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Abstract

[Excerpt] To say that part-time workers are less costly than full-time workers, however, is *not* an explanation for the trend in the use of part-time employees that has occurred. Rather, one must show that the relative cost advantage of part-time workers has increased over time *and* that variations in the relative cost advantage are associated with variations in the usage of part-time employment. Somewhat surprisingly, few researchers have tried to do this, and even these only indirectly.

This paper addresses this issue, albeit in a slightly different way, focusing on data from the United States. We begin in the next section by analyzing data on part-time employment in the United States. After controlling for cyclical factors, an increasing trend in the usage of part-time employment is observed. Moreover, it is clear from the component of part-time employment that is increasing that this is a demand-side, not a supply-side phenomenon.

Keywords

part-time employment, labor cost, labor demand, labor market, wage differentials, benefits

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Comments

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CHAPTER 13

Part-time Employment in the United States

RONALD G. EHRENBERG, PAMELA ROSENBERG and JEANNE LI

13.1 Introduction

It is well documented that part-time employment has grown as a share of total employment in most European countries since the early 1970s (see, for example, OECD, 1985). To the extent that this growth reflects voluntary behavior by workers it has important implications for economic welfare, because it allows an increased number of workers to be employed at any given level of aggregate demand. Put another way, explicit or even implicit work-sharing arrangements allow for easier absorption of a growing labor force into employment.

For the most part, explanations for the growth in part-time employment have focused on the supply side of the labor market and the changing industrial composition of employment. The growing shares of married women with children in the labor force, of older workers phasing into retirement, and (at least in the United States) of students who need to work to help finance their education, coupled with the growth in the share of service sector employment, have all been thought to contribute to part-time employment growth. (See, for example, OECD, 1983, 1985; Owen, 1979.)

Recently, though, attention has shifted to the demand side of the labor market and the role that relative costs play. In theory, *ceteris paribus*, part-time employment should expand relative to full-time employment if the hourly labor cost (wages and hour-related fringes) of part-time workers falls relative to that for full-time workers, or if the quasi-fixed cost (non-hour-related fringes and hiring and training costs) of part-time workers falls relative to that of full-time workers (see FitzRoy and Hart, 1986). Numerous studies for various OECD countries in fact show that part-time workers' wages and fringe benefits are often lower than those for full-time workers, and that the part-time workers are sometimes not eligible for employer-financed social insurance programs or redundancy payments.[1]

To say that part-time workers are less costly than full-time workers, however, is *not* an explanation for the trend in the use of part-time employees that has occurred. Rather, one must show that the relative

cost advantage of part-time workers has increased over time *and* that variations in the relative cost advantage are associated with variations in the usage of part-time employment. Somewhat surprisingly, few researchers have tried to do this, and even these only indirectly.[2]

This paper addresses this issue, albeit in a slightly different way, focusing on data from the United States. We begin in the next section by analyzing data on part-time employment in the United States. After controlling for cyclical factors, an increasing trend in the usage of part-time employment is observed. Moreover, it is clear from the component of part-time employment that is increasing that this is a demand-side, not a supply-side phenomenon.

The following two sections attempt a *partial* test of several demand-side explanations. Although, we have no data on how part-time/full-time wage and fringe benefit differentials have varied over time, we can estimate how they vary across industries at a point in time. We do this in Section 13.3 using data from the March 1984 *Current Population Survey* (*CPS*), a sample of approximately 122,000 individuals aged 14 and older. These estimated differentials are then entered in Section 13.4 as explanatory variables in a simple structural model of the inter-industry determinants of part-time employment and we attempt to estimate their effects. That is, we try to infer if there is an inter-industry relationship between the relative cost advantage of part-time employees and their usage.[3] If such a relationship exists, evidence that the relative cost advantage has been increasing over time could then be used to estimate the importance of relative costs in explaining the trend in the usage of part-time employees that has occurred.[4]

Section 13.3 also focuses on a potential demand-side influence that has been ignored to date, namely employers' alleged desire in the United States to maintain non-union work environments. It is often asserted (though the evidence does not always support this assertion) that part-time employees, who tend to have shorter expected tenure with a firm than otherwise comparable full-time employees, are less likely to be union members or vote for a union than their full-time counterparts.[5] *Ceteris paribus*, the greater the differential in the probability of being a union member between part-time and full-time workers in an industry, the greater the advantage that will accrue to employers trying to 'keep out' unions from hiring part-time workers.

We present two efforts to see if this influence may have contributed to the growth of part-time employment in the United States. First, we trace if government policy in the United States has made it increasingly difficult for part-time workers to join unions in recent years. Second, we estimate part-time/full-time probability of union membership differentials by industry, with the goal of including such differentials in our inter-industry cross-section model of part-time employment variation. Unfortunately, these differentials prove in the main to be statistically imprecise, so we are unable to include them in the inter-industry model.

A brief concluding section summarizes our findings and their implications for public policy and future research.

13.2 Trends in Part-time Employment in the United States, 1955–1984

The definition of what constitutes a part-time employee varies widely across countries, which makes international comparisons difficult. A person is classified as working part-time in the United States if he or she works less than 35 hours per week during the *CPS* survey week. This is a much higher cut-off point than exists in most OECD nations and, unlike them, bases part-time status on current week hours rather than usual weekly hours.

Part-time employees are further broken down into those employed part-time for non-economic reasons and those employed part-time for economic reasons. The former are voluntary part-timers; this category includes individuals who work part-time because of family or school responsibilities. The latter are involuntary part-timers; this category includes people who are temporarily part-time owing to cyclical or seasonal factors, or to temporary firm-specific fluctuations in demand, but who would prefer a full-time job. Changes in the part-time employment for non-economic reasons category reflect supply behavior, while changes in the total, and especially in the part-time for economic reasons category, probably reflect demand factors.

Appendix Tables 13A.1–13A.5 of our paper (pp. 277–81) present background data on the growth of part-time employment in the United States. Table 13A.1 contains annual data for 1963–84 on the percentage of employees who worked part-time for the total employed workforce, for youths aged 16–19, and for adult males and adult females. These raw data do not control for cyclical factors and they suggest that since 1980 only the percentage of teenagers who work part-time has increased.[6] This lack of increase in the overall part-time employment rate in the United States in recent years has been noted by other observers, as has the relatively low incidence of part-time employment among adult males (OECD, 1985).

Table 13A.2 contains annual data for 1955–84 on the percentage of employees employed part-time; the data are presented separately for part-timers for economic and non-economic reasons. The latter category, which represents voluntary part-time employees, almost doubled during the period, peaking around 1980. The former fluctuated substantially from year to year, suggesting that trends in it will be obscured when one ignores cyclical factors. Similar data are presented in Table 13A.3 for various demographic groups (teenagers, adult males, adult females, Whites, and non-Whites), starting in 1968 when these more detailed data were first published. Although the failure to control for cyclical factors may distort things, there does appear to be an increasing trend in part-time employment for economic reasons for each group, suggesting that demand-side forces may be important.[7]

Might part of the apparent trend in part-time employment be an artifact of the 35-hour cut-off point (which is higher than many European countries) for the classification of part-time employees? If,

through collective bargaining or unilateral employer adoption, standard workweeks for some regular full-time employees were reduced over time from, say, 40 hours to 32 hours a week, these workers would be classified as part-time in recent years. As a result, average weekly hours of part-time employees would be higher in recent than in earlier years. In fact, Table 13A.4 does indicate that average weekly hours of part-time employees in each of the demographic groups increased over the 1968–84 period. The increases probably are not sufficiently large, however, to support the contention that the apparent trends in part-time employment primarily reflect some full-time employees now working fewer than 35 hours a week.

