EMPLOYMENT POLICIES INSTITUTE

The Relative Compensation of Part-Time and Full-Time Workers

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Barry Hirsch of Trinity University

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Dr. Hirsch's research has appeared in the nation's most respected economics and industrial relations journals, including the *Journal of Labor Economics, Journal of Human Resources, Industrial and Labor Relations Review,* the *Review of Economics and Statistics* and the *American Economic Review.* He is author of *Labor Unions and the Economic Performance of U.S. Firms* and is co-author of the *Union Membership and Earnings Data Book*: *Compilations from the Current Population Survey*. He received his Ph.D. from the University of Virginia in 1977.

Recent Publications

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Rising Above the Minimum Wage, by William Even, Miami University of Ohio, and David Macpherson, Florida State University, January 2000.

Economic Analysis of a Living Wage Ordinance, by George Tolley, University of Chicago, Peter Bernstein, DePaul University, and Michael Lesage, RCF Economic & Financial Consulting, July 1999.

The Employment Impact of a Comprehensive Living Wage Law: Evidence from California, Employment Policies Institute, July 1999.

Effective Marginal Tax Rates on Low-Income Households, by Daniel N. Shaviro, New York University School of Law, February 1999.

An Analysis of the Baltimore Living Wage Study, Employment Policies Institute, October 1998.

Targeted Jobs Tax Credits and Labor Market Experience, by Frederick J. Tannery, University of Pittsburgh, June 1998.

Job Loss in a Booming Economy, 2nd Edition, Employment Policies Institute, May 1998.

Effects of the 1998-1999 Oregon Minimum Wage Increase, by David A. Macpherson, Florida State University, May 1998.

Effects of the 1999-2000 Proposed Washington Minimum Wage Increase, by David A. Macpherson, Florida State University, May 1998.

Effects of the 1998 California Minimum Wage Increase, by David A. Macpherson, Florida State University, March 1998.

Work Ethic and Family Background, by Casey B. Mulligan, University of Chicago, May 1997.

The Minimum Wage Debate: Questions and Answers, Third Edition, Employment Policies Institute, May 1997.

From Welfare to Work: The Transition of an Illiterate Population, Employment Policies Institute, January 1997.

Who Are The “Low-Wage” Workers? by Derek Neal, University of Chicago, July 1996.

The 1992 New Jersey Minimum Wage Increase: How Much Did it Affect Family Income?, by David A. Macpherson, Florida State University, May 1996.

The Impact of the Federal Unemployment Insurance Tax Ceiling, by Daniel S. Hamermesh and David Scoones, University of Texas at Austin, October 1995.

Effects of High School Work Experience a Decade Later: Evidence From The National Longitudinal Survey, by James D. Wright and Rhoda Carr, Tulane University, September 1995.

The Effects of Minimum Wages on Teenage Employment, Enrollment and Idleness, by David Neumark, Michigan State University, August 1995.

Youth Employment in the Hospitality Sector, by Bradley Schiller, School of Public Affairs, American University, Washington, D.C., June 1995.

The Consequences of Indexing the Minimum Wage to Average Wages in the U.S. Economy, by David Macpherson, Florida State University, and William Even, Miami University, May 1995.

Earnings Growth and Employment Stability of Workforce Entrants: Does the Point of Entry Matter? by Frederick J. Tannery, University of Pittsburgh and Slippery Rock University, April 1995.

Jobs Taken by Mothers Moving from Welfare to Work: And the Effects of Minimum Wages on this Transition, by Peter D. Brandon, Institute for Research on Poverty, University of Wisconsin — Madison, February 1995.

Minimum Wage Laws and the Distribution of Employment, by Kevin Lang, Boston University, January 1995.

The Effects of High School Work Experience on Future Economic Attainment, by Christopher J. Ruhm, University of North Carolina at Greensboro, May 1994.

The Early Careers of Non-College-Bound Men, by Jeff Grogger, University of California, Santa Barbara, May 1994.

Public Policies for the Working Poor: The Earned Income Tax Credit vs. Minimum Wage Legislation, by Richard V. Burkhauser, Syracuse University, and Andrew J. Glenn, Vanderbilt University, March 1994.

The Relative Compensation of Part-Time and Full-Time Workers: The Role of Worker and Job Skills

Executive Summary

Since the UPS strike in 1997, organized labor and some policy makers have been highly critical of part-time employment. At the root of these criticisms is the perceived wage gap experienced by part-time employees. At first glance it appears that part-timers earn substantially less than their full-time counterparts. Dr. Barry Hirsch shows in this study that the apparent wage gap is not nearly as large as some have claimed. In fact, once he accounts for all relevant factors, he finds the real wage gap to be either nonexistent or quite small.

The bulk of the wage differential between part-time and full-time employees is attributed to lower skill levels among part-time workers, and to the work preferences of those who choose part-time employment. Dr. Hirsch reaches much further than past examinations of part-time work by incorporating typically unmeasured worker-specific skill measures, plus measured job skill requirements and working conditions.

The major contribution of this study is this: it effectively dispels the notion that part-time employment carries any sizable *wage* penalty for similarly skilled employees (there does exist a gap in nonwage benefits). Dr. Hirsch shows that the part-time wage gap is particularly narrow among women, who compose 68% of the part-time work force.

Explaining Wage Differences

Many researchers have measured (and some have attempted to explain) the part-time "wage gap." Overall, women in part-time jobs typically earn 25.9% less than women in full-time jobs, while the comparable figure is 46.2% for men. These raw figures, however, are extremely misleading.

The first part of the Hirsch analysis concentrates on the differences between part-time and full-time workers in general. These differences are dissected using a range of important variables, including occupation, gender, years of schooling, marital status, region, industry of employment, job tenure and specific measurements of hours worked. The study also taps data on occupational skills requirements (verbal, mathematical, spatial, problem solving, etc.) and working conditions.

Dr. Hirsch draws on Census Bureau data and new Department of Labor data to show that part-time employment is systematically concentrated among jobs requiring a lower skill level compared to full-time work for both women and men. As Dr. Hirsch puts it, "Full-time work-

"There exists little wage gap between similarly skilled part-time and full-time workers."

ers are employed in occupations requiring higher levels of verbal, mathematical, problem solving, technical and system skills." These higher skills ultimately translate into higher wages for full-time employees, creating the illusion of a wage gap that is, in fact, largely explained by skill levels.

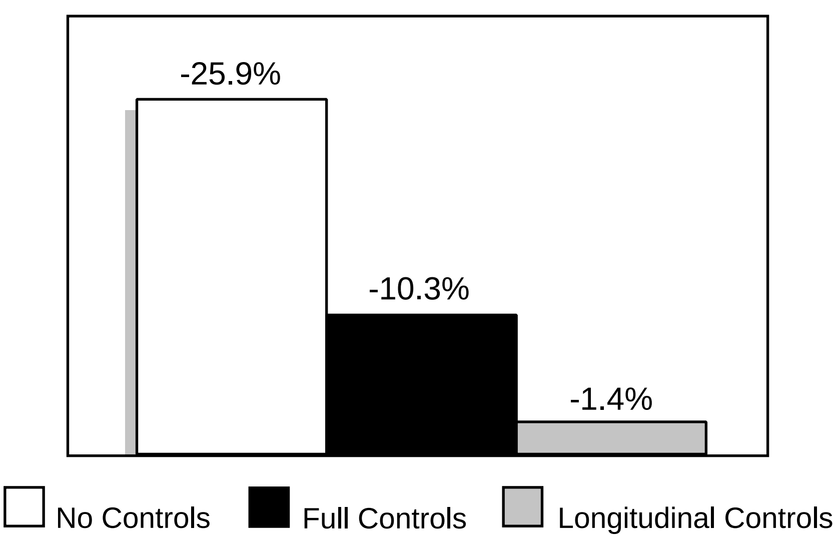
This research finds that measurable personal and location characteristics, in particular age or experience, can readily account for a large portion of the part-time wage differential. For women, 33% of the wage differential is explained by these characteristics, while the comparable figure for men is 48%. When personal, location and job characteristics are all considered, Dr. Hirsch accounts for 64% of the wage differential for women and 68% for men.

Evidence from Workers Switching Status

After taking into account measurable personal, location and occupational characteristics, the part-time wage gap is reduced to 10.3% for women and 18.1% for men. Thus, approximately two-thirds of the traditional estimate of the part-time wage gap is accounted for by these measurable characteristics. Dr. Hirsch identifies the factors explaining the remaining wage gap by accounting for characteristics that are not typically measured in economic databases.

Dr. Hirsch examines wage gains and/or losses of individual workers switching between full-time and part-time status from one year to the next. Using longitudinal estimates, he filters out skill differences between workers and is able to focus more closely on the potential wage gap itself. "There exists little wage gap between similarly skilled part-time and full-time workers."

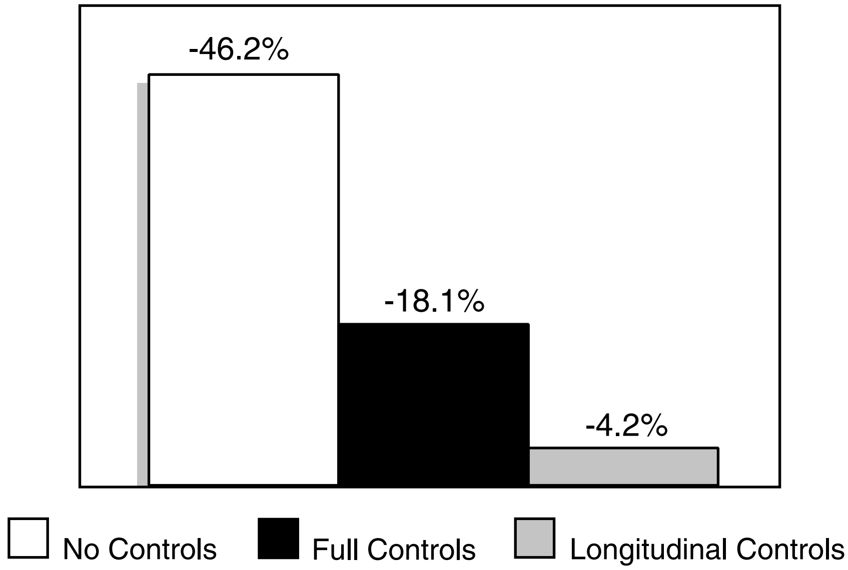
Female Part-Time Wage Gap: No Controls, Full Controls, Longitudinal Controls



On average, the wage "penalty" for women moving between part-time and full-time work is 1.4%, a fraction of the traditional estimates of the part-time wage gap. Men, on the other hand, experience a modest 4.2% penalty for being employed in part-time work compared to full-time work. These numbers show that similarly skilled and experienced individuals are subject to very little wage change as they move between part-time and full-time employment.

The only area in which Dr. Hirsch is able to find a nontrivial wage penalty is among workers who switch both industry and occupation in moving between part-time and full-time work. This effect is likely due to younger employees switching out of noncareer part-time jobs and into career-oriented full-time jobs, or vice versa. Even so, "the wage disadvantage is considerably smaller than that suggested by standard wage level analysis," according to this study. No part-time wage penalty is found among those switching to a job in the same occupation or industry.

Male Part-Time Wage Gap: No Controls, Full Controls, Longitudinal Controls



Evidence from Displaced Workers

Dr. Hirsch strengthens his analysis by examining one situation in which part-time work could play a substantial role in an individual's career — during periods of displacement (layoff, termination, etc.). Evidence from the Census Bureau Displaced Workers Surveys, in agreement with the longitudinal results described above, indicates that there is no part-time wage penalty for women and a very small penalty for men.

Experience, Skills and Tenure

The study shows that some of the part-time wage penalty can be attributed to lower human capital formation among part-time employees. Employees build skills and experience by spending time on the job. Part-timers by definition spend less time on the job than full-timers. Over a period of years, this results in significantly lower skill levels among part-time workers relative to comparable full-timers. These lower skill levels translate into lower wage levels as well.

Dr. Hirsch also points out that part-timers have less firm-specific experience because their tenure with their current employer tends to be shorter. On average, women working part-time have been with their current employer 4.5 years, as compared to 7.5 years for women working full-time. Men in part-time jobs have spent an average of 3.2 years in their current job, compared to 8.8 years for those in full-time employment. It is not surprising that employers are generally less willing to offer higher wages to part-time employees who have been in their current job for a shorter amount of time, and have worked fewer hours while there, compared to full-time employees.

Conclusion

It is important to note that part-time work is most often a voluntary arrangement. For family and personal reasons, an overwhelming majority of part-time employees prefer the flexibility of part-time employment. In 1997, only about two out of every ten part-time employees would have preferred full-time work to the part-time job they held. Dr. Hirsch's findings on the wage gap (or lack thereof) apply to all part-time workers, whether voluntary or not.

Based on the evidence, Dr. Hirsch concludes, "...virtually all of the part-time wage disadvantage can be accounted for by what are lower worker-specific skills among part-time than full-time workers…." To the extent that a part-time wage gap exists, it is largely a skills gap. Effective public policies must recognize the role of skills in determining wage levels. These findings carry tremendous value for policy makers who will no doubt be called upon to "solve" the perceived problems surrounding part-time employment.

