Chapter 1

LABOR SUPPLY OF MEN: A SURVEY

JOHN PENCAVEL*

Stanford University and National Bureau of Economic Research

1. Introduction

This survey of male labor supply covers the determinants of whether men work for pay in the labor market and, if so, the determinants of their hours of work. Issues pertaining to the size and structure of the population are not addressed. Also, I shall be concentrating on the work behavior of men prior to their retirement from the labor force.[1] Moreover, even though there are noteworthy investigations into the labor supply of men in many different countries, this survey is restricted almost entirely to the Anglo-American literature. Even with the subject so restricted, there is much material to survey. The economics literature on the determinants of work behavior of men and women is an old one, and during the past 20 years this literature has multiplied many times over as labor supply has become the most active area of all labor economics research. This early and continuing interest in the determinants of market work derives in part from questions of public policy: a century ago these questions concerned regulations on the use of child labor, compulsory school attendance, and restrictions on the length of the working day; more recently, the questions have involved income and commodity taxation, the reform of welfare programs, and movements in productivity.

*A number of friends have helped me prepare this survey. Above all, I am indebted to Angus Deaton for constructive and thorough comments on an incomplete version of this paper and to Tom MaCurdy for his careful criticism of several large portions of this manuscript. I also received comments on particular sections of the paper from Orley Ashenfelter, David Card, John Ham, Mark Killingsworth, Tom Kniesner, and Ian Walker. Jeremy Rudin was an excellent research assistant and the manuscript was typed professionally and cheerfully by David Criswell. To all these people, I am most grateful for their help. Support from the Sloan Foundation to the Department of Economics at Stanford University is gratefully acknowledged.

[1]Edward Lazear's paper in this Handbook (Chapter 5) contains information on retirement.

Handbook of Labor Economics, Volume I, Edited by O. Ashenfelter and R. Layard

Conjectures about whether an increase in remuneration brought forth more work effort can be traced back at least to the mercantile economists,[2] but the careful statement of the issues is much more recent. Robbins (1930) is usually credited with the proposition that constrained utility maximization yields an ambiguous implication about the wage-slope of the labor supply curve although Jevons (1888) was quite explicit on the matter.[3] After pointing out that in the absence of knowledge about the form of the utility function it was impossible to sign the slope of the labor supply curve, Jevons proceeded to cite instances in which the sudden increase in the prices of goods induced greater hours of work and so he surmised that, in fact, the labor supply function was negatively sloped with respect to wages. The first major empirical effort to examine the wage-slope of the labor supply curve[4] was Paul Douglas's *Theory of Wages* (1934). In one chapter drawing upon data collected from the 1920 Census of Manufactures, he regressed for each age–sex group in 38 U.S. cities the employment-to-population ratio on *real* annual earnings in manufacturing industry holding constant the fraction of the city's population who were either foreign-born or black. For men in all age groups he found a negative correlation though only for the very young and the old was this association significantly different from zero. In another chapter he examined both time-series and cross-section (across industries and across states) data on hours of work and hourly earnings and from these he concluded that the elasticity of hours with respect to wages "is in all probability somewhere between -0.1 and -0.2..." (p. 312). In his careful treatment of the data and in his awareness of the problems impeding inferences, Douglas's work is really quite outstanding.[5] After *The Theory of Wages*, the landmarks in the research on labor force participation are as follows: Schoenberg and Douglas (1937); Woytinsky (1940); Durand (1948); Bancroft (1958); and Long (1958).

[2] See the references cited in Douglas (1934, p. 270). Long (1958, p. 40) refers to Sir Edward West's summary in 1826 of evidence presented to Committees of the Houses of Parliament "that the labourer in a scarce year, when his wage will furnish him with a much less than the usual quantity of food, will, in order to attain his usual supply of necessaries, be willing to do much more work than usual, even at a reduced rate of wages".

[3] Jevons wrote (1888, pp. 179–180): "Supposing that circumstances alter the relation of produce to labour, what effect will this have upon the amount of labour which will be exerted? There are two effects to be considered. When labour produces more commodity, there is more reward, and therefore more inducement to labour. If a workman can earn ninepence an hour instead of sixpence, may he not be induced to extend his hours of labour by this increased result? This would doubtless be the case were it not that the very fact of getting half as much more than he did before, lowers the utility to him of any further addition. By the produce of the same number of hours he can satisfy his desires more completely; and if the irksomeness of labour has reached at all a high point, he may gain more pleasure by relaxing that labour than by consuming more products. The question thus depends upon the direction in which the balance between the utility of further commodity and the painfulness of prolonged labour turns. In our ignorance of the exact form of the functions either of utility or of labour, it will be impossible to decide this question in an a priori manner...."

[4] Earlier though more casual empirical work appears in Frain (1929) and Teper (1932).

[5] Rees (1979) provides a modern perspective on Douglas' labor supply research.

With respect to hours of work, there is the work of Lewis (1957), Bry (1959), Jones (1961), and Finegan (1962).