Finally, by way of background, Table 13A.5 presents data for 1968 and 1984 on the shares of adult males, adult females and teenagers in total and part-time employment. In spite of the increasing proportion of employed teens who work part-time, the teenage share of part-time employment declined over the period because the share of teens in the United States labor force declined. Part-time employment, for both economic and non-economic reasons, continues to be primarily an adult female phenomenon in the United States.

Of course, all of the above conclusions come from visual inspections of data that do not control for cyclical factors. To be a bit more precise, Table 13.1 presents estimates of the time-trends in the percentage of non-agricultural workers who are employed part-time in the United States, which we obtained from models that included the adult unemployment rate to control for cyclical factors. Separate estimates are presented for part-time workers for economic and non-economic reasons and the time-trends are estimated over a variety of periods (1955–84, 1963–84, 1968–84 and 1973–84) to see how sensitive the results are to the starting date. Results are also presented for various demographic groups during the 1968–84 period.

This table does suggest the importance of controlling for cyclical factors, of varying the starting date and of disaggregation. Over the 1973–84 period, the trends in the percentage employed part-time for economic reasons and part-time for non-economic reasons were positive and negative (but not statistically significant), respectively, with the former clearly dominating.[8] Since the positive trend in total part-time employment (after controlling for cyclical forces) coincides with a non-positive trend in part-time employment for non-economic reasons, demand-side forces *must* be responsible for the increase in total part-time employment. It is interesting to note, however, that if one starts the analyses in 1955 a positive trend in part-time employment for non-economic reasons emerges while there is no trend in part-time employment for economic reasons. The growth of part-time employment in the early years of the longer period obviously was influenced by supply-side forces.

Focusing on the 1968–84 period, one sees from the bottom panel of the table that, while adult males are increasingly working part-time voluntarily, adult females are decreasingly doing so. Similarly, the increasing

Table 13.1
Estimated Time-Trends in the Percentages of Non-agricultural Workers Employed Part-time in the United States

	Part-time for:	
	Economic reasons	*Non-economic reasons*
All workers		
1955–84	.011 (0.7)	.208 (4.2)
1963–84	.042 (2.1)	.140 (2.4)
1968–84	.083 (3.4)	.024 (0.5)
1973–84	.113 (3.8)	−.055 (1.4)
1968–1984		
Males age 20+	.039 (1.7)	.052 (4.0)
Females age 20+	.035 (0.6)	−.117 (2.4)
Both sexes 16–19	.202 (3.0)	.122 (0.9)
Whites	.088 (4.3)	.055 (1.3)
Non-whites	.091 (1.1)	−.164 (3.4)

Source: Obtained from models that used annual data, included the adult male unemployment rate to control for cyclical factors, and that corrected for autocorrelation using the Prais–Winsten method. See Tables 13A.1–13A.5 for the underlying data.
Note: Absolute value of t statistic in parentheses

probabilities observed (after controlling for cyclical factors) that teenagers and adult males involuntarily work part time are not matched by an analogous trend for females. The female share of all part-time employment increased during the period (Table 13A.5) only because of the increased share of women in the labor force.

The bottom line of this section is that in recent years in the United States there *has* been a positive trend in the proportion of people employed part-time and that the data suggest this has been due to employer, not employee, decisions.

13.3 Part-time Employee/Full-time Employee Wage, Fringe Benefit and Union Membership Differentials

The previous section suggests that there has been a trend towards increased employment of part-time workers vis-à-vis full-time workers in recent years in the United States and that this trend is probably due to demand-side factors. Do these factors include a growing cost advantage of part-time workers caused by increases over time in differentials in part-time/full-time wages, fringe benefits and probability of being a union member? Time-series data are not available to answer this question directly. Instead, in this section, we estimate the extent to which these differentials vary across industries at a point in time. In the next section, we then see if these estimated differentials can help explain inter-industry variations in the usage of part-time employment.

The March 1984 *Current Population Survey* is a national probability sample of roughly 122,000 individuals aged 14 and older. We restricted our attention to employed wage and salary workers who were working during the survey week and were *not* self-employed, and stratified the sample into 46 industry subsamples. Recalling that part-time employees are defined as those who work less than 35 hours during the survey week, Table 13.2 presents background data on the number of individuals in each subsample and the percentage of these employees classified as part time. The percentages vary widely across industries, from about 2 per cent in the petroleum and coal products industry to 40 per cent in retail trade.[9]

For each industry, we used the individual observations to estimate wage equations of the form

$$\log W_{ji} = a_{0i} + \sum_{k=1}^{m} a_{ki} X_{kj} + a_{pi} P_j + e_{ji}, \quad i = 1, 2, \ldots 46. \tag{13.1}$$

Here W_{ji} is the hourly earnings of individual j in industry i, the X_{kj} are a set of variables available in the *CPS* data to control for human capital, cost-of-living and other factors that influence wages, m is the number of these variables, P_j is a dichotomous variable that takes on the value of 1 if the individual is a part-time employee and zero otherwise, e_{ji} is a random error term, and the a's are parameters to be estimated. The estimate of a_{pi} is an estimate of the extent to which part-time workers in the industry i are paid less than full-time workers in the industry.[10]

The control variables used include years of schooling completed, proxies for years of potential labor force experience (age minus years of school minus five) and experience squared, the number of children in the family, and dichotomous variables for marital status, gender, race, veteran status, hispanic ethnicity, being a student, residence in a standard metropolitan statistical area, and residence in various census regions. As is well known, the effects of some of these variables on earnings (children, experience, education, marital status) are often found to vary with gender, so some specifications interact gender with these variables. Finally, some specifications use part-time status in the survey week as an explanatory variable, while others use usual part-time employment status.

The estimates of the part-time/full-time wage differentials that we obtained from these models are displayed in Table 13.3. As noted there, the differentials in the columns headed *RW*1 and *RW*3 (*RW*2 and *RW*4) are based on the survey week (usual) part-time/full-time dichotomy, and those in the columns headed *RW*1 and *RW*2 (*RW*3 and *RW*4) are from the models without (with) interaction terms. As might be expected, the four sets of estimates are highly correlated (in the range of .90 to .98) across industries. This suggests, and the next section confirms, that the estimated effects we will obtain from the inter-industry analyses will be relatively insensitive to the relative wage differential measure that we use.