Thomas K. Dilworth
Research Director

The Relative Compensation of Part-Time and Full-Time Workers

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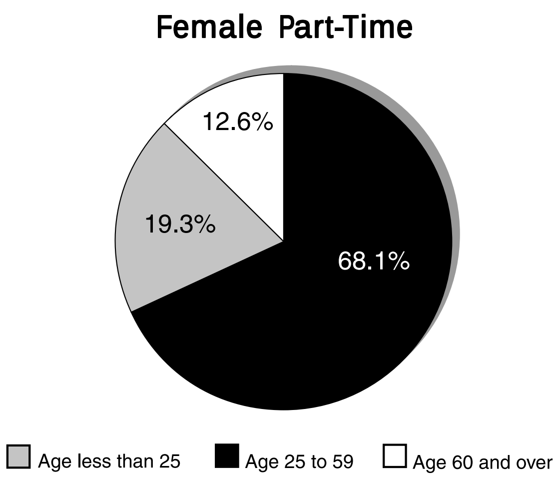
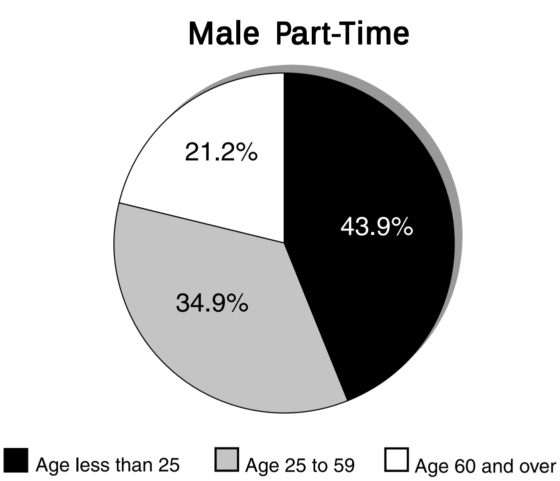
The Relative Compensation of Part-Time and Full-Time Workers: The Role of Worker and Job Skills

I. Introduction

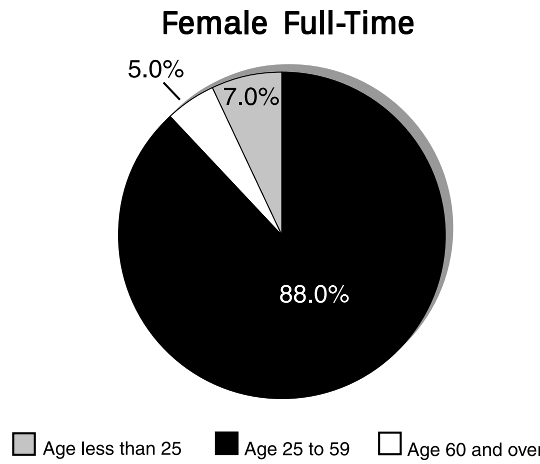
The use of part-time workers by employers has received increased public scrutiny, particularly following the summer 1997 strike at United Parcel Service (UPS), where the wages and use of part-time workers were key issues of conflict.[1] Part-time jobs pay substantially lower wages and benefits than do full-time jobs. Although some of the wage differential can be accounted for by standard measures of worker and job attributes (e.g., age, education and industry), most studies conclude that a sizable gap remains. For example, Blank (1990) obtains part-time penalty estimates of 19% and 26% for women and men, respectively, after controlling for a large number of worker, job and labor market characteristics.



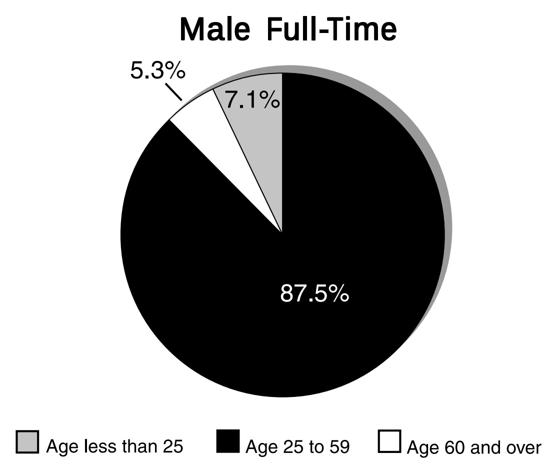


Understanding the sources of the part-time penalty is important for at least two reasons. First, part-time employment is widespread. In 1997, 17.9% of workers were part-time, 26.4% of all women and 10.7% of men (U.S. Bureau of Labor Statistics 1998c, Table 8). Second, standard theory (ignoring quasi-fixed costs) suggests that marginal labor costs to employers should be equivalent for workers with the same productivity, or, from the perspective of employees, compensation should be equivalent for workers with equal skills and preferences in equally attractive jobs. Gaining an understanding of the sources of the large part-time/full-time wage differential, therefore, is important if we are to gain insight into how labor markets operate and how closely they approximate the competitive model.

In this paper, we examine the role of worker-specific skills, occupational skill requirements and job working conditions on the part-time/full-time wage differential. The analysis extends previous research in two principal directions. Most important, we construct large panels of workers from the Current Population Survey (CPS) in order to observe wage *changes* among individual workers as they move from full-time to part-time or from part-time to full-time employment. The longitudinal analysis provides a method for controlling for otherwise unmeasured individual-specific skills or preferences that remain constant across jobs. In addition, we incorporate information from the Labor Department's new *Occupational Information Network* (known as *O*NET*) database on job skill requirements and working conditions.



Our analysis indicates that a moderate amount of the part-time penalty is explained by differences in occupational skill requirements, as measured either in *O*NET* or by occupational controls. The part-time wage disadvantage does not result because of less onerous working conditions in part-time than in full-time jobs. Longitudinal analysis indicates that worker-specific skills and preferences account for a substantial portion of the part-time wage differential. Among women and men switching part-time status but not occupation and industry, there is effectively no wage change associated with the change in part-time status. Among workers switching both occupation and industry, however, there remains a small but nontrivial wage penalty associated with part-time employment. Evidence for a part-time wage gap is particularly weak among women, who constitute 68.0% of the part-time work force (U.S. Bureau of Labor Statistics 1998c, Table 8).



In Section II, we develop various explanations for why wage differentials exist between part-time and full-time jobs, and review previous literature on part-time wage differentials. The estimation approach is outlined in Section III. Section IV describes our data, while Section V provides descriptive evidence. Regression results from wage level and wage change models are presented in Section VI. Section VII provides additional evidence on the part-time gap, followed by conclusions in Section VIII.

II. Why Do Part-Time Jobs Pay Less?

Theory

Part-time wage differentials can result from, among other things, differences in labor supply among heterogeneous workers, fixed employment costs, job search differences

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and worker skills. The labor supply explanation starts from the premise that a large number of workers are willing to work part-time but not full-time or, stated alternatively, an even larger full-time wage advantage would be required to attract more workers into full-time employment. For example, young people working while in school, adults heavily engaged in home production, or older workers who have moved out of career jobs may prefer part-time employment.

Labor supply differences by themselves should not produce a part-time wage penalty. If part- and full-time employees had identical skills and there were no fixed costs to employers associated with employment, wage rates would be equalized as firms simply adjust their worker mix toward more part-time employees. A wage differential will arise, however, if there exist fixed employment costs or if part- and full-time workers are not fully fungible. For example, college towns have many students in the labor force who are not willing to work full-time and who possess a different set of skills and preferences than the full-time labor force. A large supply of students willing to work part-time creates an equilibrium full-time/part-time wage gap that is not eliminated by employee movement across jobs or employer shifts in the hours mix within jobs. In short, different preferences and labor supply across heterogeneous or noncompeting skill groups can create an equilibrium wage differential between part- and full-time workers.[2]

An additional factor producing an equilibrium wage differential is the existence of fixed employer costs. Costs associated with recruiting, hiring, training, personnel management and employee benefits whose costs are not proportional to hours worked (e.g., health insurance) increase the average cost of part-time relative to full-time workers for any given wage rate. There must exist a wage differential that lowers the cost of employing part-time workers so that the marginal product per dollar is equivalent for part-time and full-time workers. Note that while marginal *employment* costs (which include fixed costs) are equalized, the marginal costs of *hours* will be lower for part-time workers. Hence, where part-time and full-time workers possess similar skills, one can expect employers to desire more hours from their lower-wage part-time employees than those workers are willing to offer.[3] Not only do fixed employment costs produce a part-time wage penalty, it also follows that part-time employees are likely to receive fewer nonwage benefits (e.g., health insurance, pensions, paid leave) than are full-time workers, as a means of reducing the cost and wage differential between part-time and full-time workers. Were it not for the

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lower benefits, the part-time *wage* gap would be even larger.

Dual labor market and search theory might also account for a part-time wage penalty, assuming that part-time jobs are more prevalent in secondary than primary labor markets. If workers in secondary labor markets invest less time in job search than do workers in primary markets, or "park" in secondary jobs while searching for primary jobs, the wage penalty for part-time workers will increase owing to differential search and differences in the quality of the job match. If a substantial number of workers were "involuntary" part-time workers (i.e., seeking but not obtaining full-time employment), the differential might be large if search for higher-paying part-time jobs is limited while workers search for full-time employment.

Although emphasis is typically on those factors leading to lower part-time wages, some forces work in the opposite direction. If businesses have peak customer and labor demand over short time periods (e.g., a restaurant with mealtime peaks), but workers prefer continuous hours rather than brief or split shifts, equilibrium part-time wages would be higher than full-time if few workers prefer part-time hours. More generally, firms' use of part-time workers as a low-cost means of varying labor input in the face of varying and uncertain demand requires that there be a relatively large supply of part-time workers. And to the extent that there is limited mobility or substitution across labor markets (delineated by geography, occupation and, possibly, industry), equilibrium part-time gaps should vary across markets.

The principal focus in this paper is on occupational skills and working conditions and (unmeasured) worker-specific skills. It is unlikely that differences in working conditions (e.g., job hazards, strength requirements) can account for much of the part-time penalty; indeed, we know of no evidence suggesting that part-time jobs have less onerous working conditions, on average, than do full-time jobs. Differences in job and worker skills, however, are likely to account for a sizable portion of the full-time wage advantage. Because current work hours are highly correlated with past hours of work (Blank, forthcoming), part-time workers typically have fewer skills and less general and firm- and industry-specific training. Unlike most previous work, our analysis attempts to better account for worker skills through the explicit measurement of occupational skill requirements and through the indirect measurement of individual-specific skills accounted for through the use of longitudinal analysis.

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Prior Evidence

There are surprisingly few studies whose principal focus is the wage differential between full-time and part-time workers. But there are hundreds of studies in which part-time status is included as a control variable in a log wage equation. Such studies typically find a significant wage differential between full- and part-time workers, following control for individual, job and labor market characteristics included in widely-used data sets.

We do not attempt an exhaustive survey of the literature but, rather, focus on work most closely related to our analysis. Blank (1990) provides detailed descriptive data identifying the extent and nature of part-time work and the wage penalty associated with it. As stated at the beginning of our report, Blank finds a large part-time penalty using standard regression analysis. She then attempts to control for unmeasured skill differences between part-time and full-time workers through the use of cross-sectional "instrumental variable" methods that endogenize or statistically account for the determination of part-time status. She concludes that selection into part-time employment is important, erasing the part-time wage penalty among women. Blank further notes both the importance of occupation in accounting for the part-time penalty and stresses the desirability of using panel data to control for worker-specific quality. In this paper, we provide longitudinal analysis using panel data to measure the part-time wage gap and explicitly account for a wide variety of occupational skill measures.

In a subsequent paper, Blank (forthcoming) provides longitudinal analysis from the Panel Study of Income Dynamics (PSID) tracking workers as they move into and out of part-time employment. Quite relevant to our analysis, she concludes that individual-specific unobservables are important predictors of transitions between part-time, full-time and out-of-the-labor-force status. A relatively short time series provides good predictions of future transitions (or absence thereof), providing some support for our use of a panel data set with only two observations per worker. Although Blank is most concerned with explaining employment transitions, her work supports our conjecture that unmeasured person-specific skills and tastes may be important determinants of wage differences between part- and full-time workers.

In a valuable recent paper, Lettau (1997) uses establishment level data collected by BLS for the Employment Cost Index (ECI) program. Measuring wage and total compensation differences between part-time and full-time workers within the same establishments and occupation, Lettau finds an overall part-time wage penalty of

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about 15% and a total compensation penalty of about 20% (-.164 and -.227 log points, respectively). Lettau is unable to control for worker-specific skills (e.g., schooling, experience, tenure, gender). Thus, one cannot rule out the possibility that much of the part-time wage disadvantage reflects differences in worker skills *within* establishments and occupations. An earlier study by Montgomery and Cosgrove (1995) comparing part-time and full-time wages in child care establishments among teachers and aides, controls for schooling and experience as well as establishment and occupation. They find a part-time wage disadvantage of about 8%-9% using OLS. Using IV or random effects estimation, they find effectively no wage gap among teachers and approximately a 7% gap among aides. It is not clear to what extent results from their sample can be generalized to the larger labor market. An important advantage of our study is that we account through the use of longitudinal analysis for what turns out to be important worker-specific skills and other fixed effects, factors not accounted for in most previous studies.

An issue relevant to our analysis is the definition of part-time and full-time employment. The official BLS definition is based on whether a worker usually works 35 or more hours per week *on all jobs*. That is, part-time and full-time status are defined for individual *workers* and not for individual *jobs* (Nardone 1995). Because we are interested in pay differences to workers in full-time and part-time jobs we define part-time status (and pay) based on usual hours worked (and earnings) on the individual's *principal* job.

An additional issue is whether 35 hours constitutes an appropriate breakpoint for defining part-time status. Papers by Hotchkiss (1991) and Averett and Hotchkiss (1996) explore the statistical justification for this definition, based on joint estimation of the labor supply choices of workers and wage determination among part-time and full-time jobs. The latter paper concludes that although men begin receiving a full-time premium at about 33 hours, white (black) women are not offered full-time wage premiums until roughly 37 (39) hours of work. Apart from the difficult theoretical and empirical issues involved in making such a determination, such analysis is limited by the fact that few workers report usual hours worked other than at hours amounts divisible by five or eight. Whether wages vary discretely and/or continuously with respect to hours worked is an important question, however, and we return to it subsequently.[4]

III. Estimation Approaches

The estimation strategy followed in this study is straightforward. First, relatively standard log wage regressions are estimated. Using this approach, the wages for part-time workers are compared to the wages for full-time workers with the same measured characteristics. We then extend the analysis by adding a large set of occupational skill and working condition variables. In this case, we compare part-time and full-time workers with similar characteristics *and* working in similar jobs. We then turn to longitudinal analysis, in which we measure wage changes for individuals as they move from part-time to full-time work, or vice versa. In this analysis, the comparison group is not a different set of workers with similar measured characteristics. Rather, the comparison wage for each individual worker is his or her *own* wage one year earlier, prior to a change in part-time status. In this way, unmeasurable worker attributes affecting earnings that are constant over a year (e.g., motivation, reasoning ability, preferences) are automatically controlled for, and the part-time wage disadvantage is estimated by the average change in wages across individuals changing part-time status.

In the analysis that follows, separate log wage equations are estimated for female and male workers, since many wage equation parameters, including the full-time/part-time gap, differ substantially between women and men. For ease of exposition, we use a simple dummy variable approach in order to measure the log wage differences associated with part-time status, conditional on controls.

The general form of the wage equation model is:

$$(1) \quad \ln W_{it} = X_{it}\beta + \theta PT_{it} + \varepsilon_{it}, \text{ with } \varepsilon_{it} = \Phi_i + \mu_{it}$$

where lnW_{it} is the natural log of real hourly earnings of individual i in year t; X is a vector of individual, job, and labor market characteristics defined at the individual level, with β the corresponding coefficient vector (including an intercept); PT is a binary variable equal to one if the worker's principal job is part-time (defined initially as less than 35 hours usually worked per week) and θ is an estimate of the part-time log wage penalty; and ε is the error term. ε includes both a random component μ with mean zero and constant variance, and a worker-specific fixed effect Φ. The fixed effect Φ reflects unmeasured individual-specific skills and worker preferences regarding the nature of work and pay similarly reflected across jobs in consecutive periods. The inability to measure Φ causes estimates of θ in (1) to suffer from omitted variable bias if Φ is correlated with part-time status. For example, if PT and Φ are negatively correlated owing to lower unmeasured skills among part-time workers, estimates of θ are likely to have a negative bias and overstate the part-time wage penalty.