Modern research on labor supply is characterized by a more careful attempt to separate the measurement of income from substitution effects. It dates from Mincer's (1962) paper on the labor force participation rate of married women and Kosters' (1966) dissertation on the hours worked by men. Since the mid-1960s, progress in computing technology – especially the development of more efficient methods of storing on magnetic tapes and processing information on individuals and the enormous reduction in the costs of applying multivariate statistical techniques to these data – has resulted in a vast outpouring of empirical research in labor supply. This literature has already been the subject of a number of very good surveys: Heckman and MaCurdy (1981); Heckman, Killingsworth and MaCurdy (1981); Keeley (1981); and Killingsworth (1981, 1983). Each of these tends to be a survey of the economics literature. This survey strives to be a little different, namely a survey of the topic and our knowledge of it as well as what economists have written about it. This is why I have devoted an important part, Section 2, to a summary statement of the major empirical regularities in male labor force participation and male hours of work. It is these and other regularities that economists' theories should be trying to explain and, if economics is indeed a science rather than a branch of applied mathematics, then it is the task of economists to confront the theories with the evidence. As will become clear, there has been a great deal of empirical work on male labor supply and much of it has been imbedded explicitly in the standard neoclassical allocation theory. In fact, one of the most pleasing aspects of labor supply research during the last 20 years has been its careful attention to the theoretical underpinnings. At the same time, the overwhelming proportion of this empirical work has not questioned the validity of the conventional model; this model has been treated as a maintained hypothesis. Empirical research has concentrated on quantifying the magnitude of the presumed relationships. Such quantification is naturally an important ingredient of any science, but in many laboratory sciences refined attempts at calibration represent a stage of research that usually follows, not precedes, the testing of hypotheses. In male labor supply research, very little formal testing of the standard model has been undertaken. Labor supply research cannot be indicated for "measurement without theory", but it can be described as "measurement without testing". The theory is by no means empty of refutable implications and, at least when asked, most economists would grant that ultimately the implications or assumptions of any economic theory must correspond with actual behavior. So why has the great volume of empirical work involved so little testing of the standard model?

I suspect that one reason can be attributed to the fact that not merely are we reluctant to *reject* a theory until we have a viable substitute close at hand – this is a familiar proposition in the sociology of science – but also we hesitate even to *test* a theory until an alternative, behavioral, hypothesis is available.[6] The answer "I don't know" is something that an economist will say after being pushed by careful questioning, but he will not readily volunteer this response.

A more substantive reason for the lack of hypothesis testing in labor supply research is that many economists view such tests as tantamount to questioning whether a consumer's income-compensated demand curve for a commodity slopes downwards with respect to its price. After all, so the argument would go, the neoclassical theory of labor supply is a straightforward extension of the consumer's allocation problem and surely we believe that demand curves slope downwards? Putting aside the issue of whether that basic proposition of consumer theory has itself been corroborated, it is usually agreed that, in the absence of adverse evidence, the confirmation of a hypothesis increases with the number of favorable test outcomes: if the theory of consumer behavior had been found to be an apt description of the demand for apples, oranges, cherries, bananas, and many other fruit, an economist will wager it also applies to the demand for pears. But it is by no means clear that the exchanges taking place in the labor market are well described by analogies to the individual's behavior with respect to the purchases of fruit, that the evidence about the demand curves for fruit is relevant to the supply of work effort. As Coase (1937), Phelps Brown (1960, pp. 289–293), Simon (1951), and others have emphasized, labor market transactions possess many dimensions – the wages to be paid, the level of work effort to be applied, the range of activities to which the employee may be directed, the duration of the contract, and so on – and the particular combination of wages and hours worked represents only a subset of the bundle of items involved in the exchange. It is not at all obvious that this subset may be siphoned off from the rest and appropriately characterized by the sort of allocation process that the conventional model applies. I am not suggesting that the preferences of workers have nothing to do with their market work decisions, only that what I call below the canonical model may not be the most useful characterization of the way in which preferences and opportunities come together to determine outcomes in the labor market.

Nevertheless, the research attempts to measure the relevant parameters precisely have resulted in some notable advances in our understanding of the issues. For instance, the economics and econometrics of hours of work as distinct from labor force participation decisions are much better understood than they were 20 years ago. Though the literature on nonlinear budget constraints is by no means recent, it has been only in the past ten years that their implications for empirical work have been fully explored. The development and application of tractable dynamic models of labor supply have also represented a major advance in our

[6]Cf. Lakatos (1970, p. 179, n. 2): "The reluctance of economists and other social scientists to accept Popper's methodology may have been partly due to the destructive effect of naive falsificationism on budding research programmes."

understanding of the issues. We have much more and much better information today on the major empirical regularities in work behavior and especially on the importance of unobserved variables in accounting for variations across individuals in their hours of work. In all these respects, the standards of enquiry and critical debate in labor supply research have risen tremendously compared with the state of affairs 20 years or so ago. It is in this sense that undeniable progress has been made.

An outline of this survey is as follows. In the next section, Section 2, I identify the major time-series and cross-section empirical regularities in male labor supply behavior. It is these that any economic theory should be designed to address. Section 3 presents first the canonical static model of labor supply and then it immediately proceeds to deal with the problems in applying this model at the aggregative level. The static model is then amended to handle the situation of nonlinear budget constraints. Section 3 concludes with an outline of the most popular life-cycle model of labor supply. Section 4 addresses the issues in and results from the estimation of the static model. In this section, problems in specifying the model are first considered and then the results are presented from the U.S. nonexperimental literature, the British literature, and the U.S. experimental literature. Section 5 discusses the estimates from the applications of the life-cycle model. Some conclusions and suggestions for further research are given in Section 6.

