.118	.323
Manufacturing	.219	.414	.145	.352	.346	.476	.240	.427
Transportation	.025	.155	.024	.152	.072	.258	.073	.260
Communication	.017	.130	.020	.139	.018	.133	.021	.142
Utilities	.007	.083	.010	.098	.033	.179	.031	.173
Wholesale	.190	.392	.156	.363	.145	.353	.185	.389
Finance	.068	.252	.107	.309	.019	.135	.049	.216
Hospital	.122	.328	.086	.281	.024	.152	.019	.137
Medical	.093	.290	.084	.278	.010	.100	.013	.114
Educational	.052	.221	.127	.333	.025	.157	.045	.206
Social	.025	.156	.032	.176	.005	.067	.008	.089
Professional	.026	.159	.042	.200	.012	.109	.044	.205
Forestry	.0009	.0293	.0011	.0324	.0010	.0313	.0011	.0331
Public administration	.053	.224	.061	.240	.050	.218	.061	.240
Major occupation:								
Managerial	.156	.363	.372	.483	.075	.263	.284	.451
Technical	.045	.208	.043	.204	.035	.184	.035	.184
Sales	.090	.287	.099	.299	.036	.185	.098	.298
Administration	.335	.472	.270	.444	.073	.259	.068	.251
Service	.191	.393	.128	.335	.414	.493	.306	.461
Operator	.141	.348	.061	.239	.165	.371	.086	.280
Movers	.0	.0	.008	.088	.07	.26	.073	.261
Handlers	.032	.177	.016	.127	.092	.289	.042	.201
Reason for desiring flextime:								
Family or child care	.009	.095	[1]	[1]	.0007	.0256	[1]	[1]
Transportation	.002	.046	[1]	[1]	.002	.048	[1]	[1]
Build up leave	.0004	.0207	[1]	[1]	.0003	.0181	[1]	[1]
Personal reasons	.004	.065	[1]	[1]	.002	.048	[1]	[1]
Enjoy flextime	.011	.103	[1]	[1]	.011	.106	[1]	[1]
Nature of the job	.034	.182	[1]	[1]	.038	.190	[1]	[1]

[1] 1997 survey did not report reasons for desiring flextime.

bly indicating a negative coefficient, as the negative compensating wage differential more than offsets any positive efficiency wage differential. However, if employers tend to be more willing to grant requests for flextime to workers who have proven to be more productive, then a positive efficiency wage component could emerge in these samples. In addition, when flextime is adopted because of the nature of the job, it could be that flextime is more the employer's choice than the employee's choice. This suggests a zero or negative compensating wage differential, perhaps a positive efficiency wage differential (particularly if the nature of the job requires flextime for productivity reasons), and thus a positive coefficient overall in equation (2).

Following previous studies, we anticipate positive coefficients on experience, education, metropolitan area, the white race, and union membership and negative coefficients on experience squared and the southern geographic region. We similarly expect the coefficient on married to be positive for men, but negative for women.

In addition, we estimate two other equations to quantify any systematic differences in the wage differentials associated with flextime by industry and by major occupation for 1997:

$$\ln W_i = \alpha + X_{1i}\beta_1 + \beta_2 \text{FLEX} \times \text{INDUSTRY}_i + \varepsilon_i; \quad (3)$$

$$\ln W_i = \alpha + X_{1i}\beta_1 + \beta_2 \text{FLEX} \times \text{OCCUPATION}_i + \varepsilon_i. \quad (4)$$

These decompositions will permit us to infer whether any apparent productivity effects of flextime may be relatively greater than the hedonic effects for certain industries or occupations. Although it is natural to suppose that productivity effects may be unequal across the various industry or occupation

categories, we did not hypothesize specific effects a priori.