In order to reduce bias in estimates of θ we follow two strategies. First we add explicit measures of occupational skills and working conditions to X, intended to capture typically omitted wage determinants correlated with part-time status. Second, we estimate longitudinal wage change models that account for (i.e., net out) worker-specific skills and other fixed effects Φ. Using multiple short panels with two observations per worker, one year apart, we estimate the following equation:

$$(2) \quad \Delta \ln W_i = \Delta X_i \beta' + \Delta PT_i \theta' + \Delta \varepsilon'_i, \text{ with } \Delta \varepsilon'_i = \Delta \mu_i$$

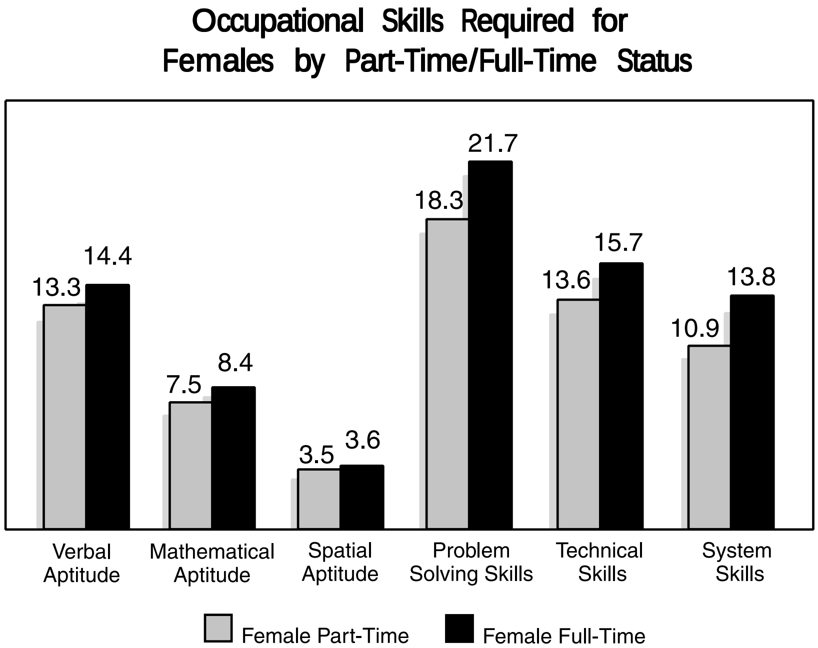
Here, Δ represents the change operator between year *t* and *t-1*. Longitudinal estimates of θ′ reflect wage changes between years *t-1* and *t* for given individuals as they switch between full-time and part-time status.

While having important advantages, longitudinal analysis is not without shortcomings. The panel sample is not fully representative, measurement error in the change variables can bias estimates of θ′ toward zero, and, while accounting for problems associated with the nonrandom assignment or endogeneity of *PT* in (1), equation (2) assumes that *changes* in part-time status are exogenous. These issues are addressed subsequently, following presentation of the paper's principal results. We conclude that none of these potential problems turn out to be serious concerns in the estimation of the part-time/ full-time wage gap.

IV. Data and Variables

The primary data used in the paper are from the Current Population Survey (CPS) Outgoing Rotation Group (ORG) earnings files for January 1989 through December 1997. The structure of the CPS permits one to match given individuals in the same month, one year apart. We construct a large panel data set for 1989/90 through 1996/97 from the CPS ORG.[5] Our sample consists of wage and salary workers, ages 16 and older (17 and older in the second year of each panel). Excluded are those who are not in wage and salary employment in consecutive years, those whose weekly earnings are top-coded by the Census (at $1,923) in either year (since the wage change would then be determined by the earnings assignment(s) in the open-ended category), workers with an implied wage (i.e., real weekly earnings in 1997 dollars divided by hours worked per week) less than $1 or more than $99.99 (the latter corresponds to someone receiving weekly earnings near the cap and working less than 20 hours), workers for whom the usual weekly hours of work variable has been allocated by the Census in either year (because reliable allocation flags on hours are not available for all years, some observations with allocated hours cannot be deleted) and workers who cannot be matched across years. Because the Census reinterviews households in fixed locations, individuals whose household moves or who move out of a household during the year are not in the sample. Young workers are most likely to be underrepresented (Peracchi and Welch 1995, Card 1996).

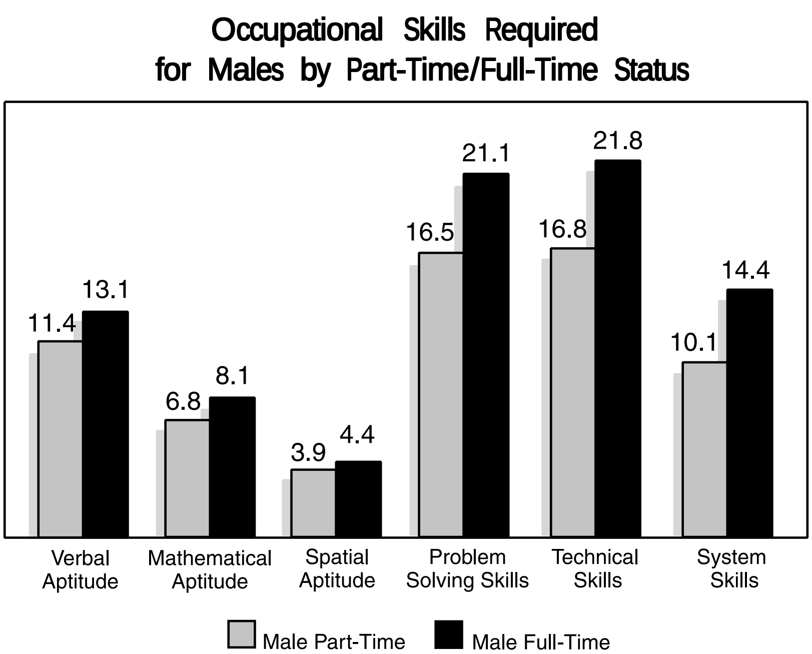
To insure comparability between the wage level and wage change samples, the panel data set is used for estimation of both equations (1) and (2), with the levels equations based on second-year observations for each worker. Wage level estimates using the full CPS ORG data set are very similar. In addition to the part-time status variable, we include the following control variables in *X* — years of schooling completed, potential experience (measured by years out of school — the minimum of age minus schooling minus 6 or of age minus 16) and its square, marital status (2 dummies included), race and ethnicity identifiers



(4), children in household (3), region (8), metropolitan size based on 1990 Census population counts (7), union membership, industry (12), occupation (11), and year (7).

*O*NET* is the new Department of Labor database that is intended to replace the *Dictionary of Occupational Titles*. *O*NET* provides hundreds of numerical descriptors of occupations. Details regarding *O*NET* are provided in the note to Appendix Table A-1. In specifications utilizing information from *O*NET* we include highly-aggregated variables measuring the following occupational skills and working conditions:

Occupational Skill Requirements: *Verbal* skill measuring Oral Comprehension, Written Comprehension, Oral Expression, and Written Expression. *Math* skill variables measure Mathematical Reasoning, Number Facility, and Mathematics. *Spatial* measures are Spatial Orientation and Spatial Visualization. *Problem Solving* skill variables measure Problem Identification, Information Gathering, Information Organization, Synthesis/Reorganization, Idea Generation, Idea Evaluation, Implementation Planning, and Solution Appraisal. *Technical* skill variables measure requirements for Operations Analysis, Technology Design, Equipment Selection, Installation, Programming, Testing, Operation Monitoring, Operation and Control, Product Inspection, Equipment Maintenance, Troubleshooting and Repairing. *System* skill variables measure Visioning Systems, Perception, Identifying Downstream Consequences, Identification of Key Causes, Judgment and Decision Making and Systems Evaluation.



Occupational Working Conditions: Six *Hazard* variables measuring the frequency times degree of injury for Radiation, Diseases/Infections, High Places, Hazardous Conditions, Hazardous Equipment and Hazardous Situations. Six *Environmental* variables — Distracting Sounds and Noise Levels, Extremely Bright or Inadequate Light, Exposure to Contaminants, Cramped Work Space or Awkward Position, Whole Body Vibration and Prolonged Exposure to Very Hot or Cold. Five *Strength* variables measuring Static Strength, Explosive Strength, Dynamic Strength, Trunk Strength and Stamina.

In the longitudinal wage change analysis, many of the variables in *X* are time invariant. Included in addition to the change in part-time status is the change in union status, changes in experience squared (the change in experience is a constant), changes in broad industry, changes in broad occupation and changes in the *O*NET* occupational variables, plus year dummies. By construction, region and city size do not change because households that move cannot be matched in the CPS panel. Remaining variables either cannot change by construction of the panel (e.g., gender) or are treated as invariant since recorded changes may result from measurement error or have insufficient time to be reflected in earnings differences (e.g., race, schooling, children in household).

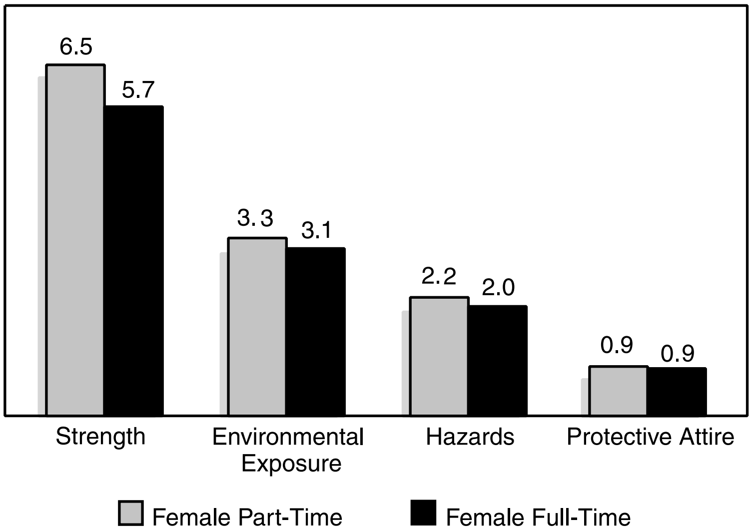
V. Descriptive Evidence

Variable Means for Part-Time and Full-Time Women and Men

Prior to evaluating the empirical analysis, a comparison of characteristics among part-time and full-time workers is informative. Table 1 provides means of selected variables for part-time and full-time women and men based on our 1989-97 CPS sample and *O*NET* occupational characteristics.

As is widely recognized, part-time workers have lower wages and are more likely to be female, either young (among men) or old, and nonunion. Part-time women are relatively more likely to be married with spouse present and have more children, while the opposite is true for men. The principal contribution of Table 1 is evidence comparing means of the occupational *O*NET* variables between part-time and full-time workers. For convenience, Table 1 provides means of *O*NET* variables following aggregation into general categories. Appendix Table A-1 presents means on each of the individual variables, aggregated into the broad variables included in the regressions. Note that both full-time and part-time workers *within* occupations are assigned identical values of the *O*NET* variables, so differences in means are due entirely to differences in the occupational structure between part-time and full-time workers. To the extent that there exist *within-occupation* differences in skill requirements and working conditions between part-time and full-time workers, as is likely, the mean differences reported in Table 1 *understate* the total part-time/full-time gap in skills and working conditions.

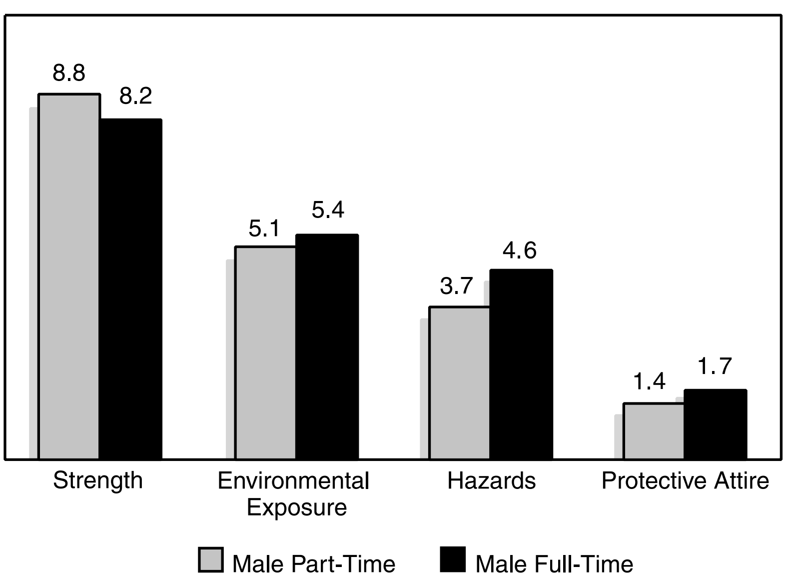
Occupational Working Conditions for Females by Part-Time/Full-Time Status



Three important patterns are evident in Table 1.[6] First, and most important, there exists a substantial gap in required occupational skills between part-time and full-time workers. Consistent with the thesis that skill differences explain some and possibly much of the part-time wage disadvantage, we find that full-time workers are employed in occupations requiring higher levels of verbal, mathematical, problem solving, technical and system skills. The one exception is spatial skills, where little difference is found.

The second pattern is that the part-time/full-time gap in occupational skill requirements (ignoring spatial skills) is systematically larger for male than female workers. This pattern helps account for what is a substantially larger part-time wage gap among men than women. Although the gender gap in wages is not the focus of this paper, it is worth noting that *O*NET* skill requirements are *not* systematically lower for women than men. Among both part-time and full-time workers, women are employed in occupations requiring somewhat higher levels of verbal, math and problem solving skills than are males, and lower levels of spatial and technical skills.

Occupational Working Conditions for Males by Part-Time/Full-Time Status



A third pattern evident from Table 1 is that there is little systematic difference in working conditions between the occupations in which part-time and full-time workers are employed. Although the *O*NET* evidence suggests that some portion of the part-time wage gap is related to job skills, occupational working conditions appear unlikely to account for the part-time gap. If anything, part-time females are employed in occupations with somewhat greater hazards, strength requirements, and environmental risks. The pattern among men is mixed, with full-time males in occupations with somewhat greater hazards and environmental risks, but not strength (the latter may not be a clear disamenity for men). Although differences in working conditions appear to be far less important than skill requirements, less attractive working conditions for part-time relative to full-time women, as compared to differences among men, are consistent with there being a smaller part-time wage gap among women than men.

[F]ull-time workers are employed in occupations requiring higher levels of verbal, mathematical, problem solving, technical and system skills.