2. Empirical regularities

2.1. Trends in work behavior

For a century or so, at least in North America and West Europe, a declining fraction of a man's lifetime has been spent at market work. This decline has been manifested in a number of ways: more years have been spent at school and the age of entry into full-time market employment has advanced; workers have been wholly or partially retiring from the labor force at younger ages; fewer hours have been worked per day and per week; and there have been more holidays and longer vacations. In addition, I suspect that work effort per hour has decreased, although this is difficult to verify. Consider now these different dimensions of work behavior.

Changes during the last 80 years or so in the labor force participation rates of men of different ages are documented for the United States, Britain, Canada, and Germany in Tables 1.1, 1.2, 1.3, and 1.4. The age group that maintained the closest association with the labor market has been men aged 25 to 44 years; for all four countries in all these years, more than 90 percent of these men were classified as members of the labor force. However, from the turn of the century

Table 1.1
United States: Male labor force participation rates (expressed as a percentage) by age over time.

Age (in years)	1890	1900	1910	1920	1930	1940	1950	1960	1970(a)	1970(b)	1982
10–13	17.8	17.7	9.2	6.0	3.3						
14/16–19	57.1	61.1	56.2	52.6	41.1	34.4	39.9	38.1	47.8	58.4	58.1
20–24	92.0	91.7	91.1	90.9	89.9	88.0	82.8	86.2	80.9	86.6	86.0
25–44	97.6	96.3	96.6	97.1	97.5	95.0	92.8	95.2	94.4	96.8	95.1
45–64	95.2	93.3	93.6	93.8	94.1	88.7	87.9	89.0	87.3	89.4	81.0
≥ 65	73.9	68.3	58.1	60.1	58.3	41.5	41.6	30.6	25.0	26.8	17.8
All	87.4	87.3	86.3	86.5	84.1	79.0	79.0	77.4	76.8	80.6	77.2

Notes: The Censuses after 1930 did not count children aged less than 14 years in the labor force. The age category "14/16–19" relates to 14–19 years for the years from 1890 to 1960 and to 16–19 years thereafter. The age category "All" describes all males aged 14 years and over from 1890 to 1960 and all males aged 16 years and over thereafter. The data for the years 1890, 1900, 1910, 1920, 1930, 1940, and 1950 are from Long (1958, Table A-2, p. 287). The data for 1960 are from U.S. Department of Commerce, Bureau of the Census, *U.S. Census of Population 1960: Employment Status and Work Experience*, Subject Reports PC(2)-6A, Table 1. The data for 1970(a) are from U.S. Department of Commerce, Bureau of the Census, 1970 *Census of Population*: *Employment Status and Work Experience*, Subject Reports PC(2)-6A, Table 1. The data for 1970(b) and for 1982 are from the monthly Current Population Survey of households and are not strictly comparable with the decennial census data in the other columns. The data for 1970(b) are from *Employment and Earnings*, January 1971, Table A-1, page 115 and those for 1982 from *Employment and Earnings*, January 1983, Table 3, page 142.

Table 1.2
Great Britain: Male labor force participation rates (expressed as a percentage) by age over time.

Age (in years)	1891	1911	1931	1951	1966	1981
< 20			84.7	83.8	70.6	64.6
20–24	98.1	97.3	97.2	94.9	92.6	89.2
25–44	97.9	98.5	98.3	98.3	98.2	97.5
45–64	93.7	94.1	94.3	95.2	95.1	90.2
65+	65.4	56.8	47.9	31.1	23.5	10.8
All			90.5	87.6	84.0	77.8

Notes: The category " < 20" relates to males aged 14–19 years in 1931, to males aged 15–19 years in 1951 and 1966, and to males aged 16–19 years in 1981. The category "All" relates to males aged 14 years and over in 1931, to males aged 15 years and over in 1951 and 1966, and to males aged 16 years and over in 1981. The data for the years 1891, 1911, 1931, 1951, and 1966 come from Department of Employment and Productivity, *British Labour Statistics Historical Abstract 1886–1968*, London, HMSO, 1971, Table 109, pp. 206–207. Those for 1981 are from Central Statistical Office, *Annual Abstract of Statistics 1983 Edition*, 1983, Table 6.16, p. 130.

Table 1.3
Canada: Male labor force participation rates (expressed as a percentage) by age over time.

Age (in years)	1911	1931	1951	1971	1980
14/15–19	64.6	51.4	48.1	46.6	51.9
20–24	92.2	92.3	91.8	86.5	79.7
25–44	97.1	97.6	96.3	92.7	92.2
45–64	94.4	94.8	90.6	85.9	83.3
65+	52.1	55.8	38.5	23.6	14.0

Notes: The youngest age category is 14–19 years in 1911, 1931, and 1951 and is 15–19 years in 1971 and 1980. For the years 1911, 1931, and 1951, the data are from Long (1958, Table A-11, p. 305). For the years 1971 and 1980, the sources are the International Labour Organization's *Yearbook of Labour Statistics* for 1975–76 and 1983, respectively.

for each and every age-group, the labor force participation rates of men in all these countries has fallen. The decline has been most marked for older men: for men aged 65 years and over, as recently as the early 1930s labor force participation rates of 58 percent, 48 percent, and 56 percent were recorded in the United States, Britain, and Canada, respectively. Twenty years later these rates had fallen by about the same 17 percentage points in each of these countries. A similar change was registered in Germany from 47 percent in 1925 to 27 percent in 1950. The post World War II period has witnessed further declines in each country in the labor force participation rates of older men. These declines have often been attributed to the expansion of government-organized social security

Table 1.4
Germany: Male labor force participation rates (expressed as a percentage) by age over time.