Results

Table 2 presents the regression results for wage equation (1) by gender. The results for 1989 indicate that flextime is associated with higher wages for women ($t = 2.53$, significant at the 0.05 level), as in Johnson and Provan's subsample of professional women.[17] This outcome is consistent with an efficiency wage effect—reflecting higher productivity—dominating any compensating wage differential. For men, no significant wage differential is associated with flextime ($t = 0.48$), suggesting that any positive efficiency wage effect is roughly offset by a negative compensating wage differential (and conversely). For 1997, flextime is associated with significantly higher wages for both men and women at the 0.01 level; the magnitude of the "flextime premium" for women is virtually unchanged from its 1989 value, while that for men is nearly the same as for women.

The majority of other control variables exhibit significant coefficients, except for occupation effects on women. Experience shows positive, but declining, marginal returns, and wages are higher in metropolitan areas, but lower in the south. Education, unionization, and being a member of the white race are all associated with higher wages, as in previous studies.

Table 3 presents the regression results for wage equation (2), distinguishing the various reasons for flextime in 1989. For each gender, only one flextime reason is associated with a significant wage differential: transportation for women and personal reasons for men, each with a positive coefficient. For the other reasons for adopting flextime, a coefficient not significantly different from zero could be consistent with a net offset of positive and negative wage differentials from productivity and compensating wage effects. However, as noted earlier, a sparse representation for some of these reasons (especially among men) makes it difficult to detect significance in

Table 2. Parameter estimates, wage equation (1)

Variable	Women				Men			
	1989		1997		1989		1997	
	Coefficient	*t*-statistic	Coefficient	*t*-statistic	Coefficient	*t*-statistic	Coefficient	*t*-statistic
Intercept	1.129	[1]5.44	0.928	[1]5.13	0.874	[1]10.57	0.745	[1]8.16
Experience	.013	[1]5.58	.024	[1]9.06	.022	[1]9.20	.026	[1]9.59
Experience squared	–.00020	[1]–4.01	–.00044	[1]–6.56	–.0003	[1]–7.05	–.00041	[1]–6.17
Education	.041	[1]10.63	.0728	[1]20.80	.040	[1]11.65	.068	[1]21.51
South	–.053	[1]–3.75	–.047	[1]–3.50	–.087	[1]–6.21	–.047	[1]–3.60
Metro	.102	[1]6.82	.137	[1]8.87	.097	[1]6.75	.119	[1]7.92
Married	–.003	–.26	.013	.99	.060	[1]4.14	.093	[1]6.91
White	.046	[2]2.49	.032	[3]1.82	.106	[1]5.53	.122	[1]6.54
Union	.230	[1]12.22	.143	[1]7.50	.235	[1]15.97	.137	[1]8.54
Flextime	.066	[2]2.53	.067	[1]4.41	.013	.48	.062	[1]4.41
Major industry:								
Mining	.231	[1]4.16	.114	[2]2.12	.347	[1]12.14	.210	[1]7.91
Manufacturing	.242	[1]8.26	.111	[1]3.92	.230	[1]8.58	.149	[1]6.28
Transportation	.266	[1]5.43	.143	[1]3.10	.242	[1]6.95	.128	[1]4.06
Communication	.243	[1]4.35	.154	[1]31.4	.289	[1]5.45	.210	[1]4.60
Utilities	.233	[1]2.83	.198	[1]2.99	.321	[1]7.59	.274	[1]6.98
Wholesale	.002	.07	–.133	[1]–4.86	.059	[2]1.99	–.060	[2]–2.41
Finance	.133	[1]3.87	.046	1.61	.052	.99	.071	[2]2.14
Hospital	.291	[1]9.64	.094	[1]3.13	.060	1.26	–.048	–1.01
Medical	.124	[1]3.93	.022	.74	–.015	–.23	–.052	–.94
Educational	.037	.98	–.095	[1]–3.16	.016	.34	–.126	[1]–3.55
Social	–.157	[1]–3.31	–.161	[1]–3.97	–.062	–.64	–.150	[2]–2.17
Professional	.115	[2]2.44	.042	1.12	.217	[1]3.43	.099	[1]2.83
Forestry	.044	.18	–.373	[3]–1.94	–.314	–1.54	.254	1.40
Public administration	.198	[1]5.33	.076	[2]2.29	.257	[1]6.96	.131	[1]4.15
Occupation:								
Managerial	.118	.59	.214	1.25	.345	[1]5.04	.417	[1]5.44
Technical	.085	.42	.105	.60	.293	[1]4.05	.378	[1]4.62
Sales	–.150	–.75	.079	.46	.067	.91	.300	[1]3.82
Administrative	–.069	–.35	–.035	–.21	.134	[2]1.98	.176	[2]2.23
Service	–.264	–1.32	–.180	–1.05	.189	[1]2.96	.171	[2]2.25
Operator	–.230	–1.14	–.170	–.98	.148	[2]2.25	.059	.76
Movers	–.093	–.43	–.004	–.02	.103	1.54	.126	1.60
Handlers	–.232	–1.14	–.181	–1.02	.021	.32	.051	.63
Observations	2,324		3,800		3,061		4,558	
Adjusted R^2	.40		.40		.35		.37	