The Distribution of Hours Worked

We next provide graphic representations of the distributions of hours worked and wages by hours worked. Figures 1a and 1b show the frequency distribution of usual hours worked per week for women and men, respectively. In order to increase sample sizes, we employ the full CPS ORG earnings files for 1989-97. There are several points worth emphasizing. First, there is a heavy concentration of workers reporting 40 usual hours worked per week, 51% among women and 57% among men. Second, the hours distribution for women is more dispersed than for men and contains more low-hour and fewer high-hour observations. And third, there exist "spikes" or "heaping" at intervals divisible by five, a common survey phenomenon.[7] If we instead examine the distribution of "hours worked last week" we obtain a more dispersed hours distribution (with similar but slightly lower means) with fewer workers reporting exactly 40 hours, 39% among women and 43% among men.[8] Although the hours worked last week variable, averaged over workers and time, provides a more accurate summary of the distribution of hours worked in a typical week, usual hours worked per week should correspond more closely to the CPS earnings measure of usual weekly earnings.

Figure 2a and 2b show mean wage rates by hours worked between 20 and 60 hours. Note first that there is a large degree of noise in both tails of the distribution and at hour intervals not divisible by five. This is to be expected, since few workers are observed at these points and there is substantial measurement error in the wage (calculated as usual weekly earnings divided by usual hours worked per week) among persons reporting very low and high hours of work. Most relevant for our analysis is the finding of a reasonably sharp break in wages around 35 hours per week, rather than a gradual increase in wages beginning at low levels of hours worked.[9] Although a jump in wages is evident for women and men, the full-time wage advantage is clearly larger among men than among women.

[O]ccupational working conditions appear unlikely to account for the part-time gap.

The wage break at 35 hours provides justification for following the convention, as we do in this paper, of defining part-time status as a binary rather than continuous variable, with the breakpoint being less than 35 versus 35 or more hours worked per week. That being said, mean wages do tend to increase with respect to hours among full-time workers. This is particularly evident if one focuses on mean wages at the rounded hour intervals (i.e., 40, 45, 50, etc.), where sample sizes are largest.[10] For this reason, we subsequently examine the sensitivity of results to the use of an alternative definition. Of course, the figures in Figure 2 are mean wages *without* control for worker or job characteristics. If worker skills vary with hours worked, reliable inferences about how wages vary with hours worked must await the empirical analysis provided below.

VI. Wage Level and Longitudinal Estimates of the Part-Time Penalty

Wage Level Estimates of the Part-Time/Full-Time Wage Gap

Table 2 provides regressions estimates of the part-time wage disadvantage, based on alternative specifications of wage level equations. Wage level regressions are for the years 1990-97, based on the second year observation for each worker in our panel data set. Longitudinal estimates, shown subsequently in Table 3, are based on wage change regressions for the periods 1989/90-1996/97.

As seen in line A-1, the part-time/full-time real log wage penalty, *unadjusted* for worker, job, or labor market characteristics is -.300 log points for women and -.620 log points for men. This implies that part-time women earn 25.9% less per hour than full-time women, while part-time men earn 46.2% less.[11] The unadjusted log point differentials are the benchmark figures that we will use in order to evaluate how much of the part-time wage disadvantage can be accounted for by measured characteristics and unmeasured worker-specific skills.

In line A-2, estimates of the part-time/full-time wage differential, θ, are based on a standard wage level equation with controls for individual and location characteristics, but not industry or occupation. We obtain part-time coefficients of -.201 and -.323 for women and men, or part-time wage disadvantages of 18.2% and 27.6%, respectively. These estimates imply that a sizable portion of the large part-time wage differential can be readily accounted for by measurable personal and locational characteristics (in particular, age or experience). Returning to our benchmark figures, 33% of the female gap (1-.201/.300) and 48% of the male gap (1-.323/.620) are readily accounted for by the variables. The remaining log wage gaps, of course, remain sizable.

In line 3 we add industry dummies, reducing estimates of θ to -.155 for women and -.263 for men. Addition of occupation dummies (line A-4) further reduces estimates of the wage gap, to -.126 for women and -.222 for men. Estimates in A-4 are based on a relatively dense specification using standard information available in micro data sets. In Line A-5, the *O*NET* occupational variables are substituted for the broad occupation dummies. These reduce the gap by more than does inclusion of the occupation dummies, but the difference is small. Our "best" estimate of the part-time wage gap using wage level analysis is shown in line A-6, with inclusion of all CPS and *O*NET* variables. Our estimates of θ are -.109 for women and -.200 for men. Overall, controlling for *measurable* personal, location and job characteristics accounts for about two-thirds of the total part-time wage disadvantage — 64% for women (1-.109/.300) and 68% for men (1-.200/.620).

Recall that one point of issue is whether a single part-time differential is a reasonable approximation of the way labor markets work, or whether the wage gap between part-time and full-time workers increases with the level of hours. In order to examine this issue, Table 2 provides estimates of wage differences associated with alternative levels of hours worked, based on inclusion of five dummy variables, with 40 hours worked per week the omitted reference category.

Line B-1, based on a regression *absent* controls, effectively summarizes the information pictured in Figures 2a and 2b. The part-time gap is roughly similar between 1-24 and 25-34 hours worked, while among full-time workers the wage increases with hours worked. A sizable wage difference exists between those working 41-49 hours versus 35-39 hours — .177 log points among women and .189 among men. Following control for worker and job characteristics (line B-2), the relationship between hours and wages is much weaker. There is no meaningful difference in wages among those working 1-24 versus 25-34 hours. And differences between those working 40 hours and those working 35-39 and 41-49 are modest.

Interestingly, women and men reporting 50 or more usual hours per week have lower implicit hourly earnings (i.e., weekly earnings divided by weekly hours) than do those working 40 hours. This may result from a high rate of overreporting or mismeasurement of hours worked among those with 50 or more hours, or because those working long hours have relatively low marginal disutility from work and require lower wages, *ceteris paribus*. Since those reporting very high hours constitute a small part of the work force, we do not explore these issues further. In general, the results reported in B-2 indicate that use of a single part-time dummy with 35 hours delineating full-time employment is not only convenient, but also not too unreasonable a strategy for approximating the part-time wage gap.

[C]ontrolling for measurable personal, location and job characteristics accounts for about two-thirds of the total part-time wage disadvantage—64% for women and 68% for men.

Longitudinal Estimates of the Part-Time/Full-Time Wage Gap: Initial Results

Longitudinal estimates of the part-time wage gap, θ', are based on the wage gains and losses of individual workers switching into full-time and part-time status, respectively. Table 3 provides estimates of θ' from wage change equations estimated for the pairs of years 1989/90-1996/97. Longitudinal estimates of θ' vary relatively little with respect to specification density (i.e., the number of explanatory variables) because much of the skill difference between workers is captured by person-specific fixed effects, which are netted out or differenced in the longitudinal analysis. Therefore, we present results only from specifications with no controls and the full set of controls.

The full model coefficient θ' on ΔPT (line A-2) is very small for women — -.014 — suggesting little wage gap for women in part-time jobs, as compared to women with equivalent skills in similar full-time jobs. The ΔPT coefficient for men of -.043 is more substantial, but still very tiny compared to the unadjusted .62 wage gap between part-time and full-time men.

The *apparent* inference from these results is that *virtually all* of the part-time wage disadvantage can be accounted for by what are lower worker-specific skills among part-time than full-time workers, with many of these worker skills not measured by standard variables. On average, individual workers realize little change in hourly pay as they move between full-time and part-time jobs. Before accepting the sweeping conclusion that there exists little part-time/full-time wage differential for truly similar workers, it is important that several dimensions of the longitudinal results be probed in some detail. We turn to this task below.

There is no meaningful difference in wages among those working 1-24 versus 25-34 hours.

Probing Longitudinal Estimates of the Part-Time Wage Penalty

Symmetry Between Wage Gains and Losses. The longitudinal results in line A of Table 3 impose symmetry between the wage gain from switching from a part-time to full-time job and the wage loss from switching from full-time to part-time. In line B of Table 3, this restriction is relaxed, with separate wage changes estimated for full-time stayers (the reference group), part-time stayers, part-time leavers and part-time joiners.[12]

[V]irtually all of the part-time wage disadvantage can be accounted for by what are lower worker-specific skills among part-time than full-time workers, with many of these worker skills not measured by standard variables.

Focusing first on results with control variables (line B-2), we find similar wage changes (in absolute value) for part-time joiners and leavers. Among women, part-time workers switching to a full-time job realize only a .020 wage gain relative to part-time stayers. Women switching from full-time to part-time realize a .013 wage loss relative to the reference group of full-time stayers. As seen by the F test on the last line of B-2, we cannot reject the null of symmetric wage change among part-time leavers and joiners. Among men, wage changes for joiners and leavers is virtually equivalent — .044 for part-time leavers versus -.042 for part-time joiners.

Although not the focus of the study, we also find no significant difference in wage growth among part-time and full-time *stayers*, as seen by the small and insignificant coefficients on the dummy variable coded 1 for *PT* in the initial period. Because the results in line B-2 indicate no substantial wage change asymmetry, in the analysis to follow we reimpose the simplifying assumptions of symmetric wage change among those leaving and joining part-time employment, and equivalent wage change among part-time and full-time stayers.

Do Wage Changes Vary with Change in Hours Worked?

Up to this point, we have made the simplifying assumption that the change in hourly earnings associated with a change in part-time status is the same for both small and large changes in hours worked per week. In this section, we allow the wage change estimates of the part-time differential to vary with the magnitude of the change in hours worked.

As seen in line C-2 of Table 3, little or no wage change is found among the relatively few part-time switchers changing usual hours worked by fewer than 10 per week. Among women, there is an approximate .02 log point wage change among those changing part-time status and hours worked by 10 or more. Among men, wage change is a more substantial .09 among those changing part-time status and hours worked by 15-19 hours, .05 among those in the 10-14 hours change interval, and relatively small among those with small or extremely large hours changes. Relatively few men who change part-time status report small hours changes.

Our reading of these results is that there exists little variation with respect to hours change in the magnitude of the small part-time wage differential among women. But among men, the change in hourly earnings does tend to increase with the change in

[L]ittle or no wage change is found among the relatively few part-time switchers changing usual hours worked by fewer than 10 per week.

hours. The exception is among men who report unusually high hours worked, which leads to a low value of the implied wage.

Are Longitudinal Estimates Biased by Misclassification Error?

Misclassification or measurement error in right-hand-side change variables presents the potential for bias toward zero in estimated coefficients. Past research has indicated that bias on longitudinal estimates of the *union* wage effect is particularly serious (Freeman 1984, Card 1996). In the case of part-time status, if there were a large number of individuals incorrectly classified as changing part-time status, relative to the number who actually change status, the bias in θ' toward zero would be severe. Although our expectation is that the rate of misclassification error in the case of part-time changers should be low, it is important that this issue be examined to avoid an incorrect interpretation of the longitudinal results.

Hirsch and Schumacher (1998) have shown that the misclassification error in union status change is reduced substantially by focusing on union status changers who also change detailed occupation and industry. The logic is that most of these individuals will be true union changers, whereas those not changing occupation and industry are less likely to be true changers. They implement this approach by interacting *ΔUnion* with dummies for those who change detailed occupation and industry, industry only, occupation only and neither occupation nor industry.

In line D of Table 3, we present results from specifications interacting *ΔPT* with dummies designating the four alternative groups of industry and occupation changers — workers changing both detailed industry and occupation, industry only, occupation only and neither industry nor occupation. Workers recorded as changing both detailed industry and occupation are more likely to be true job switchers and, therefore, true changers of part-time status. Estimates of the part-time penalty based on this group, therefore, are least likely to be biased by misclassification error.[13]

Relatively few men who change part-time status report small hours changes.

As it turns out, we later conclude that the bias from misclassification error in part-time status is small. But our analysis reveals what we believe are substantially different wage effects of part-time employment for workers switching both occupation and industry and those not. Focusing on the *ΔPT* for industry and occupation changers in the specification

[T]here does exist a non-trivial part-time wage penalty among workers switching industry and occupation.

with full controls (line D-2), we find an approximately .069 log point wage change associated with the change in part-time status among women, and .105 log wage change among men. These estimates can be compared to the .11 and .20 wage differentials for women and men found using wage level analysis (Table 2, line A-6), which did not control for unmeasured worker fixed effects. In short, there does exist a nontrivial part-time wage penalty among workers switching industry and occupation, but the wage disadvantage is considerably smaller than that suggested by standard wage level analysis, including regressions with detailed job characteristics. Even if we were to treat θ' among the industry and occupation switchers as the "true" part-time wage penalties, these differentials are a small fraction of the total unadjusted log wage gap — 23% for women (.069/.300) and 17% for men (.105/.620).

Longitudinal estimates of θ' in line D-2, based on part-time status switchers *not* changing both detailed industry and occupation, are close to zero. No evidence is found for a statistically significant part-time wage penalty among these workers (a small part-time *advantage* is found for those switching neither occupation nor industry).

For several reasons, we do not believe that misclassification error in ΔPT is large or can account for much of the difference in estimates of θ' between those who do and do not change both industry and occupation. First, the CPS industry variable is recorded with considerably greater accuracy and consistency than is occupation (see Polivka and Rothgeb 1993). If misclassification error were driving our results, estimates of θ' for those changing industry only would be similar to those for industry and occupation switchers (i.e., since both groups would be true switchers) and dissimilar from those who are recorded as changing only occupation, a large number of whom would not be job or *PT* switchers.[14] Yet industry-only switchers have estimates of θ' similar to the latter group. Second, measurement error in ΔPT, constructed primarily from responses to the question on usual hours worked per week, is not likely to be substantial. Most persons recorded as changing part-time status indicate a large change in hours worked, with relatively few workers classified as changing part-time status because of small changes in hours, say from 34 to 36 hours a week.

[A] small part-time advantage is found for those switching neither occupation nor industry.

Third, bias toward zero from measurement or misclassification error varies with the ratio of error variance to true variance in part-time status. Women display substantially higher rates of part-

[P]art of what we are calling a penalty for part-time employment is in fact a wage differential associated with career and noncareer jobs.

time status switching than do men. If misclassification rates for part-time status were similar for women and men, then coefficients on ΔPT should be most biased (i.e., driven closer to zero) for men, since they have lower true rates of part-time switching. Yet estimates of θ' are larger (in absolute value) for men than for women, -.043 versus -.014 in the model with full controls (line A-2), and the proportional decline in θ' relative to the wage level estimate θ, is *smaller* among men than women.

For the reasons above, we conclude that misclassification error is not the principal explanation for differences in longitudinal part-time gap estimates among the different groups of switchers. Rather, these wage gap differences appear to be real. We are thus left to explain why there is a wage effect associated with changing part-time status for workers changing industry and occupation, but not for other part-time status changers. We do not have a convincing explanation. The answer is *not* that it results from a greater loss in specific human capital when there is an industry and occupation change. Workers who *remain* full-time (or part-time) across years but change occupation and industry also lose specific capital. Note also that being held constant are changes in broad industry and occupation, as well as dummies designating whether a worker changed industry and occupation, industry only, or occupation only (coefficients on these three dummies are effectively zero). We can think of no reason why gains or losses in specific human capital should be larger for part-time switchers than for part-time stayers who also change industry and occupation, as this interpretation would imply.

