Women's part-time employment: a gross flows analysis

A decline in women's part-time employment since the early 1980's is due chiefly to women having become more likely to move from part-time to full-time employment and less likely to leave full-time employment once they get there

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Over the past three decades, the proportion of U.S. workers employed part time has grown rapidly.[1] For example, in 1957, the part-time employment rate was 12.1 percent, compared with 18.5 percent in 1990.[2] This increase, however, masks a significant decline in the rate during the late 1970's and a falloff from a peak of 20.6 percent in 1982 to the 18.5-percent figure of 1990.[3] The trend is primarily the result of a marked decline in the rate of part-time employment among women, set against only moderate increases in the rate among men. (See chart 1.) Still, the rate of part-time work for women is considerably greater than for men.

Although previous analyses of changes in the rate of part-time employment have focused chiefly on its growth, the insights they offer may be useful in identifying the sources of the decline in the rate as well. On the supply side has been the rapid growth of segments of the labor force with historically high propensities for part-time employment: women, teenagers, and older workers. Their greater preference for part-time work is usually attributed to a desire for greater flexibility of scheduling or fewer hours, because of home responsibilities, school, and health,[4] or to the use (among older workers) of part-time employment as a bridge to retirement.[5] One supply-side factor found *not* to have contributed to the growth of part-time work has been the overall growth in unemployment.[6] Interestingly, the supply-side explanations have zeroed in on changes in the size of groups with strong preferences for part-time work, rather than on changes in the preferences themselves.

Demand-side factors are twofold. First is the argument that firms are increasing their use of part-timers in order to decrease costs of production.[7] Lower costs are made possible through fewer fringe benefits,[8] less overtime pay,[9] the declining influence of unions,[10] and greater productivity or efficiency by part-time workers.[11] Second is the trend on the part of firms to gear more of their jobs toward part-time workers. For example, jobs in the retail sector are well suited to part-timers, with an emphasis on daily or weekly peak hours and on flexible schedules,[12] as are low-skilled jobs with routine and repetitive tasks.[13]

Of course, there may be interactions between various factors, such as the growth of the female labor force perhaps facilitating the growth of retail trade and the move toward low-skilled jobs possibly being in response to a growing low-skilled labor force.

The prevailing view of the sources of growth of part-time employment during the past three decades is that supply was likely more important through the 1960's and demand through the 1970's.[14] But what explains the decline since

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1980? True enough, the teenage and older populations have waned as a proportion of the labor force, but the share of women has continued to grow (albeit at a diminishing rate). Also, it is difficult to argue that firms have become less concerned about decreasing costs over the past decade. Likely explanations include a slowing of the transition toward industries and occupations with technologies that lend themselves to part-time work and an increased preference for full-time employment among women.