Table 13.2
Distribution Across Industries of the
March 1984 Current Population Survey Sample

Industry code	*Number of observations*	*Percentage part-time*	*Industry description*
1	971	24	Agriculture
2	707	5	Mining
3	3238	14	Construction
4	456	8	Lumber and Wood Products
5	319	5	Furniture and Fixtures
6	324	4	Stone, Clay, Glass and Concrete
7	512	4	Primary Metals
8	789	5	Fabricated Metals
9	—	—	Not Specified Metals
10	1551	4	Machinery, except Electrical
11	1336	3	Electrical Machinery
12	623	2	Motor Vehicles and Equipment
13	323	3	Aircrafts and Parts
14	455	3	Other Transportation Equipment
15	395	5	Professional and Photographic Equipment and Watches
16	84	7	Toys, Amusements and Sporting Goods
17	257	14	Miscellaneous Manufacturing
18	1061	11	Food
19	35	11	Tobacco
20	427	9	Textiles
21	836	12	Apparel
22	405	4	Paper and Allied Products
23	1076	20	Printing, Publishing and Allied Industries
24	715	3	Chemicals and Allied Products
25	113	2	Petroleum and Coal Products
26	440	9	Rubber and Miscellaneous Plastic Products
27	182	20	Leather and Leather Products
28	2507	13	Transportation
29	989	6	Communications
30	974	3	Utilities and Sanitation
31	2444	10	Wholesale Trade
32	10217	40	Retail Trade
33	1834	9	Banking and Other Finance
34	1978	12	Insurance and Real Estate
35	927	71	Private Household Services
36	1902	23	Business Services
37	744	17	Repair Services
38	1501	30	Personal Services, except Private Household
39	638	37	Entertainment and Recreation Services
40	2675	20	Hospitals

Continued

Table 13.2 (*Continued*)

Industry code	*Number of observations*	*Percentage part-time*	*Industry description*
41	2055	32	Health Services, except Hospitals
42	5542	26	Educational Services
43	952	33	Social Services
44	1910	20	Other Professional Services
45	108	13	Forestry and Fisheries
46	3329	6	Public Administration

Source: Authors' computations from the March 1984 *CPS* extract (see the text for sample restrictions).

Table 13.3
Estimated Part-time/Full-time Wage Differentials: by Industry

Industry code	*RW1*	*RW2*	*RW3*	*RW4*
1	−.199 (2.0)	−.229 (2.1)	−.155 (1.6)	−.179 (1.7)
2	−.201 (1.9)	−.259 (1.6)	−.148 (1.4)	−.153 (0.9)
3	−.153 (4.7)	−.196 (4.8)	−.147 (4.5)	−.185 (4.5)
4	−.077 (0.8)	−.281 (2.2)	−.068 (0.7)	−.288 (2.2)
5	−.002 (0.0)	−.047 (0.2)	−.107 (0.2)	−.011 (0.1)
6	−.147 (1.0)	−.012 (0.7)	−.131 (1.0)	−.086 (0.6)
7	.142 (1.3)	.125 (0.8)	.145 (1.3)	.139 (0.8)
8	−.031 (0.3)	−.066 (0.6)	−.033 (0.3)	−.077 (0.7)
9	[a]	[a]	[a]	[a]
10	−.200 (2.1)	−.143 (1.4)	−.171 (1.8)	−.110 (1.1)
11	−.238 (2.6)	−.159 (1.5)	−.230 (2.6)	−.141 (1.3)
12	−.044 (0.3)	−.029 (0.2)	−.072 (0.6)	−.052 (0.3)
13	−.213 (1.7)	−.306 (2.1)	−.233 (1.8)	−.309 (2.1)
14	−.302 (1.5)	−.566 (2.4)	−.264 (1.4)	−.530 (2.3)
15	−.326 (3.1)	−.273 (2.4)	−.330 (3.2)	−.260 (2.2)
16	.057 (0.2)	.057 (0.2)	.109 (0.3)	.109 (0.3)
17	−.321 (2.5)	−.414 (2.9)	−.338 (2.6)	−.435 (3.0)
18	−.170 (2.9)	−.289 (4.1)	−.177 (3.0)	−.296 (4.2)
19	[a]	[a]	[a]	[a]
20	−.069 (1.0)	−.198 (1.9)	−.071 (1.0)	−.203 (1.9)
21	−.165 (3.0)	−.180 (2.4)	−.178 (3.2)	−.184 (2.5)
22	−.245 (1.6)	−.419 (2.3)	−.211 (1.4)	.322 (1.8)
23	−.306 (4.1)	−.290 (3.8)	−.269 (3.6)	−.240 (3.1)
24	−.083 (0.8)	−.115 (0.9)	−.054 (0.5)	−.077 (0.6)
25	.444 (1.5)	.444 (1.5)	.361 (1.2)	.361 (1.2)
26	−.091 (1.0)	−.081 (0.8)	−.081 (0.9)	−.045 (0.4)
27	−.058 (0.6)	−.064 (0.5)	−.049 (0.5)	−.046 (0.4)
28	−.226 (6.2)	−.237 (5.9)	−.203 (5.5)	−.209 (5.2)
29	−.298 (3.7)	−.277 (3.3)	−.261 (3.3)	−.250 (3.0)
30	−.415 (4.6)	−.428 (4.7)	−.349 (3.8)	−.363 (3.9)
31	−.172 (3.2)	−.168 (3.0)	−.157 (2.9)	−.151 (2.7)

Continued

Table 13.3 (*Continued*)

Industry code	*RW1*	*RW2*	*RW3*	*RW4*
32	−.123 (7.3)	−.117 (6.8)	−.088 (5.2)	−.081 (4.6)
33	−.086 (1.9)	−.078 (1.7)	−.065 (1.5)	−.056 (1.3)
34	−.300 (4.7)	−.264 (4.1)	−.266 (4.2)	−.299 (3.5)
35	.323 (4.7)	.283 (4.1)	.330 (4.8)	.294 (4.3)
36	−.213 (5.0)	−.211 (4.9)	−.186 (4.3)	−.185 (4.2)
37	−.346 (3.8)	−.370 (3.7)	−.337 (3.7)	−.358 (3.4)
38	−.065 (1.3)	−.050 (0.9)	−.052 (1.0)	−.032 (0.6)
39	−.111 (1.4)	−.102 (1.3)	−.093 (1.1)	−.085 (1.1)
40	−.047 (1.8)	−.041 (1.6)	−.034 (1.4)	−.027 (1.1)
41	−.052 (1.5)	−.048 (1.3)	−.040 (1.1)	−.036 (1.0)
42	−.177 (9.1)	−.179 (9.1)	−.150 (7.6)	−.151 (7.6)
43	−.124 (2.5)	−.127 (2.5)	−.107 (2.1)	−.110 (2.1)
44	−.247 (4.0)	−.200 (3.1)	−.192 (3.0)	−.141 (2.2)
45	−.369 (2.1)	−.227 (1.1)	−.358 (1.9)	−.264 (1.3)
46	−.255 (7.1)	−.263 (7.0)	−.239 (6.6)	−.246 (6.6)

Notes: Absolute value of t statistics in parentheses.
"Sample size too small to compute difference.

Correlations:				
	RW2	.91		
	RW3	.98	.90	
	RW4	.90	.98	.90
		RW1	*RW2*	*RW3*

Source: Authors' computations from within-industry wage equation estimates. *RW1* and *RW3* are based on part-time status in the survey week, while *RW2* and *RW4* are based on usual part-time status. *RW1* and *RW2* are from models without interaction terms, while *RW3* and *RW4* are from models that interact gender with number of children, marital status, education and experience. See the text for details.

Most striking, the estimated differentials are primarily negative (part-time workers *do* get paid less) and vary widely across industries. For example, 40 of the 44 *RW1* coefficients are negative, with 25 of these being statistically significantly different from zero. The mean of the negative differentials is about −.18, with the largest close to −.41. All of the estimated positive differentials are statistically insignificantly different from zero, save for that for private household workers. As noted earlier, because of the unique nature of this industry, we exclude it from the inter-industry analyses reported in the next section.