Although we find no obvious explanation for why relative part-time/full-time pay should vary with industry and occupational switching, some possibilities are explored. Examining the characteristics of the different groups of part-time switchers, it is found that for the group changing occupation and industry as well as part-time status, there are more young workers and the mean change in hours worked is larger than among other part-time switchers. Yet these differences appear to account for a rather modest amount of the difference. This is not surprising since wage change associated with large hours changes are similar or smaller than among moderate sized changes (Table 3, line C-2). Nor does controlling for age differences account for much of the difference among the groups of changers (this work not shown).

The closest we can come to a satisfactory interpretation of this result is that workers switching from part-time to full-time status, occupation and industry are workers most likely to be shifting from low-paid non-career jobs into higher-paying career jobs. Likewise, workers switching from full-time to part-time along with occupation and

There exists little wage gap between similarly skilled part-time and full-time workers.

industry are most likely to be moving out of a career job. Thus, part of what we are calling a penalty for part-time employment (or a premium for full-time) is in fact a wage differential associated with career and noncareer jobs.

Although we are unable to satisfactorily explain why longitudinal estimates of θ' differ with respect to occupation and industry change, our principal conclusions are not dependent on an explanation. Whether one focuses on the very tiny *average* wage change among those changing part-time status, or the moderate wage change exhibited by occupation and industry switchers, the same conclusion follows. Most of the very large part-time wage gap can be accounted for by measurable worker and job characteristics and by unmeasured worker-specific skills. There exists little *wage* gap between similarly skilled part-time and full-time workers.

Endogenous Job Change and Bias in Longitudinal Estimates. The longitudinal analysis presented above has the important advantage of controlling for otherwise unmeasured worker-specific skills and preferences correlated with part-time status. A limitation often associated with longitudinal analysis is that the *change* in part-time status may be correlated with wage change, potentially biasing longitudinal estimates.

As it turns out, such bias is *advantageous* in this analysis because the direction of bias is known *a priori* and differs between joiners and leavers. Thus, wage gap estimates for joiners and leavers place lower and upper bounds on the part-time wage penalty. Specifically, we expect the probability of workers switching from part- to full-time employment to be positively correlated with wage change, and the probability of switching from full- to part-time employment negatively associated with the wage gap. All other things equal, wage gains observed for part-time leavers, therefore, should *overstate* the part-time penalty while wage losses for part-time joiners should *understate* the penalty. As seen in Table 3, wage gaps for leavers and joiners display such a pattern. But the more important finding is that differences in wage gap estimates between leavers and joiners are very small. This indicates that there exists little bias associated with endogenous change in part-time status, and provides us with a rather narrow band between our lower- and upper-bound estimates of the part-time penalty.

In short, longitudinal studies are typically plagued by bias from misclassification (measurement) error and endogenous job change. In our particular application, neither is found to be important. Thus, we have increased confidence in our conclusion that the true part-time wage penalty is quite small.

VII. Additional Evidence: Displaced Workers, Tenure, Students and Nonwage Benefits

This section provides additional analysis on part-time compensation. Below, we provide estimates of the part-time wage gap based on part-time workers displaced from full-time jobs, explore the relationship of experience and tenure with the part-time wage penalty, examine the sensitivity of part-time wage gap estimates to the inclusion and exclusion of students, and summarize evidence on differences in nonwage benefits for part-time and full-time workers.

Part-time Transitions Among Displaced Workers

In the previous section, we concluded that endogenous switching in part-time status appeared to cause little bias in estimates of the part-time wage penalty. This conclusion was based on our expectation from theory that bias from endogenous switching would be in opposite directions for part-time joiners and leavers, coupled with our empirical finding that these lower-bound and upper-bound gaps for joiners and leavers were similar. In this section, we take an alternative approach, estimating wage changes among a sample of workers who are permanently displaced from a full-time job and subsequently take a part-time job (or vice versa).

Gibbons and Katz (1992) have proposed using displaced workers (in particular, those displaced by plant closings) to measure industry wage differentials associated with exogenous job change (i.e., job change *not* related to the wage on either the displacement or post-displacement job). We follow their approach to estimate the wage penalty associated with an exogenous switch from full-time to part-time employment. Note that while the job loss and concomitant job transition is largely exogenous, the choice of subsequent part- or full-time employment is not. Evidence using displaced workers complements previous evidence using matched CPS panels. It does not correspond precisely to the type of conceptual experiment we might like – the random reassignment of full-time workers to part-time jobs, and vice versa.

We use the biennial CPS Displaced Worker Surveys (DWS) for January 1984, 1986, 1988, 1990, and 1992, and February 1994 and 1996 (for description of the DWS, see Farber 1997, 1999). Our sample is restricted to individuals displaced from wage and salary employment owing to a plant or company closing, insufficient work at a job, or a position or shift being abolished, and subsequently being employed in wage and salary employment at the time of the survey. Examined are wage changes between the job from which the worker was displaced and the current job held. The 1984-92 DWS provide information on job displacement over the previous five years, while the 1994 and 1996 DWS provide information on jobs over the previous three years. A limitation of the DWS for 1984-92 is that they provide information on weekly earnings and part-

Both [CPS and DWS] indicate that there is no part-time wage penalty for women and a small penalty for men.

time/full-time status of the displacement (and current) job, but *not* hours worked per week. Hence, precise calculation of hourly earnings is not possible. Usual hours worked on the previous and current jobs are provided in the 1994 and 1996 DWS. In order to retain the sample of displaced workers switching part-time status over all years, we assign hours worked for 1984-92 based on gender-specific mean hours worked on the displacement and current part-time and full-time jobs observed for workers in the 1994-96 surveys. Although this approach introduces considerable *individual* measurement error, coefficients may not be biased if average hours are correct, since measurement error is in the dependent but not independent variables (i.e., on the left-hand-side wage change variable but not the change in part-time status on the right-hand-side).

Table 4 provides results from the DWS. We focus on wage change among displaced workers switching from full-time to part-time status, relative to wage changes among displaced full-time workers who resume full-time employment. It is important that other displaced workers form the reference group, since there exist earnings losses associated with displacement (Fallick 1996, Farber 1997, 1999). Although we also present results for workers displaced from part-time jobs and switching to full-time jobs, we attach less weight to these results, at least for women. This is because displaced female workers who are part-time both in their displacement job and current job fare very well as compared to females who are full-time in both jobs, presumably because such women suffer little loss from accumulated human capital or job seniority. The reason this presents a problem is that to measure the wage differential for those switching from part-time to full-time jobs, we compare their wage change to displaced workers who remained part-time. If women who are part-time in both periods systematically suffer the smallest losses from displacement, their use as the base or comparison group will provide poor estimates of the wage change associated with a change in part-time status.

Line 1 of Table 4 presents regression results that measure the mean log wage change differences among four groups of workers in the DWS, with controls only for year. Relative to displaced full-time female workers subsequently finding full-time employment, full-time women switching to part-time employment suffer a loss of -.053. Among men, the corresponding figure is a loss of -.168. These wage losses from switching into part-time employment are somewhat larger than the losses observed in our matched CPS panels of (largely) non-displaced workers. The estimates decline slightly following control for years since displacement and dummies indicating a geographic move and whether detailed industry and/or occupation changed (line 2). Following full control for the change in occupational characteristics and the change in broad industry and occupation (line 3), wage change from a switch to part-time employment is effec-

Part-time workers generally will accumulate less human capital than full-time workers for equivalent years of potential experience.

tively zero for women (a *positive* and insignificant .011) and a nontrivial -.079 for men.

The evidence from the DWS on full-time to part-time switchers is consistent with the CPS longitudinal results. Both indicate that there is no part-time wage penalty for women and a small penalty for men. In short, much of the observed cross-sectional part-time wage gap is the result of measured and unmeasured worker-specific skills and job characteristics.[15]

Surprisingly, we observe no gain in hourly earnings among women displaced from part-time jobs who subsequently work in full-time jobs (relative to wage change for workers who remain part-time) — the coefficient for part-time leavers being -.034 log points absent controls and -.093 after accounting fully for job characteristics. Taken literally, these estimates suggest a *full-time* penalty for this group of workers. We believe these results are misleading since the wage loss for those switching from part- to full-time jobs is measured relative to a large .10 log point wage gain for the small sample of part-time stayers (or more precisely, a .10 gain relative to the losses realized by full-time stayers). Results among men switching to full-time jobs are fully consistent with expectations and prior evidence — a wage gain of .089 log points absent controls and a small and insignificant .022 following controls.

Despite some ambiguity in the DWS evidence, the results reinforce our previous conclusions from the CPS longitudinal analysis that the bias from endogenous switching is small, there exists little or no part-time wage penalty among women, and there is a modest but very real part-time penalty among men.

Are Unmeasured Skills Experience and Tenure Related?

Our longitudinal analysis has indicated that much of the observed part-time wage disadvantage is associated with worker-specific skills unmeasured in standard analysis. Part-time workers generally will accumulate less human capital than full-time workers for equivalent years of potential experience. Moreover, as shown by Blank (forthcoming) using the PSID, individuals' hours worked are correlated across time. Hence, part of what is typically interpreted as a part-time penalty reflects lower levels of accumulated human capital owing to fewer hours of *prior* work (i.e., a higher frequency and duration of part-time spells).

To examine this issue explicitly, we first return to the wage level analysis presented previously in Table 2. We estimate equations identical to those shown in lines 3 and 6, except that we allow the slopes of wage-experience profiles to vary by part-time status. As expected, we find a significantly flatter profile, or slower

[W]hat is typically interpreted as a part-time penalty reflects lower levels of accumulated human capital owing to fewer hours of prior work.

wage growth, for part-time than full-time workers. Stated alternatively, the part-time wage differential grows with respect to potential experience.[16]

By restricting wage equations to have common slopes of earnings profiles for part- and full-time workers, as we did previously (see Table 2), differences in accumulated skills are then reflected in the part-time coefficient, overstating the true part-time penalty. An alternative approach is to assume that *all* differences in the slopes of experience profiles between part- and full-time workers result from differences in human capital accumulation, rather than including it as part of the part-time penalty estimate.

Following this alternative approach, we find that differences in profile slopes account for a large share of the part-time/full-time differential. The part-time penalty is measured by the coefficient on *PT* in a specification that also includes interactions between *PT* and potential experience (and its square). Our estimate of θ in the full wage level model (line 6 of Table 2) is -.109 for women and -.200 for men absent the interaction terms. Once one adds the interaction terms, the *PT* coefficients change to -.067 for women and -.122 for men. These estimates of θ are highly similar to longitudinal estimates of θ' based on industry and occupation changers, shown on line D-2 of Table 3 (-.069 and -.105, respectively). In short, a substantial portion of the part-time wage disadvantage is associated with lower worker-specific skills, and lower skills among part-time workers appear to be due in large part to lower levels of human capital among part-time workers accumulated over the work life.

An alternative, albeit imperfect, proxy for accumulated work experience and on-the-job training is tenure, which measures years with one's current employer (for a recent review of literature on wages and tenure, see Farber 1999). We utilize the April 1993 CPS Benefit Supplement (roughly in the middle of our 1989-97 sample period), which contains data on tenure.

We find predictable differences in mean tenure for part-time and full-time workers. Among women, part-time workers average only 4.5 years with their current employer, as opposed to 7.5 years for full-time workers. Among men, the absolute and relative tenure gap is larger, part-time and full-time workers average 3.2 and 8.8 years of tenure, respectively. Because part-time and full-time workers can differ in age and other characteristics, we estimate log tenure equations, holding constant years of potential experience and other worker characteristics. As seen in Table 5, among part-time women, tenure is -.46 (37%) lower than for full-time workers, controlling for potential experience and its square, and is -.31 (27%) lower following control for individual characteristics, location, union membership, firm size, industry and occupation dummies (absent firm size dummies the

[A] substantial portion of the part-time wage disadvantage is associated with lower worker-specific skills, and lower skills among part-time workers appear to be due in large part to lower levels of human capital among part-time workers accumulated over the work life.

part-time coefficient is -.32). An even stronger pattern is found among men, with a part-time coefficient of -.60 (45%) controlling for experience and -.40 (33%) following a full set of controls (or -.41 absent firm size).

The tenure evidence demonstrates that part-time workers possess substantially lower firm-specific experience than do full-time workers with similar characteristics in similar industries and occupations. This finding is consistent with our previous conclusion that measured and unmeasured worker skill differences account for a substantial portion of the part-time wage disadvantage. That being said, tenure *per se* does not appear to be the sole or even primary explanation for low part-time wages. When we add tenure and its square to a fairly dense specification of a standard log wage equation, the part-time gap declines substantially, but remains sizable. As seen in Table 5, adding tenure reduces the estimated part-time disadvantage for women from -.113 to -.081. Among men, the wage gap without control for tenure is -.202, falling to -.171 following control. The larger relative impact of tenure on the female than male part-time coefficient is because tenure partly captures the effects of general work experience, for which the standard potential experience measure (Age minus Schooling minus 6) is a poorer proxy for true experience among women than among men (for evidence, see O'Neill and Polachek 1993).

Students

The analysis presented to this point in our study has included workers enrolled in school. Among those in the labor force, school attendance is likely to increase the labor supply of part-time relative to full-time workers, particularly in jobs most complementary to students' schedules and preferences. Moreover, students are often concentrated in labor markets where the supply (and, to a lesser extent, demand) for part-time work is large (i.e., so-called college towns).

Our principal interest here is not the magnitude of the part-time penalty for students, but rather, whether our previous estimates would have differed appreciably had the analysis been restricted to non-students. To answer this question, we first re-estimate the wage level and wage change equations shown in Tables 2 and 3, omitting all persons who were full-time students in either year *t* or *t-1* (school attendance questions are asked only of respondents ages 16-24). Omission of students reduces by several percentage points both the female and male *unadjusted* part-time differential, shown in line A-1 of Table 2. This reduction reflects the fact that, absent controls, one is comparing many younger part-time workers to older full-time workers. In wage level regressions with controls (lines A-2 through A-6), the effect is minor, decreasing the female part-time wage gap by about 1% and increasing the male gap by about 1%. In the wage change equations (Table 3), longitudinal estimates of the part-time gap were reduced by about 1 percentage point for both women and men following exclusion of students.

Among women, part-time workers average only 4.5 years with their current employer, as opposed to 7.5 years for full-time workers.

Young people attending school constitute an important segment of the part-time labor force. But because they face severe constraints on their time and geographic mobility, part-time wages realized by students need not be representative of part-time wages available to nonstudents. Our evidence suggests that the part-time wage gap faced by the larger population of nonstudents is slightly smaller than that previously shown for the combined students and non-student population. But none of the conclusions reached in the paper would be appreciably different were the analysis restricted to the nonstudent population.