Age (in years)	1895	1907	1925	1939	1950	1970	1981
14/15–19	83.6	86.1	85.0	86.0	74.2	66.6	46.4
20–24	95.1	95.7	95.0	96.2	93.4	86.4	81.4
25–44	97.2	97.4	97.4	98.0	96.3	96.7	95.8
45–64	91.8	89.4	91.4	87.0	89.6	85.7	83.7
65+	58.8	50.2	47.4	29.7	26.7	16.0	7.0

Notes: Between 1895 and 1950, the youngest age group is 14–19 years; for 1970 and 1981, the youngest age group is 15–19 years. For the years 1895, 1907, 1925, and 1939, "Germany" consists of that area defined by her post World War I frontiers without the Saar. For the other years, "Germany" means the Federal Republic of Germany, excluding Berlin. The source for the data for 1895, 1907, 1925, 1939, and 1950 is Long (1958, Table A-16, p. 313). For the years 1970 and 1981, the sources are the International Labour Organization's *Yearbook of Labour Statistics* for 1973 and 1983, respectively.

Table 1.5
United States and Britain: labor force participation rates (expressed as percentages) of males and females combined over time.

United States		Britain	
Year	Participation	Year	Participation
1890	54.0	1891	61.3
1900	54.8	1901	59.9
1910	55.7	1911	60.4
1920	55.6	1921	58.6
1930	54.6	1931	57.7
1940	52.2	1951	57.7
1950	53.4	1961	59.3
1960	55.4	1971	61.4
1970	55.7	1981	61.0

Notes: The U.S. data in all years describe males and females aged 14 years and older and the British data in all years describe males and females aged 20 years and older. The U.S. data come from the Decennial Censuses and the precise sources are the same as those given beneath Table 1.1. The sources for the British data are the same as those given beneath Table 1.2.

systems and, indeed, it is unlikely that the taxes and benefits associated with the operation of these systems have not affected the labor force participation rate of older people.[7] On the other hand, it should be noted that the participation rates of older men were already declining before the period of the great expansion of government social security.

At the same time as the labor force participation rates of men were falling, those of women were rising. Indeed, as Table 1.5 shows for the United States and Britain, these changes largely offset one another. The absence of a trend in the overall (male and female) labor force participation rate prompted Klein and Kosobud (1961) to classify it as one of the "great ratios of economics". Both in 1910 and in 1970, the participation rate of all people aged 14 years and over in the United States was 55.7 percent; in 1981 in Britain, the participation rate of all people aged 20 years and over differed by only three-tenths of one percent from the rate in 1891.

Year-to-year movements in the labor force participation rate reflect the state of the business cycle as well as underlying trends. A convenient and simple way of

[7]Parsons (1980) claims the Social Security disability program is responsible for the declines during the post World War II period in the labor force participation rate of men aged 45–54 years. This interpretation is challenged by Haveman and Wolfe (1984) and then defended by Parsons (1984).

describing these cycles and trends is to fit the following equation to annual U.S. data from 1955 to 1982 for the civilian labor force participation rates of different groups of males in the population:

$$\Delta L_{jt} = \alpha_j + \beta_j \Delta U_t^{\mathrm{r}} + \varepsilon_{jt}. \tag{1}$$

In this equation, $\Delta L_{jt} = L_{jt} - L_{jt-1}$ and L_{jt} is the civilian labor force participation rate (expressed as a percentage) of group j in year t and $\Delta U_t^{\mathrm{r}} = U_t^{\mathrm{r}} - U_{t-1}^{\mathrm{r}}$ and U_t^{r} is the unemployment rate (expressed as a percentage) of white males aged 35–44 years in year t. The unemployment rate of this group is a better indicator of the stage of the business cycle as it operates in the labor market than is the overall unemployment rate and the superscript "r" on U designates this as the "reference" group. The responsiveness of the participation rate to the business cycle is measured by β while α reflects a linear time trend. The equation error is represented by ε_t and the index j runs over nine age groups and two racial groups.

The consequences of estimating eq. (1) by ordinary least squares are shown in Table 1.6. According to these estimates, over the past 27 years there has been a downward trend of almost three-tenths of one percent per year in the participation rate of white men and of almost one-half of one percent per year in the participation of black men. These trends are especially marked for young black men and for older men, both black and white. Although most of the estimates of β are negative (suggesting the participation rate falls in a recession),[8] these effects are small and not statistically significant except for younger men.[9] In general, very little variation in annual movements of male participation rates is removed by this cyclical indicator and Mincer's (1966) summary diagnosis – "some net cycle elasticity plus much residual variation due to other factors" – remains apt.[10]

For Britain, a time series on the male labor force participation rate for different age-groups is not published for the entire post-war period.[11] So I constructed an annual series for the entire adult male labor force participation

[8]The phenomenon of the labor force contracting in a recession is sometimes described as "the discouraged worker effect", that is, the costs of searching for acceptable employment rises in a recession to a degree such that it no longer pays some individuals to continue searching.