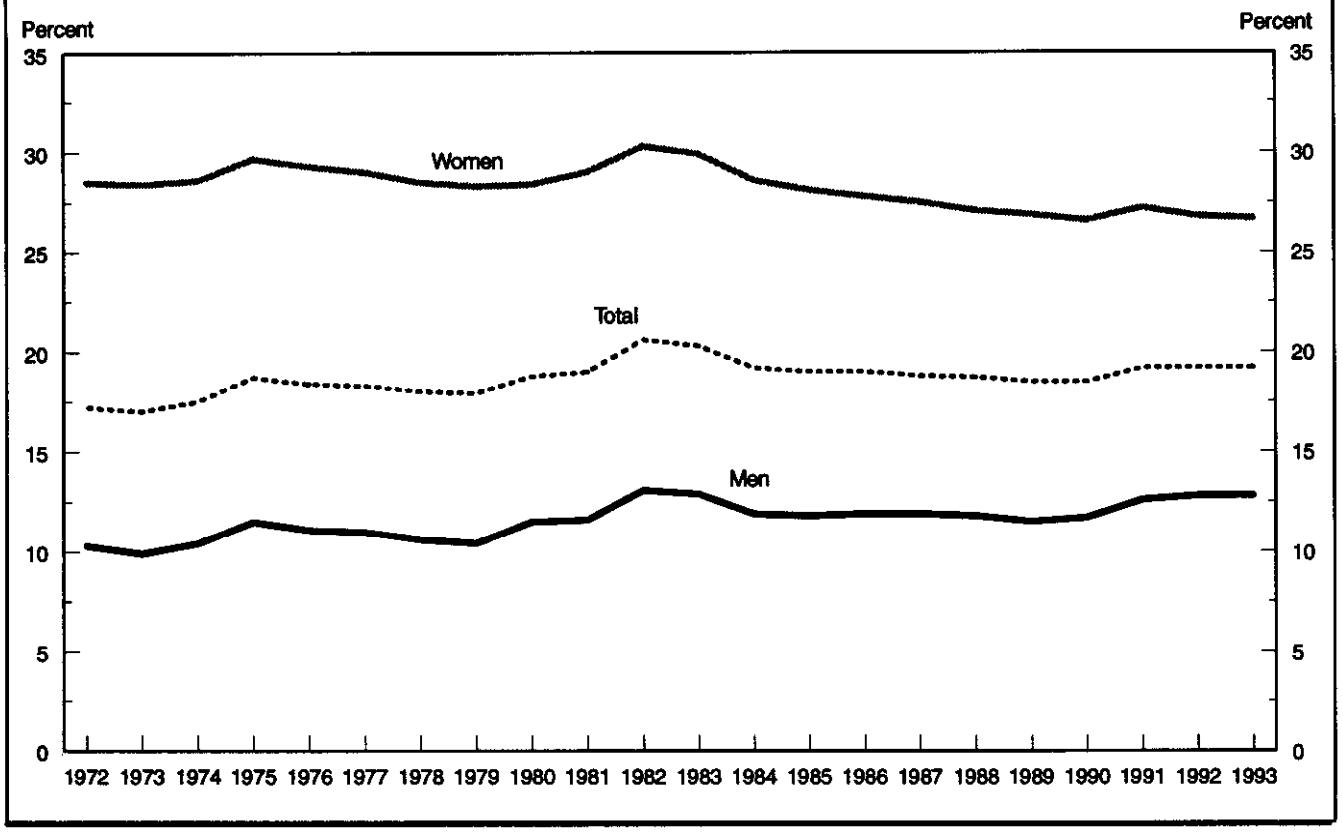
This article attempts to shed some light on the matter through an examination of gender differences in the levels of and trends in part-time employment in a dynamic context. In particular, it focuses on the transitions, or *flows*, among the labor market states of full-time employment, part-time employment, unemployment, and nonparticipation; the part-time employment *rate* at a point in time is a function of these flows. This approach has been used extensively in analyses of variations in unemployment and labor force participation rates.[15] Although the article is primarily descriptive, it can provide insights not available from simple time series analyses of the part-time employment rate alone, nor even from cross-sectional microlevel data (which would still be useful in analyzing differences in levels).

The flow approach

Let us begin by defining the following three mutually exclusive labor market states: full-time employment, part-time employment, and nonemployment.[16] Let the number of individuals from the population in each of those states at time t be F_t, P_t, and Z_t, respectively, and denote the numbers of individuals who make transitions from state to state during the interval $[t, t + 1]$ as FP_t, FZ_t, PF_t, PZ_t, ZF_t, and ZP_t. Then the transition rate between states I and J at time t is defined as $\lambda_{ij} = IJ_t/I_t$. Six transition rates describe the flows among the three states. As is the case for unemployment and labor force participation rates, the part-time employment rate can be expressed as a function of these rates of flow. Following Stephen T. Marston,[17] and defining the steady state as occurring when flows into a state equal flows out of a state, we can write the steady-state part-time employment rate, $PR = P/(P + F)$, as

$$(1)\quad PR = 1 / [1 + ((\lambda_{PF} + \lambda_{PZ})\lambda_{ZF} + \lambda_{PF}\lambda_{ZP}) / ((\lambda_{FP} + \lambda_{FZ})\lambda_{ZP} + \lambda_{FP}\lambda_{ZF})].[18]$$

Chart 1. Part-time employment rate, 1972–93



From this, it can be seen that the part-time employment rate is directly related to the rates of transition from full to part time (λ_{FP}) and from nonemployment to part time (λ_{ZP}) and inversely related to the rates of transition from part to full time (λ_{PF}) and from part time to nonemployment (λ_{PZ}). Consequently, trends in the part-time employment rate can be related to trends in these transition rates. Similarly, gender differences in the levels of and trends in the part-time employment rate can be attributed to gender differences in the levels of and trends in the various transition rates. The purpose of the empirical analysis that follows is to identify the transition rates representing the sources of the trends in these part-time employment rates and gender differentials. That is, we can determine whether the part-time rate is higher for women than for men, for example, because women are more likely to make transitions into part-time employment from nonemployment (λ_{ZP} is greater for women) or because they are less likely to make transitions from part-time to full-time employment (λ_{PF} is lower for women).