The March 1984 *CPS* data also contained information on whether an employee was covered by a health insurance plan to which an employer contributed and on whether the employee was covered by a pension plan. As a result, it was possible for us to estimate equations similar to equation (13.1) in which the dependent variables were dichotomous (1.0) variables indicating an employee's coverage under these types of plans. Estimates of the coefficients of P_j in these linear probability

Table 13.4
Estimated Part-time/Full-time Fringe Differentials: By Industry (absolute value of t statistics)

Industry code	*Employer contributes to a health insurance plan for the employee (1,0)* RH1		RH2		*Employee is covered by a pension plan (1,0)* RP1		RP2	
1	−.170	(4.6)	−.172	(4.3)	−.078	(3.1)	−.075	(2.7)
2	−.320	(4.6)	−.559	(5.2)	−.381	(4.3)	−.575	(4.3)
3	−.265	(10.8)	−.321	(10.3)	−.181	(7.6)	−.228	(7.5)
4	−.246	(3.1)	−.351	(3.3)	−.069	(0.7)	−.264	(2.3)
5	.020	(0.2)	.025	(0.2)	.018	(0.2)	−.153	(0.9)
6	−.449	(3.9)	−.509	(4.0)	−.349	(2.5)	−.446	(2.9)
7	−.240	(3.2)	−.514	(4.7)	−.111	(1.1)	−.271	(1.8)
8	−.130	(1.9)	−.228	(2.3)	−.193	(2.2)	−.287	(2.4)
9	[a]		[a]		[a]		[a]	
10	−.397	(7.8)	−.447	(8.2)	−.278	(4.0)	−.288	(3.9)
11	−.455	(7.3)	−.503	(6.8)	−.350	(4.6)	−.376	(4.1)
12	−.116	(1.2)	−.296	(2.4)	.048	(0.4)	.191	(1.2)
13	−.343	(3.8)	−.320	(3.1)	−.478	(3.6)	−.399	(2.6)
14	−.161	(1.6)	−.271	(2.3)	−.438	(3.6)	−.448	(3.2)
15	−.271	(2.9)	−.297	(3.0)	−.280	(2.4)	−.206	(1.6)
16	−.248	(1.1)	−.0.94	(0.3)	−.077	(0.3)	−.085	(0.3)
17	−.280	(3.1)	−.383	(3.9)	−.175	(1.9)	−.171	(1.7)
18	−.247	(5.9)	−.312	(6.3)	−.212	(4.3)	−.306	(5.3)
19	[a]		[a]		[a]		[a]	
20	−.135	(1.7)	−.133	(1.1)	−.106	(1.1)	−.124	(0.9)
21	−.105	(2.0)	−.199	(2.9)	−.095	(1.9)	−.142	(2.1)
22	−.398	(4.6)	−.496	(4.5)	−.338	(2.9)	−.517	(3.4)
23	−.410	(10.4)	−.398	(9.7)	−.274	(6.6)	−.274	(6.3)
24	−.356	(4.8)	−.386	(4.6)	−.413	(4.3)	−.463	(4.3)
25	−.463	(2.3)	−.463	(2.3)	−.426	(1.2)	−.426	(1.2)
26	−.269	(3.8)	−.265	(3.2)	−.257	(3.0)	−.264	(2.6)
27	−.188	(2.0)	−.289	(2.1)	−.088	(1.0)	−.036	(0.8)
28	−.384	(14.3)	−.421	(14.4)	−.290	(9.7)	−.306	(9.3)
29	−.441	(9.2)	−.415	(8.1)	−.367	(5.9)	−.359	(5.4)
30	−.610	(9.7)	−.598	(9.3)	−.647	(8.4)	−.642	(8.3)
31	−.389	(12.1)	−.408	(12.2)	−.262	(7.2)	−.270	(7.1)
32	−.265	(25.4)	−.266	(25.0)	−.121	(13.0)	−.120	(12.7)
33	−.465	(13.5)	−.472	(13.6)	−.257	(6.2)	−.257	(6.1)
34	−.446	(13.4)	−.438	(13.0)	−.269	(7.6)	−.273	(7.6)
35	−.037	(2.2)	−.050	(3.0)	−.019	(1.3)	−.027	(1.9)
36	−.381	(14.3)	−.368	(13.4)	−.206	(7.7)	−.193	(7.0)
37	−.334	(6.3)	−.287	(4.8)	−.104	(2.3)	−.098	(1.9)
38	−.240	(8.6)	−.244	(8.5)	−.118	(5.0)	−.112	(4.5)
39	−.374	(8.9)	−.382	(9.0)	−.187	(4.6)	−.190	(4.7)
40	−.348	(15.5)	−.361	(15.8)	−.298	(12.0)	−.302	(12.0)
41	−.312	(13.5)	−.309	(13.2)	−.190	(8.6)	−.189	(8.4)
42	−.419	(29.1)	−.423	(29.3)	−.396	(27.9)	−.397	(27.9)

Continued

Table 13.4 (*Continued*)

Industry code	*Employer contributes to a health insurance plan for the employee (1,0)*		*Employee is covered by a pension plan (1,0)*	
	*RH*1	*RH*2	*RP*1	*RP*2
43	−.342 (10.7)	−.352 (10.9)	−.263 (8.4)	−.263 (8.4)
44	−.370 (12.8)	−.360 (12.1)	−.239 (7.9)	−.233 (7.5)
45	−.544 (3.9)	−.543 (3.6)	−.570 (4.6)	−.580 (4.2)
46	−.411 (13.1)	−.425 (13.2)	−.429 (15.0)	−.428 (14.5)

Notes: Absolute value of *t* statistics in parentheses.
[a]Sample size too small to compute differentials.

Correlations: (*RH*1, *RH*2) = .85 (*RP*1, *RP*2) = .90

Source: Authors' computations obtained from within-industry employer contribution to health insurance and employee pension plan coverage equations. *RH*1 and *RP*1 (*RH*2 and *RP*2) are based on survey week (usual) part-time status. Results are for models without gender interactions; differentials from models with such interactions were correlated at .97 or higher level. See text for details.

function models will thus indicate the differential in the probability that a part-time worker was covered by these plans, *ceteris paribus*.

The estimates of these coefficients appear in Table 13.4. To conserve space, only the estimates from models without interaction terms are presented (estimates from the models that included interaction terms were very highly correlated with these). In all cases these estimates are negative and in virtually all cases statistically significantly so. Part-time employees do have lower probabilities, *ceteris paribus*, of being covered by a health insurance plan that an employer contributes to or by a pension plan. The mean part-time/full-time differentials across industries in these two probabilities are −.31 and −.25, respectively. As Table 13.4 indicates, however, the probabilities vary widely across industries.

Finally, for approximately one-quarter of the sample, the March 1984 *CPS* contained information on whether the individual was a union member.[11] For the subsample of individuals for which this information was present, one can estimate equations similar to equation (13.1), using a dichotomous (1, 0) variable for union membership as the dependent variable. The estimated coefficient of P_j in each industry from these regressions will be an estimate of the differential in the probability of

Notes: Absolute value of *t* statistics in parentheses.
[a]Sample size too small to compute meaningful differentials.
[f]Magnitude of coefficient too large to be believable and probably due to the very small number (4) of part-time workers in the sample.

Source: Authors' computations obtained from within-industry probability of union membership equations. *RU*1 (*RU*2) based on survey week (usual) part-time status. See the text for details.