[1]Significant at 0.01 level.
[2]Significant at 0.05 level.
[3]Significant at 0.10 level (two-tailed tests).

a number of these cases. None of the reasons exhibit a significantly negative coefficient, suggesting that no reason is valued strongly enough by workers to more than offset any positive productivity effect.

Table 4 reports regressions for 1997, incorporating interactive variables between flextime and major industry (wage equation (3)) and between flextime and major occupation (wage equation (4)). In equation (3), for women, two interactive terms (automotive and repair, and hospital) are significant at the 0.05 level, while two more (communication and professional) are significant at the 0.10 level in a two-tailed test. These findings are consistent with several possible interpretations, which the analysis presented here cannot distinguish. First, flextime may be associated with an exceptionally large improvement in productivity among women in the four industries mentioned. Second, employers in those industries may selectively grant requests for flextime (or perhaps even impose flextime) on their more productive female employees. Third, the association between productivity and flextime—whatever the causality—may be positive across all industries, but women who choose to work in manufacturing may not value flexible work schedules to the same extent as women who work in other industries.

Table 3. Parameter estimates, wage equation (2), 1989

Variable	Women		Men	
	Coefficient	*t*-statistic	Coefficient	*t*-statistic
Intercept	1.14	[1]5.48	0.87	[1]10.54
Experience	.01	[1]5.50	.02	[1]9.24
Experience squared	–0.0	[1]–3.94	–.00	[1]–7.08
Education	.04	[1]10.61	.04	[1]11.62
South	–.05	[1]–3.74	–.09	[1]–6.26
Metro	.10	[1]6.88	.10	[1]6.66
Married	–.00	–.26	.06	[1]4.12
White	.05	[2]2.49	.11	[1]5.59
Unions	.23	[1]12.27	.24	[1]16.02
Reason for desiring flextime:				
Family or child care	.07	1.18	.01	.08
Transportation	.45	[1]3.17	.04	.28
Build up leave	–.31	–1.00	–.02	–.05
Personal reasons	.12	1.21	.37	[1]2.79
Enjoy flextime	.05	.84	.06	.97
Nature of the job	.04	1.05	–.03	–.79
Major industry:				
Mining	.22	[1]3.96	.35	[1]12.27
Manufacturing	.24	[1]8.08	.23	[1]8.72
Transportation	.26	[1]5.38	.25	[1]7.11
Communication	.24	[1]4.32	.29	[1]5.54
Utilities	.23	[1]2.79	.33	[1]7.69
Wholesale	.00	.01	.06	[2]2.17
Finance	.13	[1]3.83	.06	1.11
Hospital	.29	[1]9.63	.06	1.25
Medical	.12	[1]3.88	–.01	–.17
Educational	.04	.96	.02	.45
Social	–.16	[1]–3.31	–.05	–.48
Professional	.12	[2]2.46	.22	[1]3.46
Forestry	.06	.23	–.33	–1.59
Public administration	.20	[1]5.32	.26	[1]7.10
Major occupation:				
Managerial	.11	.56	.34	[1]5.02
Technical	.08	.41	.29	[1]4.03
Sales	–.15	–.76	.07	.91
Administrative	–.07	–.36	.13	1.94
Service	–.27	–1.34	.19	[1]2.95
Operator	–.23	–1.15	.15	[2]2.20
Movers	–.10	–.45	.10	1.51
Handlers	–.23	–1.15	.02	.31
Adjusted R^2	.41		.36	