Data

The data used in this article are from table 4 of the unpublished "Gross Change Tables," available from BLS.[19] The table indicates, in a given month, the employment status of the civilian labor force by employment status in the previous month, for the entire population and by sex. The estimates are calculated by BLS using data from the Current Population Survey. Unlike other tables of gross-change data, table 4 differentiates between full- and part-time employment.[20] The data used in this article are from the tables for January 1980 through July 1989, the most recent available month. The figures are not seasonally adjusted. These raw flow data are used to calculate monthly transition rates between the four labor market states, for the entire period of the sample, by sex.

Average monthly transition rates for this period are presented in table 1. The estimates suggest that there is considerable movement between states over time. For example, on average, about 42 percent of the men employed part time in any given month moved to another state in the next month. The highest rates of flow are from part-time to full-time employment for men and from unemployment to nonparticipation for women. Note that these rates are not indicative of the raw magnitudes of the flows, because they are conditioned on the number of people initially in the state. On average in 1989, there were 38.7 million women employed full time, compared with 14.3 million employed part time, 3.0 million unemployed, and 41.6 million not participating in the labor force. For men, the respective figures were 56.9 million, 7.4 million, 3.5 million, and 20.9 million.

The table reveals some significant gender differentials in the average transition probabilities in the given period. The most striking is that men are much more likely than women to make the transition from part-time to full-time employment and less likely to make the converse transition from full-time to part-time employment. Men also are more likely to make the transition from unemployment to full-time employment. Indeed, the relative odds that an unemployed worker will move to full-time as opposed to part-time employment are about 1-1/2 times as high for men as they are for women. All of these differences cause the part-time employment rate to be lower for men than for women. Another gender difference is that men are significantly less likely to make the transition from unemployment to nonparticipation in the labor force, an observation noted in previous research.

One of the insights provided by this analysis is that the gender differential in part-time employment rates is a function of the rates of flow *out of* states, as well as the rates of flow into states. It is true that men are much more likely than women to enter full-time employment from the other states. But in addition, men are less likely to *leave* full-time employment once they get there. Less than 6 percent of full-time employed men made a transition out of that state, on average each month, compared with about 10 percent of full-time employed women. This contributes to men having a higher full-time employment rate. Similarly, although women are more likely to enter part-time employment from other states (except nonparticipation),[21] they are less likely to leave it once they get there. (The exit rate for women is 31 percent, versus 42 percent for men.) Note that women also are less likely than men to make transitions from part-time employment into unemployment and nonparticipation, as well as into full-time employment. These differences in levels of the transition rates will be examined again in a later section; the next section examines trends and the cyclical variability of the rates.

Empirical analysis

Using the flow data for the period from January 1980 to July 1989, we have a monthly time series of 115 observations for each of the 12 transition rates, for both men and women. To analyze the rates empirically, we simply estimate the parameters of the following equation for each transition, by sex:

$$(2) \qquad \ln(\lambda_{ij})_t = \alpha + \beta_1 \text{TIME}_t + \beta_2 \log(\text{URAT}_{t-1}) + \Gamma(\text{monthly dummies}).$$

Here, $(\lambda_{ij})_t$ is the rate of transition from state i to state j in month t, TIME takes the value 1 in January 1980, and URAT is the unemployment rate for men with spouse present, a commonly used measure for determining business cycle effects. A vector of monthly dummy variables is used to capture seasonal variations in the transition rates, with December being the excluded month. The coefficients α, β_1, and β_2 and the vector of coefficients, Γ, are estimated using standard multiple regression analysis.

The natural logarithm of the transition and unemployment rates is used, so that the coefficient β_2 represents the elasticity of the transition rate with respect to the unemployment rate. This makes comparisons of cyclical responsiveness fairly straightforward, across both rate and gender. The trend coefficient (β_1) can be interpreted as the average rate of growth of the transition rate. A lagged, rather than contemporaneous, unemployment rate is used simply to mitigate the effect of the simultaneous nature of the determination of the flow and unemployment rates. A specification also was estimated using the contemporaneous rate, which yielded results qualitatively the same as those presented in this article.

Five of the transition rate series exhibited evidence of first-order serial correlation for at least one gender. For those transitions, the parameters were estimated assuming a first-order autoregressive (AR(1)) process, using the Prais-Winsten procedure.[22] The parameters were estimated using ordinary least squares for the remaining seven transition rates. The estimated coefficients and their standard errors for the trend and cyclical variables are presented in table 2. (The coefficients for the seasonal dummies are available upon request.) The Durbin-Watson statistics are from the initial ordinary least squares regressions.