Table 13.5
Estimated Union Membership Proportion and Part-time/Full-time Probability of Membership Differentials: by Industry

Industry code	*Number of observations with membership data*	*Proportion/number union/union*	*Estimated differentials* RU1	RU2
1	176	.03/5	−.026 (0.8)	−.026 (0.8)
2	172	.20/34	−.254 (1.3)	−.118 (0.3)
3	672	.26/175	−.145 (2.9)	−.215 (3.2)
4	113	.29/33	−.221 (0.9)	−.122 (0.4)
5	64	.13/9	−.128 (0.7)	[a]
6	85	.29/25	−.340 (1.5)	−.340 (1.5)
7	132	.45/59	.665(2.1)[b]	−.051 (0.0)
8	164	.26/43	−0.66 (0.4)	−.092 (0.5)
9	—	—	[a]	[a]
10	400	.20/80	−.089 (0.6)	−.228 (1.4)
11	309	.20/62	.043 (0.3)	−.047 (0.3)
12	161	.57/92	.233 (0.7)	.066 (0.1)
13	83	.25/21	−.834 (2.5)	−.718 (1.8)
14	108	.20/22	−.091 (0.4)	−.091 (0.4)
15	84	.08/7	−.122 (0.4)	−.122 (0.4)
16	19	.26/5	[a]	[a]
17	56	.23/13	−.135 (1.3)	−.361 (1.5)
18	253	.36/91	.038 (0.4)	−.110 (1.0)
19	8	.63/5	[a]	[a]
20	94	.17/16	−.199 (1.4)	−.267 (1.3)
21	171	.25/43	.059 (0.6)	−.105 (0.7)
22	95	.54/51	−.379 (1.3)	−.681 (2.0)
23	255	.12/31	−.070 (1.0)	−.066 (2.0)
24	168	.17/29	−.248 (1.1)	−.206 (0.8)
25	28	.36/10	[a]	[a]
26	97	.18/17	.192 (1.4)	.136 (0.8)
27	33	.21/7	[a]	[a]
28	565	.44/249	−.187 (3.0)	−.234 (3.5)
29	239	.43/103	−.155 (0.9)	−.157 (0.9)
30	208	.32/67	−.116 (0.6)	−.261 (1.3)
31	520	.08/42	−.032 (0.7)	−.023 (0.5)
32	2310	.07/162	.007 (0.5)	.004 (0.3)
33	463	.02/9	−.023 (1.0)	−.023 (1.0)
34	420	.05/21	−.048 (1.4)	−.047 (1.4)
35	191	.01/2	−.015 (1.1)	−.015 (1.1)
36	414	.04/16	−.029 (1.2)	−.029 (0.8)
37	152	.06/9	−.060 (0.9)	−.076 (1.0)
38	346	.10/35	−.118 (3.1)	−.135 (3.5)
39	153	.14/21	−.035 (0.5)	−.109 (0.2)
40	688	.15/103	−.067 (1.8)	−.056 (1.4)
41	439	.08/35	−.017 (0.6)	−.024 (0.8)
42	1369	.32/438	−.224 (7.1)	−.224 (7.1)
43	229	.09/21	−.125 (3.0)	−.125 (3.0)
44	439	.05/22	−.007 (0.3)	−.002 (0.1)
45	30	.00/0	[a]	[a]
46	780	.26/203	−.105 (1.6)	−.112 (1.8)

union membership for part-time and full-time workers in the industry, *ceteris paribus*.

These estimated coefficients for each industry, as well as the size of the subsample available to conduct the analyses, and the proportion of employees who are union members in the subsample are found in Table 13.5. While over 80 per cent of the estimated differentials are negative, only 8 are statistically significant, perhaps because of the relatively small sample sizes.

Table 13.6 summarizes the patterns of signs and statistical significance of the part-time/full-time employee wage, fringe benefit and unionization differentials that we have obtained in this section. The fringe benefit and, arguably, wage differentials appear to be sufficiently precisely estimated to use as inputs in the inter-industry analyses that will be presented in the next section. The union membership differentials clearly are not, however, and we omit them from the subsequent analyses.

In concluding this section, it is nevertheless interesting to ask whether government policy in the United States has made it increasingly difficult for part-time workers to join unions in recent years? If so, this might encourage employers to increase their usage of part-time employees.

Union elections in the private sector in the United States are governed by the National Labor Relations Board (NLRB) and it is to the NLRB that disputes over whether part-time employees should be included in

Table 13.6
Patterns of Estimated Differentials

Equation	*Number of differentials estimated*	*Number that are negative*	*Number negative and statistically significant*
Wage			
*RW*1	44	40	25
*RW*2	44	40	27
*RW*3	44	40	23
*RW*4	44	40	24
Health insurance			
*RH*1	44	43	41
*RH*2	44	43	41
Pension			
*RP*1	44	42	35
*RP*2	44	42	38
Union membership			
*RU*1	41	33	8
*RU*2	42	36	8

Source: Authors' computations from coefficient estimates in Tables 13.3, 13.4 and 13.5.

proposed bargaining units and thus eligible to vote in union elections are brought. The *stated* policy of the NLRB has remained roughly constant over time; it has always attempted to determine bargaining units on the basis of a 'community of interest'.[12] That is, employees who share similar interests in wages, hours and other conditions of employment are placed in the same bargaining unit.

Part-time employees are generally included in a bargaining unit with full-time employees whenever the part-time employees perform work within the unit on a regular basis, for a sufficient period of time, during an appropriate calendar period. In determining in a specific case whether part-time employees share a sufficient community of interest to be placed in the same bargaining unit as full-time employees, the NLRB considers a number of factors including: the similarity (with full-time employees) and regularity of part-time employees' hours of work, the similarity (with full-time employees) of part-time employees' wage and benefit packages, common supervision for both types of employees, the similarity of their qualifications, training skills and job functions, the frequency of their contact and interchange while performing their job duties, the history of collective bargaining in the firm and the extent of union organization, the desires of both part-time and full-time employees, and the organizational structure of the firm.

Of course, to say that the *stated* policy of the NLRB has remained roughly constant over time is not to say that the *actual* policy has remained constant. To investigate if the latter has changed, we searched through NLRB decisions during the 1976–1984 period, finding 52 that dealt with part-time workers. About 60 per cent of these decisions resulted in part-time workers being included in a larger bargaining unit or allowed to set up their own unit for a bargaining election.

Given the small number of decisions each year (four to five), not surprisingly the proportion of times part-time workers were included in other units or allowed to set up their own unit fluctuated from year to year. None the less, to see if there were any trends in board decisions during the period, we estimated linear probability function models in which a dichotomous ((1, 0) part-time worker included or allowed to set up own unit) variable was regressed on a time-trend term. No significant trend showed up in the data even when higher order trend terms (i.e. a quadratic term) were included to allow for non-linearities. We thus found no evidence that changes in NLRB policy in recent years have encouraged the growth of part-time employees relative to full-time employees. That is, it does *not* appear that the NLRB is increasingly making it more difficult for part-time workers in the United States to join unions.

13.4 Inter-industry Variations in Part-time Employment in the United States

Given that estimates of the part-time employee/full-time employee

wage (RW), at least partially employer-financed health insurance coverage probability (RH), and private pension coverage probability (RP) differentials vary widely across industries, our goal in this section is to see if the variations in these differentials help to explain the pattern of inter-industry variations in part-time employment in the United States. To do this, we embed them in a simple model of the demand and supply of part-time employment.