Nonwage Benefits

The analysis in this paper has focused exclusively on *wage* differences between part-time and full-time workers. Even if there were no part-time wage penalty, however, there can exist a *compensation* penalty owing to differences in nonwage benefits. Available data do not readily permit incorporation of benefits into a compensation equation. For example, CPS benefit data providing information on *receipt* of fringes by individual workers allow measurement of benefit coverage, but not the dollar value of benefits. Establishment surveys conducted by BLS provide information on costs of nonwage benefits to employers, but do not permit matching these benefits to individual workers for whom we have measures of age, schooling and other wage-related characteristics.

All available evidence indicates that part-time workers receive substantially lower nonwage benefits than do full-time workers. The March CPS provides data on pension coverage and health insurance for the previous calendar year. Snider (1995) compares coverage rates for part-time and full-time employees using the March 1993 CPS. In 1992, 57.9% of full-time workers were employed in firms offering pension plans, of which 83.0% chose to participate, leading to a 48.0% rate of pension coverage. Among part-time workers 30.1% were employed in firms with pension plans and 38.0% participated, resulting in a coverage rate of only 11.4%. Both offer and final coverage rates increase with respect to firm size, but full-time rates greatly exceed part-time rates among workers within each firm size category (take-up rates vary little with respect to firm size). The March CPS figures indicate that part-time

[T]enure evidence demonstrates that part-time workers possess substantially lower firm-specific experience than do full-time workers with similar characteristics in similar industries and occupations.

workers have low pension take-up rates when employed in firms with pension plans. This is somewhat misleading, since firms that offer pensions need not offer them to part-time workers. Hence, the low part-time take-up rates presented by Snider reflect both the choices of part-time workers and their employers.[17]

Snider (1995) also summarizes data on health insurance coverage. Among full-time workers during 1992, 61.2% received health insurance coverage directly from their employer, as compared to only 16.4% of part-time workers. Although this is the relevant figure for employer coverage of their own workers, it greatly understates economy-wide coverage. Snider calculates that more than a third of part-time workers are indirectly covered as dependents on other workers' employer plans (indirect coverage is substantially higher for voluntary than for involuntary part-time workers). In total, 52.0% of part-time workers are covered (directly or indirectly) by an employer health plan, as compared to 73.3% among full-time workers. Adding in public and other private sources of coverage, she finds that 79.5% of part-time workers have health insurance coverage, as compared to 84.1% of full-time workers.

Farber and Levy (1998) provide a comprehensive analysis of changes in health insurance coverage over time, using various CPS benefit and tenure supplements from 1979 through 1997. They are particularly interested in understanding recent declines in coverage, distinguishing among full-time jobs and workers with at least a year of tenure, full-time jobs where tenure is less than a year, and part-time jobs. They conclude that the decline in insurance coverage among full-time core workers has resulted from a decline in take-up rates, and not offer or eligibility rates. By contrast, they find evidence that employers have decreased insurance eligibility or offer rates for short-term and part-time employees, with most of the change occurring between the May 1988 and April 1993 CPS benefit surveys.

The relatively low pension take-up rate among part-time workers, coupled with previous evidence that many part-time workers are covered by health insurance through someone other than their employer, is suggestive. A possible inference is that many part-time workers place a low value on pension and health insurance benefits and that low coverage rates among part-time workers may not constitute so serious a social problem as suggested by their low coverage rates. Even were this inference correct, however, the evidence points to *compensation* differentials between part-time and full-time workers more sizable than are *wage* differentials.

Additional evidence is available from the BLS's *Employee Benefits Survey* of medium and large size private establishments of 100 or more workers (BLS 1997) and their survey of small private establishments of fewer than 100 workers (BLS 1998a). The medium and large establishment survey finds direct employer health coverage of 19% for part-time versus 77% for full-time workers in 1995 (BLS 1997, Tables 1, 4). Coverage rates among small establishments for 1996 indicate rates of 6% and 64% for part- and full-time workers, respectively (BLS 1998a, Tables 1, 2). The BLS Surveys also provide pension plan coverage rates. These figures differ rather substantially from Snider's calculations from the CPS, although evidence for a large part-time gap is similar. Among medium and large establishments an 80% coverage rate is found for full-time workers, compared to 37% among part-time workers (BLS 1997, Tables 1, 4). For small establishments, the corresponding rates are 46% and 13% (BLS 1998a, Tables 1, 2). The BLS surveys also offer evidence on paid time off (holiday, vacation, personal, funeral, jury, military, sick and family leave), other forms of insurance (disability, dental and life), tax-deferred earnings arrangements and details on retirement plans (whether a defined benefit and the type of defined contribution plan). Differences between part-time and full-time workers in all benefit categories are substantial.

Differences in coverage rates for nonwage benefits tell us little about *cost* differences or the magnitude of the part-time compensation gap. The BLS's *Employment Cost Index* (ECI) program does provide information on costs by establishment, although such data cannot be linked to individual workers and their characteristics. Data from the ECI (BLS 1998b, Table 9) indicate that in March 1998 total compensation costs for full-time employees were roughly double that for part-time employees ($20.95 versus $10.01 per hour), with nonwage benefits accounting for 28.3% and 19.0% of total compensation costs for full-time and part-time workers, respectively. Legally required benefits constitute 57% of all part-time benefits ($1.08 out of $1.90 per hour) as compared to 30% of those for full-time workers ($1.78 out of $5.93).

In a paper using March 1994 ECI data, BLS economist Michael Lettau (1997) calculates part-time wage and compensation differentials, based on 567 part-time and 571 full-time observations on jobs within the same establishment and 3-digit occupation. Lettau finds a -.164 part-time log wage gap (within establishment and occupation) for private nonunion jobs and a -.227 log compensation differential. The log benefit gap is estimated to be -.475. Although Lettau is comparing part-time and full-time workers in the same occupation and establishment, he cannot control for individual differences in schooling, age or experience, tenure, gender, race and the like between part-time and full-time workers within these jobs. The evidence in this report shows that the part-time wage gap is accounted for in no small part by accumulated training and other individual differences in worker skills. This leads us to believe that the true part-time wage penalty is substantially less than the -.164 gap found by Lettau.

Of greater interest to us is the .06 change in the part-time gap found by Lettau following addition of nonwage benefits, from -.164 to -.227. The .06 change *may* provide an *upper-bound* approximation of how our gap estimates would change were we able to measure the cost of nonwage benefit received by individual workers in the CPS. It should be an upper-bound estimate because benefits increase with skill level, and analysis with the ECI does not control fully for worker-specific skills. Just as part-time *wage* penalty estimates fall sharply as one controls for skill, *benefit* penalty estimates should do so as well. We have previously concluded that average part-time/full-time *wage* gaps for similar workers in similar jobs are approximately zero among women and very small for men. If we add to these wage gap estimates the "upper-bound" .06 difference between the wage and compensation gap, our best estimates of part-time *compensation* penalties are now nontrivial in magnitude, yet still far smaller than implied by standard estimates.

VIII. Conclusions

In this paper we have examined the role of worker-specific skills, occupational skill requirements, and job working conditions on what are large differences in wages between part-time and full-time workers. The analysis shows that roughly two-thirds of the part-time wage disadvantage for women and men can be accounted for by measurable differences in workers and jobs. Much of the remaining differential reflects unmeasured worker-specific skills and tastes, as captured through longitudinal analysis that measures wage changes among individual workers switching from part-time to full-time employment, or vice versa.

Our "preferred" longitudinal estimates of part-time wage gaps indicate little if any part-time penalty for women, who comprise two-thirds of part-time workers, and a modest penalty among men. Substantial longitudinal wage gaps are found only for workers changing detailed industry and occupation in addition to part-time status, and for male workers displaced from full-time jobs who eventually take part-time jobs. Unmeasured worker-specific skill differences between part-time and full-time workers appear to stem primarily from accumulated human capital associated with work experience. Neither misclassification error nor endogeneity bias, either of which could bias longitudinal estimates, appears to be of consequence in this analysis.

It is widely acknowledged that employer-fixed costs and the large number of workers preferring to work part-time lead to lower compensation for part-time than for full-time workers. But the analysis here demonstrates that most of the rather sizable part-time *wage* disadvantage results from differences between part- and full-time workers in job characteristics, preferences and, most importantly, accumulated worker skills. For similar workers in otherwise similar jobs, part-time wage penalties are very small, leading to what appears to be a modest gap in total compensation. Nothing in our analysis is inconsistent with the expectation from economic theory that compensation is determined largely through the interaction of labor demand and supply and that outcomes approximate those to be expected in relatively competitive markets.

Data Appendix:

Construction of the Longitudinal Database from the CPS ORG Files

The longitudinal CPS database was created from the CPS Outgoing Rotation Group (ORG) Earnings Files for 1989-97 in the following manner. Households are included in the CPS for 8 months — 4 consecutive months in the survey, followed by 8 months out, followed by 4 months in. Outgoing rotation groups 4 and 8 are asked earnings supplement questions (weekly earnings, hours, union status, etc.). The CPS contains household identification numbers (ID) and record line numbers, but not individual identifiers until 1994). Individuals potentially can be identified for the same month in consecutive years; that is, individuals in rotation 4 in year 1 can be matched to individuals in rotation 8 in year 2.

Separate data files were created for males and females, and for pairs of years. Within each file, individuals were sorted as appropriate on the basis of ascending and descending household ID, year and age. To be considered an acceptable matched pair, a rotation 8 individual had to be matched with a rotation 4 individual with identical household ID, identical survey month, and an age difference between 0 and 2 (since surveys can occur on different days of the month, age change need not equal 1). Several passes were necessary because a single household may contain more than one male or female pair. Checks were provided to ensure that only unique matches were selected. For each rotation 8 individual, the search was made through all rotation 4 individuals with the same ID to make sure there was only 1 possible match; the file was resorted in reverse order and each selected rotation 4 individual was checked to ensure a unique rotation 8 match. Incorrect changes in the variables marital status, veteran status, race and education (e.g., a change in schooling other than 0 or 1, a change from married to never married, etc.) were used to delete "bad" observations in households where there were multiple observations and ages too close to separate matched pairs. Several passes at the data were made. In households where two pairs of individuals could be separated based on a 1-year but not the 0-to-2-year age change, a 1-year criterion was used. If a unique pair could not be identified based on these criteria, they were not included in the data set. For years since 1994, CPS individual identifiers were used as the principal criterion to select matches.

The principal reasons that matches cannot be made or for exclusion from the longitudinal sample are if a household moves, if an individual moves out of a household, if a worker becomes self-employed, if the Census is unable to reinterview a household and/or receive information on the individual, or if an individual drops out of the labor market or fails to meet other sample selection criteria (see discussion in the text). Peracchi and Welch (1995) analyze attrition rates among matched March CPS files and conclude that age is the most important determinant of a successful match. Other factors that lessen match probabilities are poor health, low schooling, and not a household head, while sex and race are unimportant predictors following control for other factors.

Endnotes

[1] Public discussion of part-time work has become entangled with discussion of firms' use of contingent or temporary workers. For descriptive evidence and analysis of contingent workers, see Segal and Sullivan (1997) and articles in the October 1996 *Monthly Labor Review* (e.g., Polivka 1996a, 1996b). Newspaper articles include Uchitelle (1997), who links the part-time issue at UPS to what is occurring in the larger labor market, and Phillips (1997), who questions whether the part-time issue warrants such attention. Nor has organized labor or Washington-based policy organizations been silent on the issue. The AFL-CIO (1998) provides strong criticism of employer use of and compensation for part-time employees. Similarly, Mishel and Bernstein (1994) express concerns about part-time jobs and the growth of a contingent work force. In contrast, Lyons (1997) concludes, following a review of published data and academic studies, that part-time employment should raise few concerns.

[2] A part-time wage differential can also arise if workers are willing to take low paid part-time jobs as a way of bidding or queueing for full-time jobs with efficiency wages or other forms of rents. We doubt that a substantial share of part-time jobs are of this sort.

[3] This argument is reinforced by the Fair Labor Standards Act (FLSA) mandate of an overtime premium for hourly workers, which raises the cost of varying hours among full-time workers.

[4] There are other recent articles on part-time employment less closely related to our work. Montgomery and Cosgrove (1993) find that childcare establishments reduce the use of part-time workers in response to higher nonwage benefits. Stratton (1994) finds that official statistics overstate the level of "involuntary" part-time workers. Elsewhere Stratton (1996) concludes that most part-time workers categorized as involuntary are likely to be such and more likely than other part-timers to switch to full-time employment. Fallick (1998) examines the relationship between changes in industry growth and changes in part-time employment. Papers cited in the text provide references to a limited number of earlier studies.

[5] Details on construction of the CPS panel are provided in the Data Appendix. The Census ended one set of area samples after May 1995 and adopted new area samples beginning September 1995, not permitting the matching of households across years for June 1994/95 through August 1995/96.

[6] The same three patterns are evident when we match to the CPS occupational skill and working condition measures from the *Dictionary of Occupational Titles*, a predecessor to *O*NET*.

[7] There are also significant numbers of workers at hour intervals divisible by eight – 24, 32, and 48. This raises an issue not addressed in this report. Is it low weekly hours or daily hours that are associated with lower part-time wages? That is, does a part-time worker with four eight-hour shifts make more or less than a worker with 32 hours spread over five days? The limited evidence available in CPS data suggests that part-time workers with eight-hour shifts make more. This is consistent with the conclusion reached by Lettau (1997), who is able to examine this issue directly using data on length of shift included in the Employment Cost Index (ECI).

[8] These figures are based on data since 1994. Prior to 1994, the hours worked last week variable applies to *all* jobs rather than just the primary job. Usual weekly earnings and usual weekly hours both apply to the primary job. Beginning in 1994, all earnings and hours measures are provided separately for the primary and other jobs. A smoother distribution of hours worked is also found when using time diary surveys where individuals report their activities for each 15 minute block of time.

[9] For a careful analysis of this issue, albeit with a far smaller data set, see Averett and Hotchkiss (1996).

[10] Mean wages tend to be higher at the rounded hour intervals than at other hour intervals. We do not have a convincing explanation for this finding.

[11] Letting θ be the part-time log wage differential, the percentage differential is approximated by $100[\exp(\theta)-1]$. For relatively small values of θ or θ' (about .10 or less) the log differential is nearly identical to the percentage differential. As values of θ or θ' move further away from zero the divergence is greater. Log differentials are invariant to the base, whereas percentage differentials are not. For example, the log differential $\theta = -.20$ implies that part-time workers have hourly pay -.20 log points or 18.1% lower than do full-time workers. This is absolutely equivalent to saying that full-time workers have a +.20 log point or 22.1% pay advantage as compared to part-time workers.