Trends. For men, there are significant negative trends in the probabilities of transitions from full-time employment to unemployment, from part-time employment to nonparticipation in the labor force, from unemployment to full-time employment, from nonparticipation to full-time employment, from nonparticipation to part-time employment, and from nonparticipation to unemployment. There is a significant positive trend in the probability of transition from unemployment to part-time employment. Some of these trends have contributed to the slight overall increase in the rate of part-time employment exhibited by men in the 1980–90 period (for example, the increase in λ_{UP} and decreases in λ_{UF} and λ_{NF}) and probably on through 1993, while others have worked against it and explain the decreasing rate since 1983 (for example, the decreases in λ_{FU} and λ_{NP}).

Table 1. **Percent of individuals making transitions among labor market states each month, by sex, January 1980–July 1989**

Transition	Men		Women	
	Mean	Standard deviation	Mean	Standard deviation
Full-time employment to:				
Part-time employment	3.58	0.47	7.47	1.06
Unemployment	1.42	.29	1.05	.21
Not in labor force	1.11	.46	2.20	.66
Part-time employment to:				
Full-time employment	27.66	4.42	19.44	3.57
Unemployment	5.56	.99	2.47	.49
Not in labor force	9.41	1.95	9.15	1.37
Unemployment to:				
Full-time employment	16.55	3.41	11.08	2.28
Part-time employment	10.41	1.70	11.99	1.70
Not in labor force	14.94	2.07	27.04	2.10
Not in labor force to:				
Full-time employment	2.42	1.00	1.41	.42
Part-time employment	3.24	.56	2.70	.41
Unemployment	3.49	.72	2.35	.37

SOURCE: Calculations from BLS data on gross change, January 1980–July 1989.

The transition rates for women have exhibited significant negative trends for the following transitions: full-time employment to unemployment, full-time employment to nonparticipation in the labor force, part-time employment to unemployment, and part-time employment to nonparticipation. All of these signify an increased degree of attachment to work among women over the decade of the 1980's; that is, at the end of the decade, women were less likely to leave full-time, as well as part-time, employment than they were at the beginning of the decade. The magnitudes of the coefficients indicate that the trends have been strongest in the flows from full-time employment, thereby contributing to the decline in the part-time employment rate. Also contributing to this decline are the positive trends in the probabilities of transitions from part-time employment to full-time employment and from nonparticipation in the labor force to full-time employment. At the same time, the rates of flow from unemployment to part-time employment and from nonparticipation to part-time employment also increased (reflecting the growth in labor force participation among women), which would tend to increase the part-time employment rate for women.

Note that there is no evidence of a significant positive trend in the transition rate from unemployment to full-time employment among