On the demand side, the ratio of part-time to full-time employees (E_P/E_F) is postulated to be a function of the relative costs of the two groups, as measured by the above differentials, and the production technology in use in an industry. Since the latter is not directly observable, it is proxied by a vector of variables (Y) that indicate the share of an industry's workforce in each of seven major occupational groups. A negative value for each of the differentials indicates a cost advantage for part-time workers; the larger (in absolute value) the differential is, the greater the cost saving from part-time workers. Thus, we expect the coefficients of RW, RH, and RP all to be negative.

$$(E_P/E_F) = D(R\bar{W}, R\bar{H}, R\bar{P}, Y). \tag{13.2}$$

On the supply side, the larger in absolute value these differentials are the less attractive is the relative reward to being a part-time worker and thus the smaller the fraction of workers who will want to work part time. The relative supply of part-time workers will also depend upon the characteristics of workers 'attached' to the industry (13.2). For example, *ceteris paribus*, married women with children, students and older workers phasing into retirement may all find part-time employment attractive. Similarly, unions may try to discourage their members from working part time if they perceive that widespread use of part-timers may discourage new workers from joining unions. Thus, we have on the supply side

$$(E_P/E_F) = S(R\dot{W}, R\dot{H}, R\dot{P}, Z). \tag{13.3}$$

Linear versions of the system in (13.2) and (13.3) were estimated using the estimates of RW, RH and RP obtained in the last section, and mean values (by industry) of the other variables obtained from the May 1984 *CPS*. The analyses reported below use 43 observations, dropping only the two industries whose sample sizes were too small to estimate the part-time/full-time wage differentials (see Table 13.3) and also the private household services industry (industry 35). Restricting the sample further to only those industries for which we estimated negative values of RW, or still further to those for which these estimates were negative and statistically significant, did not lead to more precise estimates of the model.

Table 13.7 contains estimates of the model when survey week part-time status is used to classify workers, and $RW1$, $RH1$ and $RP1$ are used as explanatory variables. The complete list of other demand (Y) and

supply (Z) side variables included in the model is found in this table, along with their definitions. All results presented in the table are for unweighted regressions; weighting by the square root of the number of individuals in each underlying industry sample (Table 13.2) did not appreciably change the sign pattern or significance of the coefficients.

Column 1 presents OLS estimates of the structural demand curve. While the relative wage differential variable performs as expected, with larger part-time employee wage rate savings leading to increased use of part-time employees, the coverage by pension plan differential perversely appears to have a positive coefficient, implying that the less likely it is that part-time employees are covered by a pension, the fewer part-time employees will be employed. The pattern of occupational share coefficients suggests that industries that employ a relatively large number of blue-collar skilled workers (the omitted reference group in the equation) also tend to employ relatively few part-time employees.

Might the above pattern of results be affected by simultaneous equations bias? To answer this question, the demand (and supply) model is re-estimated by 2SLS. Columns 3, 4 and 5 in Table 13.7 present, respectively, the instrumental variable estimates we obtained for the wage, health insurance coverage probability and pension coverage probability differentials. The explanatory power of the wage differential equation is very low; indeed no individual coefficient is statistically significant. The health insurance and pension coverage probability differential equations are somewhat better. It is interesting to note that females and veterans are both less likely to be covered by either type of plan; females possibly because of coverage under other family members' plans and veterans possibly because of their access to medical care and retirement benefits through veterans' administration programs. Union membership, however, increases the probability of pension coverage, as does the average number of children in each family.

The structural demand and supply equations appear in columns 5 and 6 of Table 13.7. The 2SLS demand estimates in 5 are very similar to the OLS estimates, save that they are slightly less significant. The only cost differential that is significant in the supply curve is the wage differential; as expected, greater (more negative) part-time/full-time wage differentials lead to relatively fewer employees wanting to work part-time. The coefficients of the personal characteristics variables confirm that an increased number of children per worker and an increased percentage of workers who are students both increase the likelihood that employees will want to work part-time. Increases in the percentage of workers who are union members, however, have only an insignificant negative effect on part-time employment.

To assure the reader that the effects of the part-time/full-time wage, health insurance coverage probability and pension coverage probability differentials that we obtained are not unique to our usage of $RW1$, $RH1$ and $RP1$, Table 13.8 presents the coefficients of the differentials in the relative demand equations that we obtained when we used the other

Table 13.7
Inter-industry Cross-Section Regressions

Variable	(1) OLS *PTNOW*	(2)	(3) *Instruments*	(4)	(5) 2SLS	(6)
		*RW*1	*RH*1	*RP*1	*PTNOW*	*PTNOW*
*RW*1	−12.022 (2.0)					
*RH*1	−7.137 (0.6)					
*RP*1	29.566 (3.1)					
*RW*1					−15.825 (1.3)	36.469 (2.5)
*RH*1					−5.532 (0.3)	−28.817 (1.4)
*RP*1					32.450 (2.8)	10.951 (0.8)
PROF	.279 (2.7)	−.008 (0.8)	.000 (0.0)	−.004 (0.7)	.318 (2.6)	
SALES	.298 (2.7)	−.002 (0.3)	.002 (0.4)	.003 (0.7)	.283 (2.4)	
ADS	−.154 (1.3)	−.005 (1.1)	−.002 (0.6)	.003 (1.2)	−.140 (1.0)	
SERV	.545 (6.4)	.002 (0.5)	.001 (0.2)	.003 (0.9)	.543 (6.1)	
AGE	.204 (2.9)	−.001 (0.1)	−.003 (2.1)	−.005 (2.7)	.211 (2.9)	
BCU	.433 (1.3)	−.000 (0.0)	−.013 (1.9)	−.018 (2.6)	.553 (1.5)	
FEM		−.041 (0.1)	−.668 (1.6)	−1.513 (3.8)		2.723 (0.2)
CHILD		−.112 (0.3)	.232 (0.9)	.572 (2.3)		22.445 (2.7)
AGE		−.031 (0.6)	.018 (0.6)	.031 (1.0)		2.560 (2.2)
MNOW		.016 (1.7)	.000 (0.1)	−.008 (1.5)		−.969 (3.1)
RACE		−.007 (0.5)	−.008 (1.1)	−.008 (1.0)		−0.83 (0.3)
VET		−.010 (0.7)	−.014 (1.7)	−.036 (4.3)		−.011 (0.0)
HISP		−.003 (0.3)	.004 (0.7)	−.009 (1.6)		.074 (0.2)
EDUC		.118 (0.8)	−.055 (0.6)	−.014 (0.2)		−.047 (0.0)
STUD		.005 (0.2)	−.005 (0.4)	−.021 (1.6)		1.399 (3.0)
*A*55		.013 (0.5)	−.001 (0.0)	.002 (0.1)		−.658 (1.3)
UN		.002 (0.5)	.000 (0.0)	.004 (2.0)		−.145 (1.5)
R^2	.728	−.154	.415	.598	—	—

where:

PTNOW = percent of industry employees employed part-time last week
*RW*1 = estimated part-time/full-time wage differential (Table 13.3)
*RH*1 = estimated part-time/full-time employer contributes to a health insurance plan differential (Table 13.4)
*RP*1 = estimated part-time/full-time employee is covered by an employers pension plan differential (Table 13.4)
*RW*1 = instrumental variable estimate for *RW*1
*RH*1 = instrumental variable estimate for *RH*1
*RP*1 = instrumental variable estimate for *RP*1
PROF = percentage of workers in the industry who are professionals
SALES = sales
ADS = administrative support
SERV = service
AGF = agriculture or farm
BCU = blue-collar unskilled

} omitted group is blue-collar skilled

FEM = fraction of workers who are female
CHILD = average number of children per worker
AGE = mean age
MNOW = percentage of workers married now
RACE = percentage of workers who are White
VET = percentage of workers who are veterans
HISP = percentage of workers with Spanish surnames
EDUC = mean years of education
STUD = percentage of workers who are students
A55 = percentage of workers age 55 and older
UN = percentage of workers who are union members

Note: Absolute value of t statistics in parentheses.
$n = 43$ for all equations.