[12] Rather than include a dummy coded 1 for part-time stayers, a "PT-Initial Period" dummy, coded 1 for all those part-time in year *t-1* is included. Its coefficient measures wage change for part-time stayers relative to full-time stayers, while the part-time leaver dummy now measures the wage gain from switching to a full-time job, relative to remaining part-time.

[13] Separate 0/1 dummies are included in line D-1, controlling for whether workers changed detailed industry and occupation, industry only and occupation only (these coefficients are effectively zero). Line D-2 adds, among other things, variables measuring the change in broad industry and occupation (i.e., industry and occupation dummies in differenced form, with values of -1, 0, and 1).

[14] For such evidence with respect to longitudinal union premium estimates, see Hirsch and Schumacher (1998).

[15] It is not clear whether estimates from line 2 or line 3 are preferable. The occupational and industry variables included in line 3 control for the fact that many displaced workers gain part-time employment in lower paying occupations and industries. Thus, the lower wages associated with jobs in those industries and occupations are not included as part of the part-time penalty in line 3, whereas the part-time penalty estimate in line 2 includes such factors. It is likely that the specification in line 2 controls for too little and that in line 3 for too much.

[16] This relationship need not hold for very young workers. In work not shown, a *larger* part-time differential for very young (less than age 25) females than for older females is found. The opposite pattern is found for males.

[17] That is, reported offer rates overstate the rate at which part-time workers are offered benefits, and reported take-up rates understate the rate at which part-time workers offered benefits take them. The final coverage rate, which represents the product of the two, should be correct.

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Table 1: Selected Variable Means for Part-Time and Full-Time Female and Male Workers

	Females Part-Time	Females Full-Time	Males Part-Time	Males Full-Time
Individual Characteristics:				
Usual Hours per Week	22.009	40.882	21.366	43.250
Hourly Earnings (1997$)	9.911	12.753	9.353	16.059
Years Schooling	12.981	13.506	12.525	13.334
Age less than 25	0.193	0.070	0.439	0.071
Age 25 to 59	0.681	0.880	0.349	0.875
Age 60 and over	0.126	0.050	0.212	0.053
Married, Spouse Present	0.624	0.598	0.364	0.722
Separated, Divorced, Widowed	0.143	0.212	0.074	0.094
Children in Household	1.034	0.777	0.671	0.957
Hispanic	0.045	0.059	0.067	0.070
Non-Hispanic Black	0.062	0.114	0.074	0.074
Asian, Pacific Islander	0.022	0.032	0.037	0.029
Other Race	0.006	0.009	0.010	0.008
Union Member	0.090	0.169	0.090	0.231
Metropolitan Area over 2.5 million	0.349	0.370	0.356	0.362
O*NET Aggregate Occupational Skills:				
Verbal Aptitude	13.256	14.448	11.393	13.106
Mathematical Aptitude	7.500	8.394	6.787	8.087
Spatial Aptitude	3.534	3.568	3.903	4.361
Problem Solving Skills	18.338	21.737	16.497	21.085
Technical Skills	13.573	15.717	16.752	21.849
System Skills	10.851	13.811	10.138	14.357
O*NET Aggregate Occupational Working Conditions:				
Strength	6.485	5.714	8.834	8.217
Environmental Exposure	3.287	3.092	5.146	5.429
Hazards	2.190	2.023	3.692	4.582
Protective Attire	0.912	0.877	1.358	1.679
Sample Size	41,690	135,223	13,852	165,790

The sample includes wage and salary workers ages 16 and older employed during consecutive years, from matched panels of the CPS Outgoing Rotation Group (ORG) earnings files for 1989/90 through 1996/1997. Deleted are workers not matched across years and those who in either year have top-coded earnings, allocated hours worked per week (in years where possible), or extreme wages or wage changes. Means are based on the second year observation for each worker, years 1990-97. Each of the O*NET occupational variables, matched to individuals in the CPS, are the aggregate of several more detailed O*NET variables, as shown in Appendix Table A-1. The detailed rather than aggregate O*NET variables are included in the regressions. Part-time status is determined by whether individuals' reported usual hours worked per week on their principal job are less than 35 hours.

Table 2: Wage Level Estimates of the Part-Time/ Full-Time Log Wage Differential

	Females		Males	
	Coefficient	(s.e.)	Coefficient	(s.e.)
A. Wage Level Equations (PT coefficients)				
1. No Controls	-.300	(.0028)	-.620	(.0045)
2. Personal & Location Variables	-.201	(.0024)	-.323	(.0041)
3. 2 plus Industry Dummies	-.155	(.0024)	-.263	(.0040)
4. 3 plus Occupation Dummies	-.126	(.0023)	-.222	(.0039)
5. 3 plus O*NET Job Variables	-.110	(.0023)	-.205	(.0039)
6. 3 plus O*NET and Occ Dummies	-.109	(.0023)	-.200	(.0039)
B. Wage Level Equations (Hours Range coefficients)				
1. No Controls				
Part-Time:				
1-24 hours	-.289	(.0035)	-.634	(.0057)
25-34 hours	-.265	(.0041)	-.569	(.0072)
Full-Time:				
35-39 hours	-.021	(.0038)	-.106	(.0063)
40 hours	—	—	—	—
41-49 hours	.156	(.0052)	.083	(.0040)
50+ hours	.124	(.0052)	.028	(.0033)
2. Full Controls				
Part-Time:				
1-24 hours	-.113	(.0029)	-.209	(.0048)
25-34 hours	-.107	(.0032)	-.214	(.0056)
Full-Time:				
35-39 hours	.000	(.0029)	-.037	(.0048)
40 hours	—	—	—	—
41-49 hours	.047	(.0039)	.022	(.0030)
50+ hours	-.069	(.0040)	-.065	(.0026)

The sample includes wage and salary workers ages 16 and older employed during consecutive years, from matched panels of the CPS Outgoing Rotation Group (ORG) earnings files for 1989/90 through 1996/1997. Deleted are workers not matched across years and those who in either year have top-coded earnings, allocated hours worked per week (in years where possible), or extreme wages or wage changes. Wage level estimates above are based on second year observations for the years 1990-97. Control variables included in the full wage level regressions (lines A-6 and B-2) are schooling, potential experience (minimum of age-schooling-6 or age-16) and its square, marital status (2 dummies included), children (3), race and ethnicity identifiers (4), union membership, region (8), metropolitan size (6), industry (12), occupation (11), year (7) and all O*NET variables shown in Appendix Table A-1.

Table 3: Wage Change Estimates of the Part-Time/ Full-Time Log Wage Differential

	Females		Males	
	Coefficient	(s.e.)	Coefficient	(s.e.)
A. Wage Change Equations (ΔPT coefficients)				
1. No Controls	-.029	(.0028)	-.069	(.0041)
2. Full Controls	-.014	(.0028)	-.043	(.0041)
B. Wage Change Equations (asymmetrical change coefficients)				
1. No Controls				
FT to PT (PT joiners)	-.017	(.0043)	-.051	(.0064)
PT to FT (PT leavers)	.038	(.0042)	.069	(.0067)
PT initial period	-.000	(.0023)	.014	(.0040)
FT stayers	—	—	—	—
F test: PT leave = -PT join	12.027	(.0005)	3.569	(.0589)
(Prob > F)				
2. Full Controls				
FT to PT (PT joiners)	-.013	(.0043)	-.042	(.0064)
PT to FT (PT leavers)	.020	(.0042)	.044	(.0067)
PT initial period	-.001	(.0023)	.006	(.0040)
FT stayers	—	—	—	—
F test: PT leave = -PT join	1.635	(.2010)	0.033	(.8563)
(Prob>F)				
C. Wage Change Equations (ΔPT interacted with dummies for absolute change in usual hours worked)				
1. No Controls				
1-4 ΔHours * ΔPT	-.014	(.0107)	-.009	(.0263)
5-9 ΔHours * ΔPT	-.001	(.0057)	-.034	(.0101)
10-14 ΔHours * ΔPT	-.034	(.0059)	-.069	(.0086)
15-19 ΔHours * ΔPT	-.036	(.0065)	-.118	(.0098)
20+ ΔHours * ΔPT	-.045	(.0050)	-.065	(.0065)
2. Full Controls				
1-4 ΔHours * ΔPT	-.013	(.0106)	-.003	(.0261)
5-9 ΔHours * ΔPT	.004	(.0057)	-.024	(.0100)
10-14 ΔHours * ΔPT	-.020	(.0059)	-.051	(.0086)
15-19 ΔHours * ΔPT	-.020	(.0065)	-.090	(.0097)
20+ ΔHours * ΔPT	-.019	(.0050)	-.029	(.0065)

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Table 3 Continued: Wage Change Estimates of the Part-Time/Full-Time Log Wage Differential

	Females Coefficient	Females (s.e.)	Males Coefficient	Males (s.e.)
D. Wage Change Equations (ΔPT coefficients for each Ind/Occ Change group)				
1. No Controls (includes separate change dummies)				
Ind/Occ Change*ΔPT	-.100	(.0047)	-.148	(.0061)
Ind Only Change*ΔPT	-.011	(.0104)	-.027	(.0160)
Occ Only Change*ΔPT	-.010	(.0060)	-.026	(.0095)
No Ind/Occ Change*ΔPT	.025	(.0045)	.016	(.0076)
2. Full Controls				
Ind/Occ Change*ΔPT	-.069	(.0047)	-.105	(.0062)
Ind Only Change*ΔPT	-.009	(.0103)	-.022	(.0159)
Occ Only Change*ΔPT	.000	(.0059)	-.010	(.0094)
No Ind/Occ Change*ΔPT	.027	(.0045)	.021	(.0076)

	Females	Males
Sample Sizes:		
Full Wage Level and Panel Samples	176,913	179,642
PT Changers	18,432	9,023
FT to PT (PT joiners)	8,131	3,793
PT to FT (PT leavers)	10,301	5,230
1-4 Hours	1,215	221
5-9 Hours	4,272	1,495
10-14 Hours	4,033	2,070
15-19 Hours	3,288	1,601
20+ Hours	5,624	3,636
Ind/Occ Change	6,407	4,082
Ind Only Change	1,296	597
Occ Only Change	3,912	1,700
No Ind/Occ Change	6,817	2,644

The sample includes wage and salary workers ages 16 and older employed during consecutive years, from matched panels of the CPS Outgoing Rotation Group (ORG) earnings files for 1989/90 through 1996/1997. Deleted are workers not matched across years and those who in either year have top-coded earnings, allocated hours worked per week (in years where possible), or extreme wages or wage changes. The change equations are estimated over worker-year pairs for 1989/90-1996/97. Controls included in the full wage change equations, in addition to the change in part-time status variables, are changes in experience squared, change in union membership (interacted with Ind/Occ change status), change in broad industry (12), change in broad occupation (11), dummies for a change in detailed Ind/Occ (3), year dummies (7) and change in all O*NET variables shown in Appendix Table A-1.

Table 4: Wage Change from Part-Time Transitions Among Displaced Workers

	Females		Males	
	Coefficient	(s.e.)	Coefficient	(s.e.)
DWS Wage Change Equations				
1. Year Controls Only				
FT to PT (PT joiners)	-.053	(.0165)	-.168	(.0178)
PT to FT (PT leavers)	-.034	(.0276)	.089	(.0453)
PT initial period	.079	(.0207)	-.031	(.0380)
FT stayers	—	—	—	—
2. Plus Years Since Displacement, Relocation, and Ind/Occ Switch Controls				
FT to PT (PT joiners)	-.042	(.0164)	-.140	(.0178)
PT to FT (PT leavers)	-.044	(.0273)	.084	(.0449)
PT initial period	.098	(.0206)	-.005	(.0377)
FT stayers	—	—	—	—
3. Plus Changes in Broad Ind/Occ and O*NET Variables				
FT to PT (PT joiners)	.011	(.0163)	-.079	(.0176)
PT to FT (PT leavers)	-.093	(.0269)	.022	(.0438)
PT initial period	.102	(.0202)	-.002	(.0367)
FT stayers	—	—	—	—
Sample Sizes	7,793		11,750	

Dependent variable is change in log of hourly earnings between the displacement job and current job. Data source is biennial CPS Displaced Worker Surveys (DWS) for 1984 through 1996. The sample includes workers displaced from wage and salary employment owing to a plant or company closing, insufficient work at a job, or a position or shift being abolished, and subsequently being employed in wage and salary employment at the time of the survey. For the 1994 and 1996 DWS hourly earnings are calculated by usual weekly earnings divided by usual weekly hours. For the 1984-92 DWS, hourly earnings is calculated based on weekly earnings and gender-specific mean hours worked on the displacement and current part-time and full-time jobs observed for workers in the 1994-96 surveys. Line (1) includes the transition variables shown, plus dummies for survey year. Line (2) adds the years since displacement, a dummy for geographic relocation, and dummies for changing detailed industry and occupation, industry only, and occupation only. Line (3) further adds dummies measuring changes in broad occupation and industry (with values of -1, 0, and 1) and the change in O*NET occupational skill and working condition variables listed in Appendix Table A-1.

Table 5: Tenure Differences Between Part-Time and Full-Time Workers

	Females		Males	
A. Mean Years of Tenure and Age (s.d.)				
Part-Time Workers:				
Tenure	4.5	(5.0)	3.2	(4.8)
Age	37.7	(11.8)	33.3	(12.9)
Full-Time Workers:				
Tenure	7.5	(7.0)	8.8	(8.5)
Age	38.7	(10.6)	38.8	(10.7)
B. Part-Time Coefficients (s.e.) in Log(Tenure) Equations with Alternative Sets of Control Variables:				
1. Experience Controls Only	-.463	(.0229)	-.599	(.0401)
2. Individual and Location	-.452	(.0232)	-.551	(.0401)
3. Individual, Location, and Job	-.311	(.0236)	-.396	(.0390)
C. Part-Time Coefficients (s.e.) in Log(Wage) Equations with and without Tenure Controls:				
1. Standard without Tenure	-.113	(.0101)	-.202	(.0181)
2. Standard with Tenure	-.081	(.0099)	-.171	(.0178)
Sample Sizes	9,144		9,494	

Data source is April 1993 CPS Benefit Supplement. Sample includes wage and salary workers ages 20-64. Line B Log(Tenure) regressions: (1) includes part-time dummy, potential experience, and experience squared; (2) adds years of schooling, number of children, and dummies for race, marital status, region, and large metropolitan area; (3) adds firm size dummies, union status, and industry and occupation dummies. Line C Log(Wage) regressions: (1) includes part-time dummy, years of schooling, experience and its square, number of children, and dummies for race, marital status, region, large metropolitan area, union status, and industry and occupation dummies; (2) adds tenure and tenure squared.