[9]A finding of long standing is that school enrollment rates of young people rise in a recession. See, for example, Duncan (1956).

[10]Equation (1) was also estimated with a different cyclical indicator, namely, the inventory-sales ratio in manufacturing and wholesale and retail trade. Very similar results were obtained with this variable as those reported in Table 1.6. Note that this is also the case for white men aged 35–44 years for whom there is a real danger of a spurious correlation between L and U^{r} in eq. (1).

[11]A series exists on an important subset of the male labor force (namely, all except employers, the self-employed, some part-time employees, and the military), but for men this was discontinued in January 1971. Analyses of these series are in Corry and Roberts (1970, 1974).

Table 1.6
United States: Estimates of trend (α) and cycle (β) in male civilian labor force participation rates by race and age, 1955–1982.

Age (in years)	α	β	R^2	D–W
White				
Total, ≥ 16	−0.284*(0.051)	−0.094(0.059)	0.09	1.59
16–17	0.181(0.246)	−1.103*(0.285)	0.37	1.61
18–19	0.078(0.229)	−0.800*(0.266)	0.26	1.29
20–24	0.015(0.158)	−0.201(0.184)	0.04	1.81
25–34	−0.057(0.038)	−0.121*(0.044)	0.22	1.78
35–44	−0.075*(0.027)	−0.042(0.031)	0.07	2.24
45–54	−0.169*(0.039)	0.056(0.046)	0.05	1.00
55–64	−0.651*(0.123)	0.008(0.143)	0.01	1.82
≥ 65	−0.796*(0.142)	−0.085(0.165)	0.01	1.49
Black and other				
Total, ≥ 16	−0.492*(0.116)	−0.162(0.134)	0.05	1.48
16–17	−0.626*(0.388)	−1.105*(0.449)	0.19	2.44
18–19	−0.780*(0.329)	−0.634(0.382)	0.10	2.24
20–24	−0.438(0.222)	−0.711(0.257)	0.23	1.59
25–34	−0.256*(0.115)	−0.125(0.133)	0.03	2.41
35–44	−0.220*(0.097)	−0.090(0.112)	0.02	2.18
45–54	−0.319(0.212)	−0.215(0.245)	0.03	2.66
55–64	−0.686*(0.324)	0.008(0.375)	0.01	2.02
≥ 65	−0.861*(0.273)	−0.147(0.316)	0.01	2.17

Notes: Estimated standard errors are in parentheses next to their associated regression coefficients. "D–W" is the Durbin–Watson statistic. For ease of reading, an asterisk has been placed next to those point estimates more than twice their estimated standard errors. The data are taken from the *Employment and Training Report of the President 1981* and from recent issues of *Employment and Earnings*.

rate over the 31 years from 1951 to 1981[12] and estimated the cyclical and trend movements in this labor force participation rate by fitting eq. (1) to the data. As a cyclical indicator, however, I used the deviations of the index of industrial production from a linear time trend, positive deviations corresponding to a low level of aggregate business activity and negative deviations to a high level of business activity. The labor force participation rate (expressed as a percentage) and this cyclical indicator were first-differenced and then, as in eq. (1), an

[12] To be precise, I constructed the ratio of the male labor force (called in Britain the working population) to the male home population aged 15 years and over. Both numerator and denominator are measured at the same moment, the middle of each year, and both relate to Britain (not the United Kingdom). The sources for the data were issues of the *Annual Abstract of Statistics* published by the Central Statistical Office. The mean value of this male labor force participation rate over the 1951–81 period is 0.836 with a standard deviation of 0.047.

ordinary least-squares equation was fitted to the data over the years 1952–1981.[13] The resulting estimates (with estimated standard errors in parentheses) are as follows:

$$\hat{\alpha} = \underset{(0.094)}{-\ 0.446^{*}}, \quad \hat{\beta} = \underset{(0.022)}{-\ 0.015}, \quad R^2 = 0.02, \quad \text{D–W} = 1.36.$$

According to these estimates, the male labor force participation rate in Britain over the last 30 years displays a small procyclical movement that would not be deemed significantly different from zero by conventional criteria and a negative trend of almost one-half of a percentage point per year. A comparison of these estimates with those in Table 1.6 for the entire U.S. male labor force indicates that movements in the British male labor force participation rate look very similar to those in the United States.

Hours worked by men declined markedly during the first four decades of the twentieth century. For the United States, this is evident from the data in Table 1.7 which are taken from the decennial Censuses of Population and which relate to men working in manufacturing industry only. They show that, whereas in 1909, 92 percent of all males were working more than 48 hours per week, the percentage had fallen to 54 percent in 1929 and then to 7 percent in 1940. This dramatic decline between 1929 and 1940 was in part the consequence of the Fair Labor Standards Act of 1938 which required that all hours over a standard workweek be compensated at the rate of 1.5 times the regular wage. Initially the standard workweek was set at 44 hours; since 1940 it has been 40 hours.[14]

The U.S. trends from 1940 onwards are indicated by the data in Table 1.8 which are not restricted to manufacturing industry. This table suggests that there has not been a pronounced change in hours worked per week since 1940 except for a reduction in the fraction working 41–48 hours and a greater bunching in the