Table 13.8
Relative Price Coefficients from the Part-time Worker Employment Share Equations: Various Specifications

	Part-time last week			*Part-time usual*	
	OLS	*2SLS*		*OLS*	*2SLS*
*RW*1	−12.022 (2.0)	−15.825 (1.3)	*RW*2	−7.785 (1.8)	−5.964 (0.8)
*RH*1	−7.137 (0.6)	−5.532 (0.3)	*RH*2	−3.260 (0.4)	−5.908 (0.4)
*RP*1	29.565 (3.1)	32.450 (2.8)	*RP*2	22.910 (3.5)	33.653 (3.3)
*RW*3	−14.273 (2.2)	−18.807 (1.4)	*RW*4	−7.192 (1.5)	−3.409 (0.4)
*RH*3	−12.252 (1.1)	−24.981 (1.2)	*RH*4	−1.546 (0.2)	−4.871 (0.3)
*RP*3	34.846 (3.5)	43.894 (3.2)	*RP*4	21.577 (3.2)	33.451 (2.8)

Note: Absolute value of t statistics in parentheses.
Source: Regressions in Table 13.7 and analogous ones for other variable specifications.

estimates of the differentials (i.e. *RW*2, *RW*3, *RW*4, ...). As can be seen there, the pattern of coefficients is very similar across all four specifications, although the specifications based on part-time employment in the survey week 'perform' better than those based on usual part-time employment.

13.5 Concluding Remarks

Our analyses of the aggregate time-series data for the United States suggest that there has been a tendency towards increased employment of part-time workers in the United States in recent years, a trend that is observed *after* one controls for cyclical factors. Moreover, this trend has come from an increase in 'involuntary' part-time employment, not from an increase in voluntary part-time employment. Searches for explanations for the recent growth of part-time employment in the US should therefore focus on the demand side of the labor market.

Such a search led us to ask if a growing cost differential between part-time and full-time employees might provide part of the explanation. We addressed this issue by focusing on inter-industry variations in the part-time employment/full-time employment ratio and seeing if variations in the relative cost differential across industries could help explain this part-time/full-time employment variation. In fact, relative wage costs did appear to influence relative employment levels, as predicted, on the demand side of the market.

In contrast, the larger the differential between the probability of pension coverage for full-time and part-time workers in an industry, the smaller the relative demand for part-time employees tended to be. At first glance this result seems inconsistent with our model. However,

upon reflection, it may make sense. It is well known that pension coverage tends to reduce turnover and increase employees' expected tenure with firms (see, for example, Mitchell, 1983). The additional costs of pension coverage for part-time employees may be offset by savings in turnover and training costs if in fact this coverage induces them to have longer job tenure, making part-time employees more, rather than less, attractive to employers. To begin to test if this is occurring, one would want to see if the expected job tenure of part-time workers, by industry, is correlated, *ceteris paribus*, with the probabilities of pension coverage that we have estimated. Sadly, however, job tenure data are not available in the March 1984 *CPS*.

Finally, it is worth noting that the relative cost of part-time workers influences the relative supply of them (vis-à-vis full-time workers) as well as the relative demand. Indeed, our estimates, at least for the relative wage cost variable, suggest that supply responses exceed demand responses. Of course, given that some part-time workers are 'involuntarily' part-time, it is not obvious that the structural demand and supply model we have estimated is an entirely appropriate one.

Notes

1 For example, evidence that part-time workers earn less than full-time workers, either in raw data or, more appropriately, after controlling for personal characteristics, is found in Ballard (1984) for the United Kingdom, Labor Canada (1983) for Canada, and Owen (1979) and Parsons (1974) for the United States. Similarly, evidence on part-time workers' poorer access to fringe benefits, such as health insurance, vacations, sick leave and private retirement plans, are found in Ballard (1984) for the United Kingdom, Labor Canada (1983) for Canada, Nakakubo (1985) for Japan, and Ichniowski and Preston (1986) for the United States. Finally, Disney and Szyszczak (1984) discuss how coverage of part-time workers under various social insurance programs and protective labor legislation has varied over time.

2 Disney and Szyszczak (1984) show that employment of part-time workers in Great Britain expanded most rapidly in periods when they were covered by fewer employer-financed social insurance programs and less protective labor legislation.

3 See Owen (1979) for an earlier effort in this direction. While Owen had estimates of relative wage cost differences, he had no data on fringe benefits.

4 A similar approach was used by Ehrenberg and Schumann (1982) in investigating the growth of overtime hours in the United States.

5 Evidence on the part-time employment–union membership relationship is very weak. For example, in the United Kingdom, Bain and Elsheikh (1979) and Richardson and Catlin (1979) found no strong relationship between part-time employment ratios and unionization percentages across industries. Similarly, Dickens (1983) found in a sample of roughly 1,000 workers who voted in 31 union elections in the United States in the early 1970s that, *ceteris paribus*, part-time workers' were some 6–7 per cent less likely to vote for a union, although this relationship was not statistically significant. Somewhat surprisingly, virtually all studies seeking to explain the well-publicized decline in unionization in the United States have failed to

consider if the growth of part-time employment has played any role. (See, for example, Dickens and Leonard, 1985, and their bibliography.)

6 The latter undoubtedly owing to cutbacks in financial aid for college students that increasingly forced college students in the US to work to help finance their education. For evidence on the increasing hours of work of college students, see Ehrenberg and Sherman (1987).

7 A quick look at Table 13A.1, however, will caution the reader that whether or not one observes an apparent trend may depend heavily on the starting date one chooses. More on this point below.

8 Similar results are reported in Ichniowski and Preston (1986) who use monthly data over the 1973–83 period. Both their results and ours fail to control for minimum wage changes. Matilla (1981) provides some evidence that increases in the minimum wage are associated with increases in the part-time/full-time employment ratio of teenagers, while Ehrenberg and Marcus (1982) find the opposite, at least for teens from low-income families.

9 The percentage is actually as high as 71 per cent for private household service workers. However, because the 'employers' in this industry are typically private individuals (not firms) and most employees work for a number of different people in any one week, we will ignore data from this industry in most of what follows.

10 To be a bit more precise, given two otherwise identical individuals except for their part-time status, $a_{pi} = \log(W_P/W_F)$, where the subscripts P and F refer to part-time and full-time workers respectively. Consequently, the proportional a wage differential of part-time workers is $(W_P - W_F)/W_F = e^{a_{pi}} - 1$. For small a values of $a_{pi} \cdot a_{pi} \approx e^{a_{pi}} - 1$.