[1]Significant at 0.01 level.
[2]Significant at 0.05 level.
[3]Significant at 0.10 level (two-tailed tests).

In equation (4), for women, two interactive variables are highly significant and positive: flextime × sales, with $t = 4.17$, and flextime × administrative, with $t = 3.51$. Each of these is significant at the 0.01 level. The positive sign of both coefficients suggests either a stronger positive productivity effect of flextime in those occupations (again, whichever way the causality runs) or a systematically weaker personal preference for flextime in those occupations, combined with a positive productivity effect.

For the sample of men, equation (3) exhibits significantly positive coefficients for four major industries. As with women, flextime ×automotive and repair and flextime × professional exhibit positive coefficients, with $t = 4.31$ and 1.67, respectively. In contrast to the sample of women, however, flextime × manufacturing and flextime × medical are significant, with $t = 1.84$ and 2.30, respectively. These coefficients are consistent with a stronger association between flextime and productivity or with weaker preferences for flextime in those four industries. For men, equation (4) exhibits positive coefficients that are significant for all major occupations except operators.

From equations (3) and (4), the emergence of distinct gender-based marginal wage effects of flextime across some industries and occupations raises questions that could usefully be addressed in future studies. Are the differences due primarily to differences in productivity or in hedonic preferences? Can such findings identify those industries or occupations which could benefit more than others from a more widespread adoption of flextime? Do the differences reflect systematic discrimination by gender, or do they instead point to additional factors that must be controlled for in studies aimed at measuring wage discrimination? To what extent do any positive productivity effects that are observed result from flextime itself, as opposed to reflecting an employer's selective offering of flextime to a more productive subset of workers?

FLEXTIME IS AN EMERGING TREND IN THE MODERN WORKPLACE, with potential benefits for employers as well as employees. Theoretically, the net impact of flextime on wages depends on the relative strengths of two opposing effects and therefore raises the important empirical question of which effect is stronger either in general or in a given case. The CPS supplements from 1989 and 1997 offer a rich data set that may be used to answer that question.

Table 4. Parameter estimates for interactive flextime terms, 1997

Variable	Women			Men		
	Coefficient	*t*-statistic	Number of observations[1]	Coefficient	*t*-statistic	Number of observations[1]
Equation (3)						
Flextime × industry:						
Flextime × mining	0.0843	0.74	15	0.0456	1.05	103
Flextime × manufacturing	.0204	.46	97	.0545	[4]1.84	252
Flextime × transportation	.0857	.93	23	.0671	1.26	71
Flextime × communication	.1798	[4]1.84	21	.1304	1.42	26
Flextime × wholesale	.0348	1.00	165	.0191	.63	247
Flextime × finance	.1558	[2]3.65	108	.0668	1.21	81
Flextime × automotive and repair	.1410	[3]2.18	45	.2070	[2]4.31	108
Flextime × services	.0385	.40	20	.0168	.14	16
Flextime × entertainment	–.1055	–.89	16	–.0310	–.28	18
Flextime × hospital	.1051	[3]1.97	63	.1272	1.26	20
Flextime × medical	.0278	.52	63	.2505	[3]2.30	20
Flextime × educational	–.000278	–.00	46	–.0187	–.27	43
Flextime × professional	.0777	[4]1.65	104	.0866	[4]1.67	101
Flextime × public administration	.0412	.77	74	.0377	.73	86
Equation (4)						
Flextime × occupation:						
Flextime × managerial	.0284	1.24	415	.0576	[2]2.55	547
Flextime × technical	–.0608	–.84	36	.1542	[2]2.38	61
Flextime × sales	.1773	[2]4.17	116	.0843	[3]2.21	182
Flextime × administrative	.1050	[2]3.51	203	.0900	[4]1.62	66
Flextime × service	.0637	1.37	81	.0657	[3]2.25	223
Flextime × operators	.0368	.42	20	–.0057	–.15	130