Table 2. Regression results from equation 2

Transition	Intercept[1]	TIME[1]	log (URAT)[1]	*R* squared	Durbin-Watson statistic
Men					
Full-time employment to:					
Part-time employment	[2]0.9213	0.0005	[2]0.1803	0.5500	2.204
	(.0893)	(.0003)	(.0476)		
Unemployment[3]	.0815	[2]–.0022	[2].2775	.6980	1.314
	(.1173)	(.0004)	(.0649)		
Not in labor force	.0087	–.0005	[4]–.1231	.9091	1.989
	(.0912)	(.0004)	(.0486)		
Part-time employment to:					
Full-time employment[3]	[2]3.159	.0002	–.0118	.8013	2.586
	(.0671)	(.0002)	(.0340)		
Unemployment[3]	[2]1.073	–.0004	[2].4326	.6671	1.467
	(.1115)	(.0004)	(.0610)		
Not in labor force	[2]2.542	[4]–.0007	[2]–.2345	.8563	2.062
	(.0680)	(.0003)	(.0362)		
Unemployment to:					
Full-time employment	[2]3.308	[5]–.0006	[2]–.5473	.8249	2.183
	(.0867)	(.0003)	(.0462)		
Part-time employment[3]	[2]2.612	[2].0016	[2]–.2926	.7987	2.266
	(.0648)	(.0003)	(.0357)		
Not in labor force	[2]3.307	[6]–.0000	[2]–.3916	.5355	1.802
	(.1060)	(.0004)	(.0565)		
Not in labor force to:					
Full-time employment	[2].9883	[2]–.0016	[2]–.3509	.9226	2.269
	(.0938)	(.0004)	(.0499)		
Part-time employment	[2]1.199	[4]–.0006	[2]–.1364	.7227	2.258
	(.0858)	(.0003)	(.0457)		
Unemployment[3]	[2].6024	[2]–.0017	[2].3877	.8548	1.947
	(.0737)	(.0003)	(.0393)		
Women					
Full-time employment to:					
Part-time employment	[2]1.748	–.0001	[4].0569	.8520	1.639
	(.0498)	(.0001)	(.0265)		
Unemployment[3]	[4]–.2903	[2]–.0026	[2].2132	.6322	1.710
	(.1251)	(.0005)	(.0676)		
Not in labor force	[2].8082	[2]–.0023	[2]–.1569	.9078	1.924
	(.0801)	(.0003)	(.0427)		
Part-time employment to:					
Full-time employment[3]	[2]2.862	[4].0005	[4]–.0701	.9256	2.077
	(.0462)	(.0002)	(.0245)		
Unemployment[3]	[2].3168	[4]–.0009	[2].3101	.7675	1.633
	(.1026)	(.0004)	(.0558)		
Not in labor force	[2]2.343	[2]–.0014	[2]–.1218	.8482	2.160
	(.0544)	(.0002)	(.0290)		
Unemployment to:					
Full-time employment	[2]2.880	–.0004	[2]–.5020	.7549	2.094
	(.0986)	(.0004)	(.0525)		
Part-time employment[3]	[2]2.768	[2].0009	[2]–.2547	.7497	1.450
	(.0774)	(.0003)	(.0425)		
Unemployment	[2]3.552	–.0001	[2]–.1520	.4748	2.108
	(.0504)	(.0002)	(.0268)		
Not in labor force to:					
Full-time employment	[2].2454	.0005	[2]–.2320	.9058	1.744
	(.0842)	(.0003)	(.0448)		
Part-time employment	[2]1.024	[4].0005	[2]–.1277	.7982	2.296
	(.0620)	(.0002)	(.0330)		
Unemployment[3]	.1307	–.0003	[2].3184	.8225	1.545
	(.0796)	(.0003)	(.0435)		

[1] Coefficient followed by standard error, in parentheses.

[2] Significant at the 0.01 level.

[3] Estimates are based on the assumption of a first-order autoregressive process. The Durbin-Watson statistics are from the original ordinary least squares regressions.

[4] Significant at the 0.05 level.

[5] Significant at the 0.10 level.

[6] Between –.00005 and 0.

women. (Indeed, the trend coefficient is of the opposite sign.) The growth in full-time employment of women, therefore, is not the result of an increase in the proportion of unemployed women finding full-time employment. Rather, it is the joint product of increases in the proportion of women moving from part-time to full-time employment and decreases in the proportion leaving full-time employment when they get there.

Finally, there are significant gender differences in transition rate trends that should be highlighted. First, there is a significant (0.01-level) difference in the trend coefficient for the transition from full-time employment to nonparticipation in the labor force, with women exhibiting a greater decline. Similarly, the trend coefficient is significantly larger (in absolute value) for women for the transition from part-time employment to nonparticipation. And the coefficients for all of the transitions from nonparticipation are significantly different between the genders, even exhibiting different signs for the transitions from nonparticipation to part-time employment and from nonparticipation to full-time employment.

Cyclical variability. Although the focus of this article has not been on the cyclical variability of the part-time employment rate for either gender, one of the most striking features of table 2 is the strong cyclical responsiveness of nearly all of the transition rates. Both men and women exhibit strong decreases in the rate of flow out of unemployment as unemployment rates rise (as would be expected), as well as increases in the rate of flow into unemployment from all of the other states. For both sexes, there is a difference between the responsiveness of the flows into and out of full- and part-time employment: flows from unemployment to full-time employment are more cyclically sensitive than those from unemployment to part-time employment, while flows from part-time employment to unemployment are more cyclically sensitive than those from full-time employment to unemployment.

Consistent with evidence regarding adjustments of hours over the business cycle,[23] the rate of flow from full-time to part-time employment increases for both genders in an economic downturn. By contrast, the rate of flow from part-time to full-time employment decreases only for women as the unemployment rate rises. The findings regarding the cyclical responsiveness of the transitions from nonparticipation in the labor force to unemployment and from unemployment to nonparticipation are consistent with those from earlier periods,[24] indicating no evidence of discouraged worker effects, but significant added worker effects.