[13] If I_t is the index of industrial production in year t (published in issues of the *Monthly Digest of Statistics*) and if T_t is a linear time trend, then I fitted to the annual data for the years 1948–81 the following ordinary least-squares equation:

$$I_t = \underset{(2.28)}{64.28} + \underset{(0.114)}{2.277}\ T_t,$$

where the figures in parentheses are estimated standard errors. (The mean value of I_t over these years is 104.1.) I then formed as a cyclical indicator, C_t, the difference between the predicted value of the index, $\hat{I}_t$, and the actual value of the index, I_t: $C_t = \hat{I}_t - I_t$. Thus, when C_t is positive, a recession is implied while when C_t is negative a high level of aggregate business activity is implied. (Defining it in this way, C_t moves in the same direction as the unemployment rate, the cyclical indicator used in describing variations in U.S. labor force participation rates.) Then, in accordance with the specification in eq. (1), annual changes in the male labor force participation rate were regressed on ΔC_t, where $\Delta C_t = C_t - C_{t-1}$. The results are not altered if the cyclical indicator is formed from regressing I_t on a quadratic time trend nor if a linear time trend is added to eq. (1).

[14] Some evidence assessing the effects of the FLSA on hours worked (especially in the 1940s) is contained in Lewis (1958). More information gauging the importance of the overtime provisions for hours worked is found in Ehrenberg and Schumann (1981).

Table 1.7
United States: Percentage distribution of weekly hours in manufacturing industry by employed males from the decennial censuses of population.

<table>
<tr><th>Hours worked</th><th>1909</th><th>1919</th><th>1929</th><th>1940</th><th>1950</th><th>1960</th><th>1970</th></tr>
<tr><td>≤ 34</td><td rowspan="6">7.9</td><td rowspan="4">12.2</td><td rowspan="2">0.5</td><td>13.3</td><td>6.9</td><td>7.8</td><td>10.1</td></tr>
<tr><td>35–39</td><td>4.9</td><td>3.5</td><td>4.7</td><td>4.9</td></tr>
<tr><td>40</td><td>2.8</td><td>51.3</td><td>64.3</td><td>56.4</td><td>53.1</td></tr>
<tr><td>41–43</td><td>1.0</td><td>14.2</td><td rowspan="3">17.1</td><td rowspan="3">19.3</td><td rowspan="3">18.6</td></tr>
<tr><td>44–47</td><td>3.8</td><td>14.8</td><td>1.9</td></tr>
<tr><td>48</td><td>32.6</td><td>26.9</td><td>7.4</td></tr>
<tr><td>49–53</td><td>7.3</td><td>16.4</td><td>25.1</td><td rowspan="3">3.9</td><td rowspan="3">5.6</td><td rowspan="3">7.8</td><td rowspan="3">8.8</td></tr>
<tr><td>54</td><td>15.4</td><td>9.1</td><td>6.3</td></tr>
<tr><td>55–59</td><td>30.2</td><td>13.7</td><td>15.1</td></tr>
<tr><td>60</td><td>30.5</td><td>9.1</td><td rowspan="2">7.5</td><td rowspan="2">3.0</td><td rowspan="2">2.6</td><td rowspan="2">4.0</td><td rowspan="2">4.4</td></tr>
<tr><td>> 60</td><td>8.7</td><td>3.0</td></tr>
</table>

Notes: The data relate to all employed males in 1960 and 1970 and to all employed wage and salary workers in the years earlier. The 1970 data describe males aged 16 years and over. In the years 1929–60, the data describe males aged 14 years and over. The Census collected data on "prevailing hours of labor" in 1909 and 1919 and on "customary hours of labor" in 1929. In the Census of 1940 and in subsequent years, the hours of work relate precisely to the census week. A small number of workers whose hours were not reported in 1929, 1940, and 1950 are not included in constructing the frequency distributions above. The 1970 data are from the *Industrial Characteristics* volume (Table 39) of the 1970 Census of Population. The 1960 data are from the *Industrial Characteristics* volume (Table 9) of the 1960 Census of Population. The 1950 data are from the *Industrial Characteristics* volume (Table 11) of the 1950 Census of Population. The 1940 data are from *Sixteenth Census of the United States 1940: Population Vol. III The Labor Force* Part I: U.S. Summary, Table 86, p. 259. The 1929 data are from *Fifteenth Decennial Census of the United States 1930: Manufactures 1929*, Vol. I, General Report, Table 5. The 1919 data are from the *Fourteenth Census of the United States Taken in the Year 1920*, Vol. VIII, Manufactures 1919, General Report and Analytic Tables, Table 17. The 1909 data are from the *Thirteenth Census of the United States Taken in the Year 1910*, Vol. VIII, Manufactures 1909, General Report and Analysis, Chapter XII, Table 8, p. 316.

distribution of hours worked at 40 hours. This spike at 40 hours per week is typically attributed to the overtime provisions of the Fair Labor Standards Act and the rising fraction of employees working these hours corresponds to the expansion of the Act's provisions: at the time of its implementation, less than one-fifth of all employees were covered by the overtime provisions; by the late 1970s, this figure had grown to approximately 58 percent.

The absence of a strong trend in hours worked during the post World War II period is consistent with the series on hours worked compiled from household interviews as part of the Current Population Survey (Table 1.9). These data are available on a consistent basis from 1955 and, as distinct from the data derived from the establishment surveys, they do not describe hours paid for, but hours worked by those at work. (Individuals on vacation, ill, or on strike are not covered by these hours of work data in Table 1.9.) The annual observations on hours worked per week by male wage and salary workers clearly reveal procycli-

Table 1.8
United States: Percentage distribution of hours worked of employed males during the Census week in 1940, 1950, 1960, and 1970.