[1]The number of observations is the number of flexing employees in each industry or profession.
[2]Significant at 0.01 level.
[3]Significant at 0.05 level.
[4]Significant at 0.10 level (two-tailed tests).

This article has found evidence of a positive wage differential associated with flextime for a sample of 2,324 women in 1989 and 3,800 in 1997, presumably reflecting a positive productivity effect that more than offsets any compensating wage differential reflecting hedonic preferences for flextime. No significant wage differential accompanied the adoption of flextime for the 1989 sample of more than 3,000 men, a finding that is consistent with the hypothesis that any productivity effects are approximately offset by hedonic effects within that sample. These results are all generally consistent with earlier findings obtained by Johnson and Provan for a much smaller and more locally limited sample, with the exception of their results for nonprofessional women.[18] However, the 1997 sample of more than 4,500 men exhibited a significantly positive wage differential associated with flextime, consistent with the findings from the sample of women.

Decomposing the 1989 observations by stated reason for adopting flextime, the analysis presented finds that only a single reason was associated with measurable wage effects for each gender: transportation for women on flextime and personal reasons for men on flextime. Both of those reasons exhibited positive wage differentials, suggesting productivity benefits of flextime in those cases. This issue has apparently not been previously studied, and the omission of reasons for flextime from the 1997 survey prevented its further exploration.

Decomposing the observations by industry and by occupation for 1997 reveals positive wage differentials for women in communication, finance, automotive and repair, hospitals, and professional services and for men in manufacturing, automotive and repair, medical services, and professional services. Positive wage differentials were associated with women on flextime in sales and administrative occupations and with men on flextime in managerial, technical, and service occupations. Again, these decompositions appear never to have been addressed in the literature. The differences found across industries and occupations by gender may warrant further research to determine whether they are specific to the samples used or more systematic.

Further research on the incidence and causes of a positive flextime wage differential appears warranted. Some may find the efficiency wage hypothesis an unconvincing explanation in this context, despite more direct evidence that flextime may enhance productivity.[19] As discussed earlier, one variant of this idea is that some employers may allow only their most productive and reliable employees the option of flextime, using it as a nonpecuniary form of compensation that complements pecuniary compensation, or possibly relying on the personal integrity of their best workers to mitigate a greater

difficulty involved in monitoring the effort contributed by employees on flextime. An alternative, more cynical, explanation is that employers who offer flextime are, on average, simply less serious about maximizing profits and may also pay above-market wages as another dimension of corporate inefficiency. If data on employers as well as employees were available, this hypothesis could be tested by comparing the overall cost efficiency, profit efficiency, or other kind of efficiency of employers who allow their employees to use flextime, as opposed to those who do not.

Another question revolves around the stated reasons for adopting flextime: might these reasons mask a pattern of strategic misreporting as workers seek to conform to entrenched organizational and cultural norms or to avoid signaling that they place a large hedonic value on flextime? For instance, other things being equal, are women on flextime paid more if their stated motivation is transportation rather than family and child responsibilities? Are fathers on flextime paid more if their stated motivation is unspecified personal reasons rather than family and child responsibilities? The empirical results reported in this article are consistent with these hypotheses and others, but are merely suggestive, given the data currently available. □

Notes

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[1] See John D. Owen, "Flexitime: Some Problems and Solutions," *Industrial and Labor Relations Review*, January 1977, pp. 152–160; Steven G. Allen, "An Empirical Model of Work Attendance," *Review of Economics and Statistics*, February 1980, pp. 77–87; D. R. Dalton and D. Mesch, "The Impact of Flexible Scheduling on Employee Attendance and Turnover," *Administrative Science Quarterly*, June 1990, pp. 370–87; and Edward M. Shepard, Thomas J. Clifton, and Douglas Kruse, "Flexible Work Hours and Productivity: Some Evidence from the Pharmaceutical Industry," *Industrial Relations*, January 1996, pp. 123–39.