There are significant gender differences in the cyclical responsiveness of several of the transition rates. The effect of an increase in the unemployment rate is significantly greater for men for the transitions from full-time to part-time employment, from part-time employment to unemployment, from part-time employment to nonparticipation in the labor force, from unemployment to nonparticipation, and from nonparticipation to full-time employment. These findings are consistent with results from earlier work that did not differentiate between full- and part-time employment.[25] They have implications for that work, however, as the gender differences in exit rates from employment appear to be from part-time rather than full-time employment, at least for the period examined in this article.

Discussion

The results of the preceding empirical analysis suggest that there are several key transition rates contributing to the gender differential in the level of the part-time employment rate and to trends in that rate over the 1980–90 period and probably on through 1993. First, the part-time employment rate for women is higher than that for men because women have higher probabilities of transitions from full-time employment to both part-time employment and nonparticipation in the labor force, and lower probabilities of transitions from part-time employment to both full-time employment and unemployment, as well as from unemployment to full-time employment. But at the same time, the part-time rate for women has been falling because their rates of transition from full-time employment to both unemployment and nonparticipation are falling, while their rates from part-time employment and nonparticipation to full-time employment have been rising.

What are the sources of these differences and trends? In earlier work, I presented a dynamic model in which differences or changes in the rates of flow among the part-time, full-time, unemployment, and nonparticipation states arise because of differences or changes in the wage rate earned on the job, the costs of searching for work while unemployed, and the value of leisure time.[26] In the context of that model, an increase in the wage increases the rate of flow into and decreases the rate of flow out of full-time employment, while the rates into and out of part-time employment are unchanged. An increase in the value of leisure increases the rate of flow into nonparticipation from all states, decreases the rates of flow into full-time employment and unemployment, increases the rates of flow out of full-time employment and unemployment, and increases the rates of flow into part-time employment from full-time employment and unemploy-

ment. Increases in the costs of searching for a job affect only the flows into and out of unemployment. In a more general model, the flows into and out of part-time employment would also depend on the costs of a job search while employed, as well as on the difference between the wages in part- and full-time jobs. Finally, the rates of flow into full- and part-time work depend on the rates at which workers are offered full- versus part-time jobs.

Given these hypotheses, the gender differentials in the transition rates from full-time to part-time employment, from part-time to full-time employment, and from unemployment to full-time employment could result from higher wages among men and a higher value placed upon leisure among women. One source of the latter is the unequal distribution of responsibilities at home, including responsibilities for housekeeping, cooking, and child care. The gender differentials could also result from higher rates of arrival of full-time (versus part-time) job offers for men.

It is possible that changes in wages are one source of the *trends* in transition rates as well, as wages of women increased relative to those of men in the 1980's.[27] According to the model, this trend in wages would increase women's rates of flow into, and decrease their rates of flow out of, full-time employment, relative to men. But this hypothesis is somewhat weakened by the fact that the earnings ratio of women to men was also rising (albeit at a variable rate) for a long period during which the rate of flow into part-time employment for women was increasing. A similar statement could be made with regard to the hypothesis that the trends in flows are the result of a decline in the value of leisure among women. However, an increase in the availability of child care in the 1980's may have accelerated such a deadline. Nevertheless, both the rising-wage and the falling-value-of-leisure hypotheses are consistent with the finding that women's rates of flow out of full-time employment have been falling.

Another hypothesis regarding the decline in part-time employment is that the rate at which full-time jobs are being offered has been rising. However, although this hypothesis is consistent with the findings regarding the transition rates from nonparticipation to full-time employment and from part-time to full-time employment, it is not especially convincing, as we find no evidence that the rate of transition from unemployment to full-time employment has been rising. Moreover, a major cause of the decrease in the part-time employment rate is the tendency for women to be more likely to *stay* in full-time jobs rather than more likely to *get* them.

IN SUM, over the 1980–89 period, several rates of transition to and from full-time employment, part-time employment, unemployment, and nonparticipation in the labor force have exhibited trends that contribute to the declining propensity to work part time, especially among women. The analysis in this article points to one important source of such change: a decreased propensity, particularly among women, to *leave* full-time employment. Changes in wages, in the value of leisure, and in the rate at which full-time jobs are being offered are potential contributors to the trends uncovered, including the decline in part-time employment among women. □

Footnotes

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[1] The Bureau of Labor Statistics classifies an individual as a part-time worker if the total hours the individual has worked during the week is less than 35. Recent work by Julie L. Hotchkiss ("The Definition of Part-time Employment: A Switching Regression Model with Unknown Sample Selection," *International Economic Review*, November 1991, pp. 899–917) indicates that the 35-hour cutoff does not apply to many part-time workers in a practical sense. Still, the BLS definition is used throughout this article. Note that the definition refers to total hours at *all jobs*, not hours per job, so that the statistics commonly reported do not measure the prevalence of part-time *jobs*. Indeed, an individual holding several part-time jobs could be counted as one full-time person, depending on the total weekly hours he or she has worked.