Hours worked	1940	1950	1960	1970
1–14	1.59	2.02	4.42	4.54
15–29	5.79	4.75	4.60	5.71
30–34	4.47	3.37	3.09	5.03
35–39	4.56	2.86	4.48	4.91
40	33.53	41.45	41.59	43.06
41–48	29.37	19.29	19.59	17.41
49–59	8.87	10.69	10.36	9.99
≥ 60	11.83	15.57	11.87	9.35

Notes: These data describe all U.S. males aged 14 years and over who were employed during the Census week and who reported their hours of work. The 1940 data relate to wage and salary workers only. Also, in 1940, the categories labelled above as "1–14" and "15–29" are, in fact, less than 14 hours and 14–29 hours, respectively. The 1940 data are from *Sixteenth Census of the United States*: *1940*, *Vol. III*, *The Labor Force*, Part 1: U.S. Summary, Table 86, p. 259. The 1950 data are from *U.S. Census of Population 1950*, *Vol. IV*, *Special Reports*, Part I, Chapter A, Employment and Personal Characteristics, Table 13. The 1960 data are from *U.S. Census of Population 1960 Subject Reports*, Employment Status and Work Experience, Table 12. The 1970 data are from *U.S. Census of Population 1970 Subject Reports*, Employment Status and Work Experience, Table 17.

cal movements,[15] but after accounting for these cyclical effects there is little evidence of a trend over the past 27 years. These inferences come from fitting the following equation to the annual observations on weekly hours worked:

$$\Delta h_{jt} = \alpha_j + \beta_j \Delta U_t^{\mathrm{r}} + \varepsilon_{jt}, \tag{2}$$

where $\Delta h_{jt} = h_{jt} - h_{jt-1}$ and h_{jt} is the average weekly hours worked by group j in year t, $\Delta U_t^{\mathrm{r}} = U_t^{\mathrm{r}} - U_{t-1}^{\mathrm{r}}$ and U_t^{r} is the unemployment rate (expressed as a percentage) of white men aged 35–44 years in year t (the superscript "r" denoting my choice of these men as a reference group), and ε_{jt} is a stochastic error term. Any linear trend in hours worked is measured by α while β is supposed to reflect business cycle influences on hours. The index j runs over the six groups identified for the U.S. data in Table 1.9 and the ordinary least-squares estimates of the parameters α_j and β_j are given in Table 1.10. There are significant cyclical movements in hours worked for all workers except those in the older age groups. Most of the estimated trend terms (the α's) are negative, but none would be judged significant by conventional criteria except for that for

[15]For an analysis of weekly hours worked over the business cycle, see Bry (1959).

Table 1.9
United States, 1955–82, and United Kingdom, 1938–82:
Average weekly hours worked by male employees.

	United Kingdom: All adults	United States					
		All	14/16–17 years	18–24 years	25–44 years	45–64 years	≥ 65 years
1938	47.7						
1946–49	46.9						
1950–54	47.9						
1955–59	48.4	42.6	20.9	40.2	44.2	43.6	38.0
1960–64	47.5	42.5	18.4	39.9	44.5	43.7	35.7
1965–69	46.4	42.7	21.0	39.2	45.1	44.0	35.0
1970–74	45.2	41.8	22.5	38.1	44.1	43.3	32.5
1975–79	44.0	41.6	22.3	38.0	43.8	43.1	30.8
1980–82	43.0	40.8	20.6	37.1	43.0	42.2	30.6

Notes: The U.K. data relate to full-time manual workers and are taken from each October's earnings and hours enquiry of the major industries. The data are published in various issues of the *Ministry of Labour Gazette* and of the *Department of Employment Gazette*. The United States' data derive from household interviews in the Current Population Survey and they measure the average hours actually worked (not those paid for) of male employees in nonagricultural industries at work. (Consequently, those absent from work because of illness, vacation, or strike are not represented in these figures.) For the years 1955–58, the data are published in the *Current Population Reports*, Labor Force Series P-50, issues number 63 (Table 3), 72 (Table 18), 85 (Table 18), and 89 (Table 24). For the years 1959–64, the data are from *Special Labor Force Reports*, Table D-7 of each issue, Report numbers 4, 14, 23, 31, 43, and 52. For the years 1965–82, the data are taken from each January's issue of *Employment and Earnings* which give the figures for the preceding year. Before 1967, the youngest age group relates to those aged 14–17 years and from 1967 it relates to 16–17 years.

workers aged 65 years and over who reveal a declining trend of about 0.3 hours per year over the 1956–1982 period.

Although the downward trend in weekly hours worked in the United States seems to describe the data up to 1940 and not after that date, the length of the work year may have fallen because of increases in paid vacations and holidays. The only consistent time-series data relating to this dimension of work of which I am aware are the occasional surveys of employee compensation, a summary of which is presented in Table 1.11. Although the data in this table suggest that hours actually worked have fallen compared with hours paid for, the recorded changes are small.[16]

British long-term experience with weekly hours worked has been similar to that for the United States. The standard working week for manual workers set down in various collective bargaining agreements ranged from 48 to 60 hours or more

[16]There exist several studies investigating whether the absence of a trend during the post World War II period in weekly hours worked is spurious. Jones' (1974) study may be most thorough, but anyway the conclusions of Kniesner (1976b) and Owen (1979) are similar: hours worked have fallen little or not at all during this period and this influence survives adjustments for paid vacations and holidays.