[2] Marni Ezra and Melissa Deckman, "Balancing Work and Family Responsibilities: Flextime and Child Care in the Federal Government," *Public Administration Review,* March–April 1996, pp. 174–79.

[3] Nancy Johnson and Keith Provan, "The Relationship between Work/Family Benefit and Earnings: A Test of Competing Predictions," *Journal of Socio-Economics*, Winter 1995, pp. 571–84.

[4] See Owen, "Flexitime"; and Barney Olmstead, "Flexible Work Arrangements: From Accommodation to Strategy," *Employment Relations Today*, summer 1995, pp. 11–20.

[5] David Lewin and Daniel J. Mitchell, *Human Resource Management: An Economic Approach,* 2d ed. (Cincinnati, South-Western College Publishing, 1995), see especially p. 155.

[6] See Owen, "Flexitime"; Allen, "Model of Work Attendance"; Dalton and Mesch, "Impact of Flexible Scheduling"; and Shepard, Clifton, and Kruse, "Flexible Work Hours and Productivity."

[7] Joni Hersch and Leslie Stratton, "Housework, Fixed Effects, and Wages of Married Workers," *Journal of Human Resources*, spring 1997, pp. 285–307.

[8] Francine D. Blau and Marianne A. Ferber, *The Economics of Women, Men, and Work*, 2d ed. (Englewood Cliffs, NJ, Prentice Hall, 1992).

[9] Sherwin Rosen, "Hedonic Prices and Implicit Markets: Product Differentiation in Pure Competition," *Journal of Political Economy*, January–February 1974, pp. 34–55.

[10] See Harvey Leibenstein, *Economic Backwardation and Economic Growth* (New York, John Wiley and Sons, 1957); and Joseph E. Stiglitz, "The Efficiency Wage Hypothesis, Surplus Labour, and the Distribution of Income in L.D.C.s," *Oxford Economic Papers*, July 1976, pp. 185–207.

[11] The CPS is conducted by the U.S. Bureau of the Census for the Bureau of Labor Statistics.

[12] Full-time salaried workers whose usual weekly hours are not less than 35 were included in the sample with an imputed hourly wage rate equal to the ratio of weekly earnings to usual hours. Note that, because CPS wage data are top coded, average hourly wage data will be biased downward.

[13] At the time of the survey, the Federal minimum wage was $3.35 per hour. However, some States had minimum wage rates that were lower than the Federal minimum, and some jobs did not fit the Federal definition of interstate commerce and so were exempt from the minimum. The value of $2.00 was chosen to correspond to known wage rates of certain jobs (for example, waitress) at the time of the survey. Observations reporting a wage rate lower than $2.00 per hour were treated as miscoded responses and were ignored.

[14] Potential work experience is defined as age, minus education, minus 6 years and is usually a larger number than actual work experience.

[15] See Rosen, "Hedonic Prices and Implicit Markets"; and Charles Brown, "Equalizing Differences in the Labor Market," *Quarterly Journal of Economics*, February 1980, pp. 113–34.

[16] Johnson and Provan, "Work/Family Benefit and Earnings."

[17] *Ibid.*

[18] *Ibid.*

[19] See Owen, "Flexitime"; Allen, "Model of Work Attendance"; Dalton and Mesch, "Impact of Flexible Scheduling"; and Shepard, Clifton, and Kruse, "Flexible Work Hours and Productivity."