[2] Recent articles highlighting this growth are by Chris Tilly ("Reasons for the continuing growth of part-time employment," *Monthly Labor Review*, March 1991, pp. 10–18) and Bernard E. Ichniowski and Anne E. Preston ("New Trends in Part-time Employment," *Proceedings of the 38th Annual Meeting of the Industrial Relations Research Association* (Madison, WI, Industrial Relations Research Association, 1986), pp. 60–67).

[3] Rates are calculated from *Labor Force Statistics Derived from the Current Population Survey, 1948–87*, Bulletin 2307 (Bureau of Labor Statistics, 1988); and *Employment and Earnings* (Bureau of Labor Statistics, various issues). Note that there is a difference between the proportion of the *employed* who work part time, which is the focus of this article, and the proportion of the *labor force* or of the *population* who work part time. It is possible to have the first of these fall and the other two rise over time if the overall employment rate increases sufficiently.

[4] See, for example, Tilly, "Growth of part-time employment"; and Thomas J. Nardone, "Part-time workers: who are they?" *Monthly Labor Review*, February 1986, pp. 13–19.

[5] See Christopher J. Ruhm, "Bridge Jobs and Partial Retirement," *Journal of Labor Economics*, October 1990, pp. 482–501.

[6] See Tilly, "Growth of part-time employment"; and Ichniowski and Preston, "Part-time Employment."

[7] For an interesting analysis of the demand for part-time workers, see Mark Montgomery, "On the Determinants of Employer Demand for Part-time Workers," *Review of Economics and Statistics*, February 1988, pp. 112–17.

[8] See Ichniowski and Preston, "Part-time Employment"; and 9to5, *Working at the Margins* (Cleveland, National Association of Working Women, 1986). Ichniowski and Preston found that in 1977 the probability of receiving a fringe benefit was between 11 and 25 percentage points lower for part-time than for full-time workers. In 1987, 17.8 percent of part-time workers had direct health insurance coverage, compared with 68.2 percent for full-time workers. (See Sheila R. Zedlewski, *Expanding the Employer-Provided Health Insurance System*, Report 91–3 (Washington, Urban Institute Press, 1991).) The difference in treatment is largely attributable to institutional factors. For example, the Employee Retirement Income Security Act does not require employers to provide pensions to employees working fewer than 1,000 hours per year (about 20 hours per week). The historically high proportions of married women and teenagers among the part-time worker population, who often are covered by the insurance plan of a spouse or parent, have afforded a further incentive to employers to forego offering health care benefits to part-timers.

[9] Richard S. Belous, *The Contingent Economy: The Growth of the Temporary, Part-time and Subcontracted Workforce* (Washington, National Planning Association, 1989).

[10] See Tilly, "Growth of part-time employment"; and 9to5, *Working at the Margins*.

[11] Jean Hallaire, *Part-time Employment: Its Extent and Its Problems* (Paris, Organization for Economic Cooperation and Development, 1968).

[12] *Ibid.*

[13] Stanley D. Nollen, Brenda Broz Eddy, and Virginia Hider Martin, *Permanent Part-time Employment* (New York, Praeger, 1978).

[14] Tilly, "Growth of part-time employment."

[15] For analyses focusing on the unemployment rate, see Stephen T. Marston, "Employment Instability and High Unemployment Rates," *Brookings Papers on Economic Activity*, no. 1 (The Brookings Institution, 1976), pp. 169–203; Ronald G. Ehrenberg, "The Demographic Structure of Unemployment Rates and Labor Market Transition Probabilities," in Ronald G. Ehrenberg, ed., *Research in Labor Economics*, vol. 4 (Greenwich, CT, JAI Press, 1980); and Larry DeBoer and Michael C. Seeborg, "The Unemployment Rates of Men and Women: A Transition Probability Analysis," *Industrial and Labor Relations Review*, April 1989, pp. 404–14. For analyses of participation rates, see Donald R. Williams, "Employment in recession and recovery: a demographic flow analysis," *Monthly Labor Review*, March 1985, pp. 35–42; and *Labor Force Participation of Black and White Youth* (Ann Arbor, MI, UMI Research Press, 1987); and Ralph Smith and Jean Vanski, "The Volatility of the Teenage Labor Market: Labor Force Entry, Exit, and Unemployment Flows," in *Conference Report on Youth Unemployment: Its Measurement and Meaning* (Department of Labor, 1978). See also Olivier J. Blanchard and Peter Diamond, "The Cyclical Behavior of the Gross Flows of U.S. Workers," *Brookings Papers on Economic Activity*, no. 2 (The Brookings Institution, 1990), pp. 85–143. Rebecca M. Blank, "The Dynamics of Part-Time Work," mimeograph (July 1992), presents a dynamic analysis of transition rates for part-time employment in a cross section of women from the Panel Study of Income Dynamics.