Appendix Table A-1: O*NET Variable Means for Part-Time and Full-Time Female and Male Workers

	Females		Males	
	Part-Time	Full-Time	Part-Time	Full-Time
Occupational Skill Characteristics:				
Verbal Aptitude	13.256	14.448	11.393	13.106
Oral Comprehension	3.485	3.699	3.070	3.386
Written Comprehension	3.299	3.685	2.883	3.408
Oral Expression	3.594	3.764	3.053	3.414
Written Expression	2.878	3.300	2.387	2.898
Mathematical Aptitude	7.500	8.394	6.787	8.087
Mathematical Reasoning	2.207	2.506	1.924	2.371
Number Facility	2.703	2.945	2.424	2.766
Mathematics	2.590	2.943	2.438	2.950
Spatial Aptitude	3.534	3.568	3.903	4.361
Spatial Orientation	1.606	1.518	1.764	1.821
Visualization	1.928	2.050	2.139	2.540
Problem Solving Skills	18.338	21.737	16.497	21.085
Problem Identification	2.829	3.190	2.585	3.196
Information Gathering	2.644	3.131	2.391	2.992
Information Organization	2.691	3.063	2.338	2.721
Synthesis/Reorganization	1.854	2.273	1.540	2.040
Idea Generation	2.031	2.450	1.912	2.506
Idea Evaluation	2.103	2.529	1.946	2.545
Implementation Planning	1.841	2.343	1.664	2.364
Solution Appraisal	2.344	2.758	2.121	2.722
Technical Skills	13.573	15.717	16.752	21.849
Operations Analysis	1.351	1.802	1.368	2.082
Technology Design	0.959	1.092	1.067	1.486
Equipment Selection	1.913	2.150	2.122	2.585
Installation	0.714	0.849	1.164	1.579
Programming	0.237	0.420	0.216	0.460
Testing	0.759	1.025	0.994	1.644
Operation Monitoring	1.213	1.292	1.449	1.805
Operation and Control	1.782	1.894	2.049	2.293
Product Inspection	2.060	2.339	2.193	2.682
Equipment Maintenance	0.907	0.969	1.437	1.737
Troubleshooting	0.942	1.103	1.370	1.882
Repairing	0.737	0.783	1.322	1.613

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Appendix Table A-1 Continued: O*NET Variable Means

	Females		Males	
	Part-Time	Full-Time	Part-Time	Full-Time
System Skills	10.851	13.811	10.138	14.357
Visioning	1.733	2.232	1.648	2.335
Systems Perception	1.693	2.200	1.571	2.327
Identifying Downstream Consequences	1.551	2.084	1.454	2.198
Identification of Key Causes	2.232	2.678	2.138	2.740
Judgment and Decision Making	2.289	2.714	2.075	2.719
Systems Evaluation	1.352	1.902	1.252	2.038
Occupational Working Conditions:				
Strength	6.485	5.714	8.834	8.217
Static Strength	1.809	1.515	2.328	2.099
Explosive Strength	0.794	0.731	1.358	1.357
Dynamic Strength	0.851	0.762	1.425	1.357
Trunk Strength	2.026	1.853	2.279	2.139
Stamina	1.006	0.852	1.444	1.266
Environmental Exposure	3.287	3.092	5.146	5.429
Sounds, Noise Levels are Distracting	0.808	0.782	1.132	1.224
Very Hot or Cold Exposure	0.649	0.582	1.074	1.041
Extremely Bright or Inadequate Light	0.654	0.597	0.925	0.927
Contaminants	0.788	0.733	1.150	1.197
Cramped Work Space, Awkward Position	0.340	0.329	0.630	0.706
Whole Body Vibration	0.049	0.069	0.235	0.333
Hazards	2.190	2.023	3.692	4.582
Radiation	0.053	0.035	0.021	0.034
Disease/Infections	0.994	0.746	0.203	0.158
High Places	0.057	0.063	0.386	0.561
Hazardous Conditions	0.200	0.254	0.801	1.182
Hazardous Equipment	0.394	0.551	1.483	2.082
Hazardous Situations	0.717	0.573	1.352	1.233
Protective Attire	0.912	0.877	1.358	1.679
Common Protective or Safety Attire	0.782	0.733	1.136	1.348
Specialized Protective/Safety Attire	0.130	0.144	0.222	0.331
Sample Size	41,690	135,223	13,852	165,790

Means of the O*NET occupational variables are calculated across individuals in the CPS (see the text and note to Table 1). The top variable listed in each category is an aggregate variable representing the sum of O*NET variables listed below. Scaling of detailed O*Net measures varies, but most run from 0-7. Detailed O*NET variables are included in the regressions. The Occupational Information Network (O*NET) is a comprehensive database system for collecting, organizing, and describing data on job characteristics and worker attributes.

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O*NET is sponsored by the U.S. Department of Labor's Employment and Training Administration and is intended to replace the Dictionary of Occupational Titles (DOT). Data are from O*NET 98, Version 0.9, containing 445 variables for 1,122 occupations. A more comprehensive version of O*NET is planned for the year 2001 and will include additional variables and information from ongoing job analyses. The O*NET database was created within the past several years by job analysts, based primarily on detailed job analyses, most of which were conducted as part of the DOT. Occupational information is being gathered on an ongoing basis in order to complete the revised O*NET. O*NET 98 provides a crosswalk mapping 1,122 O*NET codes to the approximately 500 Census occupational codes used in the CPS. For the most part the crosswalk between O*NET and the CPS is clear-cut, with many O*NET and Census occupations mapping one-to-one. Where more than one O*NET occupation is assigned to a Census occupation, mean values of the O*NET variables are calculated. In a small number of cases where no O*NET occupation maps directly to the Census, close occupational matches were readily identified. The matching process is more direct and far simpler than the complex mapping from the more than 12,000 DOT occupations to Census occupations.

Figure 1a: Frequency Distribution (in %) of Usual Hours Worked per Week by Women, 1989-97

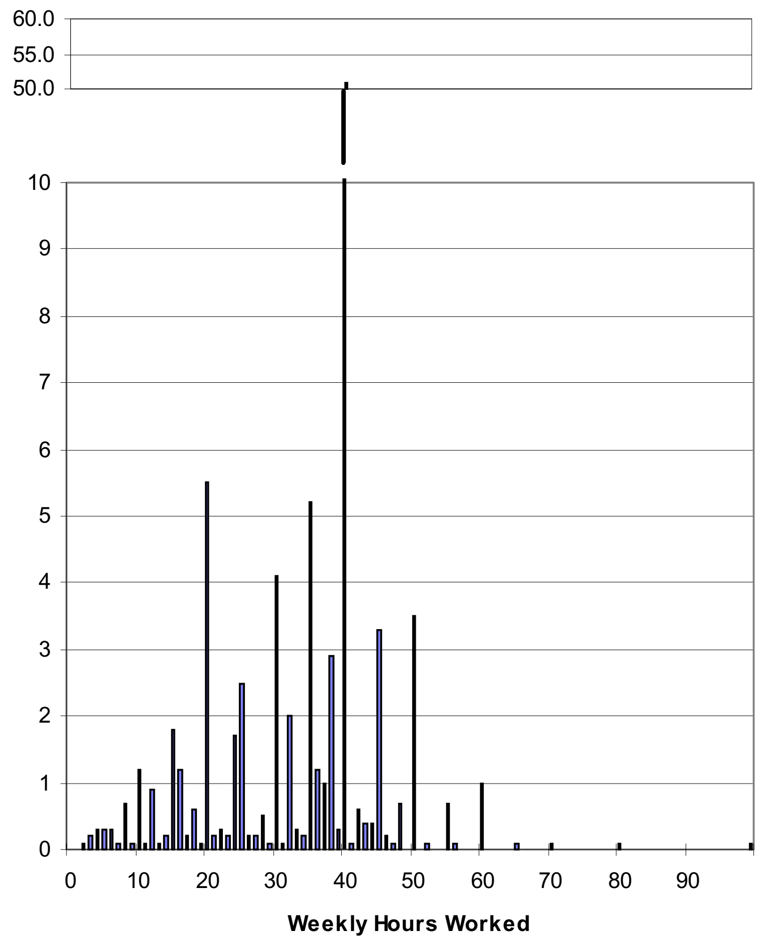


Figure 1b: Frequency Distribution (in %) of Usual Hours Worked per Week by Men, 1989-97

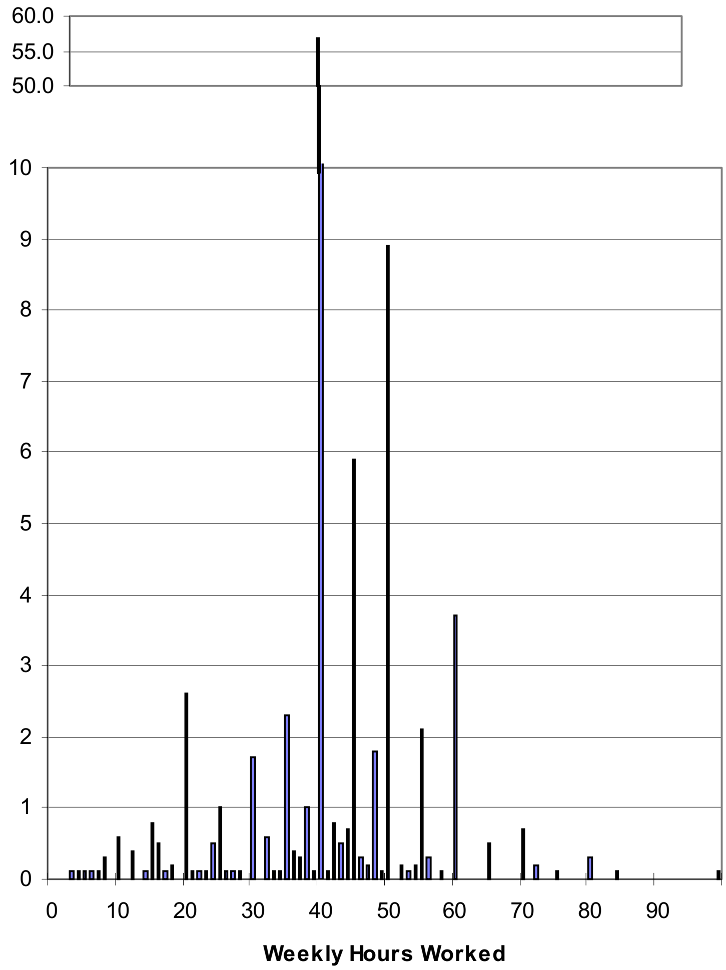


Figure 2a: Mean Wage by Hours Worked Among Women, 1989-97
(in constant 1997 dollars)

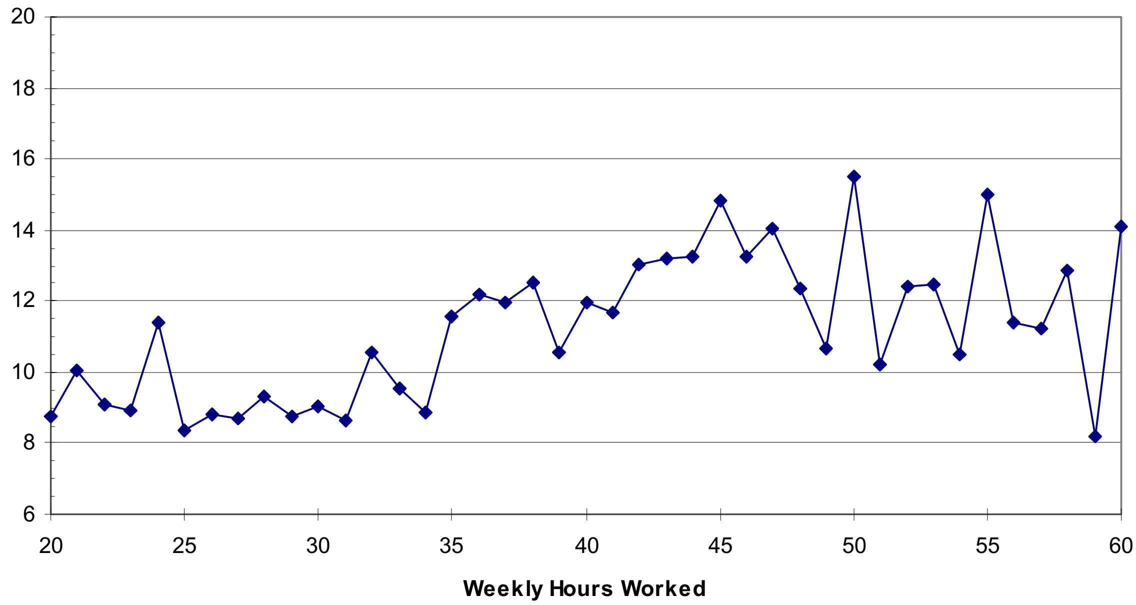
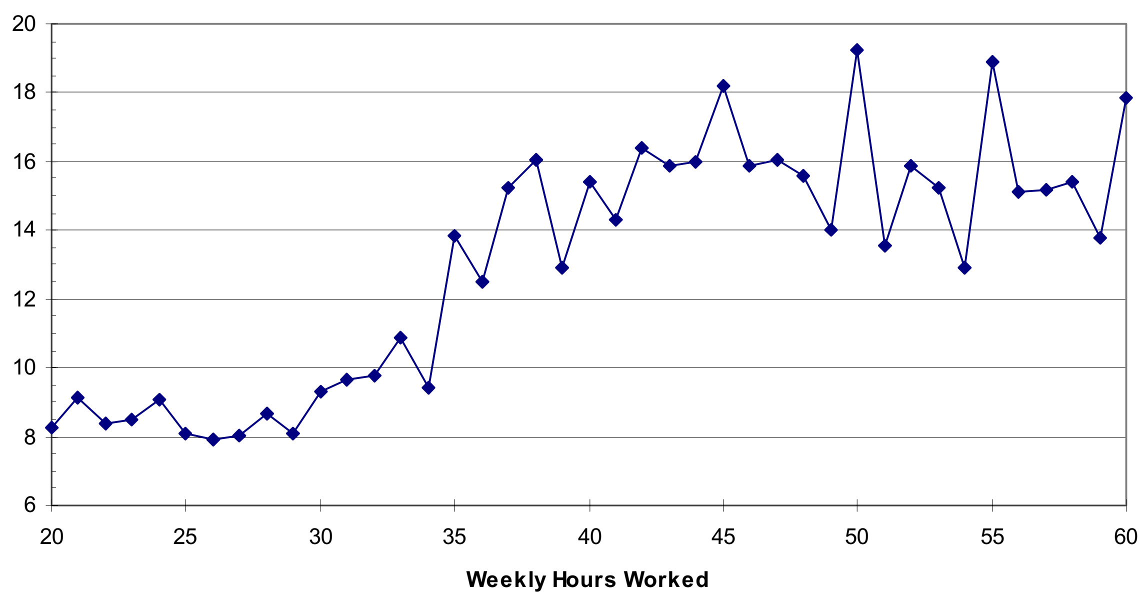


Figure 2b: Mean Wage by Hours Worked Among Men, 1989-97
(in constant 1997 dollars)



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