11 The *CPS* consists of eight 'rotation groups' and only two of the groups were asked about union membership.

12. The material in this paragraph and the next two are drawn from Morris (1983) and Nash and Blake (1979).

Table 13A.1
Percentage of Part-time Employees in the United States Economy, 1963–84

Year	*All individuals* 16+	16–19	*Males* 20+	*Females* 20+
1963	10.7	37.8	3.6	19.5
1964	11.0	40.5	3.7	19.5
1965	11.2	40.7	3.5	19.3
1966	11.7	41.2	3.6	19.6
1967	12.4	44.1	4.0	20.3
1968	12.8	44.9	4.1	20.8
1969	13.3	46.1	4.4	20.9
1970	13.7	47.0	4.7	21.5
1971	13.9	47.8	4.7	21.6
1972	14.0	46.2	4.8	21.6
1973	14.0	44.3	4.8	21.6
1974	14.0	44.1	4.8	21.5
1975	14.3	46.4	5.0	21.4
1976	14.3	46.5	4.9	21.4
1977	14.3	45.6	5.0	21.2
1978	14.3	45.6	5.0	20.9
1979	14.2	46.1	4.9	20.7
1980	14.4	47.7	5.1	20.5
1981	14.2	48.4	5.0	20.5
1982	14.3	50.1	5.2	20.4
1983	14.0	49.6	5.3	20.0
1984	13.7	50.2	5.1	19.5

Source: Authors' calculations from data in US Bureau of Labor Statistics (BLS), *Labor Force Statistics Derived from the Current Population Survey: A Databook*, Bulletin 2096 (Washington, DC, September 1982), Table A11 (for 1963–81), and *Employment and Earnings*, various issues (for 1982–84).

Table 13A.2
Percentage of Part-time Employees in Non-agricultural Industries, 1955–84, by Reason for Part-time Status

Year	*Part-time for economic reasons*	*Part-time for non-economic reasons*
1955	3.4	7.1
1956	3.6	7.9
1957	3.9	8.2
1958	5.4	8.3
1959	4.2	8.7
1960	4.5	9.0
1961	4.9	9.3
1962	4.0	9.7
1963	3.8	10.1
1964	3.5	10.5
1965	3.0	10.6
1966	2.5	11.4
1967	2.9	12.0
1968	2.5	12.4
1969	2.6	12.9
1970	3.1	13.3
1971	3.4	13.4
1972	3.3	13.5
1973	3.0	13.5
1974	3.5	13.6
1975	4.6	13.8
1976	4.2	13.8
1977	4.0	13.9
1978	3.8	13.9
1979	3.8	13.8
1980	4.5	13.9
1981	4.9	13.7
1982	6.5	13.8
1983	6.5	13.4
1984	5.7	13.1

Source: Authors' calculations from data in BLS Bulletin 2096, Table A18 (1955–81), *Employment and Earnings*, various issues, thereafter.

Table 13A.3
Percentage of Part-time Employees in Non-agricultural Industries, 1968–84, by Age, Sex, Race and Reason for Part-time Status

Year	*Males 20+*	*Females 20+*	*All 16–19*	*All Whites*	*All non-Whites*
Part-time for economic reasons					
1968	1.7	3.3	5.9	2.1	5.6
1969	1.7	3.2	6.1	2.3	5.2
1970	2.2	3.7	6.9	2.8	6.1
1971	2.4	4.2	8.0	3.1	6.2
1972	2.1	3.9	8.5	3.0	5.7
1973	2.0	3.7	7.4	2.8	5.1
1974	2.4	4.2	7.9	3.2	5.9
1975	3.4	5.2	10.0	4.2	7.3
1976	2.9	4.7	10.0	3.8	6.8
1977	2.7	4.7	9.6	3.8	6.3
1978	2.4	4.5	9.0	3.5	6.1
1979	2.5	4.6	8.4	3.5	5.7
1980	3.3	5.1	9.8	4.2	6.5
1981	3.5	5.6	11.0	4.6	7.2
1982	5.0	7.2	14.2	6.1	9.2
1983	4.8	7.4	15.4	6.1	10.3
1984	4.1	6.6	13.2	5.3	9.5
Part-time for non-economic reasons					
1968	3.6	20.2	45.1	12.5	11.5
1969	3.8	20.3	46.3	13.0	12.0
1970	4.1	20.8	47.3	13.4	11.9
1971	4.1	20.8	47.9	13.6	11.5
1972	4.3	20.8	46.2	13.7	11.8
1973	4.3	20.9	44.2	13.8	11.4
1974	4.3	20.8	44.1	13.8	11.5
1975	4.4	20.7	46.7	14.1	11.9
1976	4.4	20.6	46.5	14.2	10.8
1977	4.6	20.4	45.7	14.3	11.1
1978	4.5	20.1	45.6	14.3	11.1
1979	4.5	19.9	45.9	14.2	10.8
1980	4.7	19.7	47.6	14.3	10.9
1981	4.6	19.6	48.4	14.2	10.1
1982	4.7	19.6	50.0	14.3	9.7
1983	4.8	19.0	49.8	14.0	8.9
1984	4.7	18.6	50.3	13.7	9.1

Source: Authors' calculations from data in BLS Bulletin 2096, Table B22 (1968–81), *Employment and Earnings*, various issues, thereafter.

Table 13A.4
Average Weekly Hours of Part-time Employees in Non-agricultural Industries 1968–84 by Age, Race, and Sex

Year	*All*	*Males 20+*	*Females 20+*	*All 16–19*	*All Whites*	*All non-Whites*
1968	18.6	18.4	19.3	15.6	18.4	19.2
1969	18.2	19.3	19.3	15.8	18.7	19.1
1970	18.3	20.1	19.2	15.9	18.2	19.6
1971	18.6	19.8	19.6	16.0	18.1	19.4
1972	18.6	19.6	19.1	16.5	18.8	19.3
1973	18.4	19.0	19.6	17.0	19.2	19.0
1974	18.8	20.2	19.8	16.9	18.9	19.0
1975	19.3	21.0	19.8	16.8	19.0	19.5
1976	19.2	20.8	19.6	17.1	19.0	19.8
1977	18.8	19.6	19.9	17.4	19.2	18.9
1978	19.1	20.0	19.7	17.4	19.4	19.3
1979	19.3	19.8	20.2	17.6	19.8	19.7
1980	19.7	20.2	20.1	17.6	19.6	20.1
1981	20.0	20.1	20.4	17.3	19.9	20.2
1982	19.6	21.0	20.4	17.0	19.9	20.1
1983	19.9	21.2	20.2	17.1	19.6	20.3
1984	20.2	20.6	20.5	17.6	19.5	20.6

Source: Authors' calculations from data in BLS Bulletin 2096, Table B22 (1968–81), *Employment and Earnings*, various issues, thereafter.

Table 13A.5
Shares of Various Groups in Total and Part-time Employment, – 1968 and 1984

Category	*Year*	*Males 20+*	*Females 20+*	*All 16–19*
Total employment	1968	58.5	33.8	7.6
	1984	52.8	41.0	6.1
Part-time employment	1968	20.7	53.1	26.2
	1984	24.5	54.8	20.6
Part-time employment (economic reasons)	1968	38.3	43.8	17.9
	1984	38.3	47.6	14.1
Part-time employment (non-economic reasons)	1968	17.1	55.0	27.9
	1984	18.6	57.8	23.5

Source: Authors' calculations from data in BLS Bulletin 2096, Table B22 (1968) and *Employment and Earnings.*