[16] This three-state case is presented only for expositional purposes. The full four-state case is analyzed subsequently.

[17] Marston, "Employment Instability."

[18] This equation is derived in the appendix.

[19] For a general description of the data on gross change and their problems, see Paul O. Flaim and Carma R. Hogue, "Measuring labor force flows: a conference examines the problems," *Monthly Labor Review*, July 1985, pp. 7–17. For a method of adjusting the data, see John M. Abowd and Arnold Zellner, "Estimating Gross Labor Force Flows," *Journal of Business and Economic Statistics*, July 1985, pp. 254–83. In the analysis that follows, I use the raw, unadjusted data.

[20] The table gives the number of persons employed full time, part time for economic reasons, and part time "voluntarily"; the number with jobs but not at work (broken down by reason); the number unemployed; and the number out of the labor force (again, broken down by reason). Unlike published figures in *Employment and Earnings*, in these tables, those with jobs but not at work have not been allocated to the full- and part-time employment categories. In the empirical analysis that follows, I have allocated them according to the ratio used by the Bureau for its published tables, which is based on whether the individual "usually" worked full or part time.

[21] Using the absolute rates in the table, we see that women are less likely than men to make the transition from nonparticipation in the labor force to part-time employment. If one considers that rate relative to the rate of flow from nonparticipation to full-time employment, however, then women are more likely to end up in part-time employment.

[22] The parameters also were estimated with AR(2) and AR(1,1) specifications, yielding essentially the same results.

[23] See Robert W. Bednarzik, "Short workweeks during economic downturns," *Monthly Labor Review*, June 1983, pp. 3–11.

[24] Williams, "Employment in recession and recovery" and DeBoer and Seeborg, "Unemployment Rates of Men and Women."

[25] *Ibid.*

[26] Donald R. Williams, *A Dynamic Analysis of Recent Changes in the Rate of Part-time Employment*, Working Paper No. 9120 (Cleveland, Federal Reserve Bank of Cleveland, December 1991). The variables cited in the text are the same as those found in static models of the decision to work full as opposed to part time. See, for example, Mark Killingsworth, *Labor Supply* (Cambridge, U.K., Cambridge University Press, 1983); for empirical analyses, see Richard D. Morgenstern and William Hamovitch, "Labor Supply of Married Women in Part-time and Full-time Occupations," *Industrial and Labor Relations Review*, October 1976, pp. 59–76; and Ethel B. Jones and James E. Long, "Part-Week Work by Married Women," *Southern Economic Journal*, January 1980, pp. 716–25.

[27] As is well known, there has been no significant upward movement of average wages in the economy overall during the 1980's, so I focus on the men-women differential here. If we consider *after-tax* wages, however, then the hypothesis might be consistent with the trend, even for men.

APPENDIX: Derivation of the part-time employment rate

In the steady state, the flow into a state is equal to the flow out of that state. For part-time employment, this condition implies that

$$(A1) \qquad P(\lambda_{PF} + \lambda_{PZ}) = F\lambda_{FP} + Z\lambda_{ZP},$$

and for full-time employment,

$$(A2) \qquad F(\lambda_{FP} + \lambda_{FZ}) = P\lambda_{PF} + X\lambda_{ZF}.$$

Multiplying equation (A1) by λ_{ZF} and equation (A2) by λ_{ZP}, and subtracting the latter equation from the former, we obtain

$$(A3) \qquad P[(\lambda_{PF} + \lambda_{PZ})\lambda_{ZF} + \lambda_{PF}\lambda_{ZP}] = F[(\lambda_{FP} + \lambda_{FZ})\lambda_{ZP} + \lambda_{FP}\lambda_{ZF}].$$

Solving this equation for P, substituting for P in $PR = P/(P + F)$, and rearranging terms yields equation (1) in the text.