Table 1.10
United States and Britain: Estimates of trend (α) and cycle (β) in weekly hours worked by male employees.

	α	β	R^2	D–W
Britain, 1949–1981				
All adult manual workers	−0.073(0.083)	−0.082*(0.020)	0.34	1.81
United States, 1956–1982				
All	−0.075(0.055)	−0.163*(0.062)	0.22	2.29
14/16–17 years	−0.088(0.197)	−0.731*(0.223)	0.30	1.29
18–24 years	−0.145(0.081)	−0.328*(0.091)	0.34	1.51
25–44 years	−0.044(0.066)	−0.194*(0.074)	0.21	2.19
45–64 years	0.003(0.321)	−0.525(0.363)	0.08	2.95
≥ 65 years	−0.329*(0.088)	0.103(0.100)	0.04	2.05

Notes: Estimated standard errors are given in parentheses next to their associated regression coefficients. "D–W" is the Durbin–Watson statistic. For ease of reading, an asterisk has been placed next to those point estimates more than twice their estimated standard errors. The data sources are given in the notes beneath Table 1.9.

before World War I. This fell further to 44 and 45 hours after World War II. A comprehensive survey of hours actually worked by British manual workers was conducted in October 1938 by the Ministry of Labour. In the principal industries, it found that the average hours worked by adult male manual workers were 47.7 while the frequency distribution of hours worked was as follows: 15.5 percent of these employees worked less than 44 hours, 16.4 percent worked from 44 hours to less than 47 hours, 27.6 percent worked between 47 and 48 hours (inclusive), and 39.2 percent worked more than 48 hours.

The movement since 1938 in weekly hours worked by male manual workers is given in the first column of Table 1.9. Again, to determine whether or not a trend exists in these post World War II data, eq. (2) was fitted to the annual observations on hours worked from 1949 to 1981. As was the case when eq. (1) was fitted to the British male labor force participation rate, eq. (2) was estimated using as a cyclical indicator the deviation of the index of industrial production from its fitted linear trend. The ordinary least-squares estimates of eq. (2) fitted to the British data are given in the first line of Table 1.10 and they are similar to the U.S. results: there is a strong procyclical variation in hours worked in Britain and no significant time trend. The strong cyclical influence on hours worked probably accounts for much of the difference in the frequency distribution of hours between September 1968 and April 1981 as shown in Table 1.12. That is to say, the fraction of male employees working between 35 and 39 hours increased from 18.5 percent in September 1968 to 22.0 percent in April 1977 and to 28.3 percent in April 1981 while the percentage working in each of the categories above 42 hours decreased uniformly from 1968 to 1977 to 1981. However, these

Table 1.11
United States: Paid leave hours as a percentage of total hours paid for, 1958, 1966, 1977.

	1958	1966	1977
Manufacturing:			
Nonoffice workers	6	6	8.4
Office workers		8	10.5
All workers		7	9.0
Nonmanufacturing:			
Nonoffice workers		4	5.5
Office workers		7	8.9
All workers		5	6.9
All nonfarm industries:			
Nonoffice workers		5	6.6
Office workers		7	9.2
All workers		6	7.6

Notes: The 1958 data are from U.S. Department of Labor, *Composition of Payroll Hours in Manufacturing, 1958*, Bureau of Labor Statistics Bulletin number 1283, October 1960. The 1966 data are from U.S. Department of Labor, *Employee Compensation in the Private Nonfarm Economy, 1966*, Bureau of Labor Statistics Bulletin number 1627, June 1969. The 1977 data are from U.S. Department of Labor, *Employee Compensation in the Private Nonfarm Economy, 1977*, Bureau of Labor Statistics, Summary 80-5, April 1980.

years exhibited a growing slack in the level of aggregate business activity as indicated, for instance, by the male unemployment rate (seasonally unadjusted and including school leavers) which stood at 3.2 percent in September 1968, at 7.0 percent in April 1977, and at 12.6 percent in April 1981.

What is not reflected in these data on hours worked in a given week is the increasing length of paid vacations in Britain over the post-war period. I know of no data that document the number of days paid for, but not worked in Britain. However, the information in Table 1.13 suggests that there has been a substantial increase in paid vacations. These data are taken from national collective bargaining agreements and they concern the length of paid vacations to which covered workers are entitled. Whereas, in fact, annual paid vacations were unusual for manual workers in Britain before World War II, the data in Table 1.13 indicate that there have been substantial increases in the length of paid vacations during the last 30 years. The increases in paid vacations were especially pronounced during periods of government-mandated wage controls and incomes policies that diverted attention to less visible ways (than cash) of increasing employee compensation.[17]

The discussion above has documented the trends this century in male labor

[17]See Department of Employment, *Employment Gazette*, Vol. 89, No. 4, April 1981, p. 184.

Table 1.12
Britain: Percentage distribution of weekly hours worked by male employees in 1968, 1977, and 1981.

	September 1968	April 1977	April 1981
$0 < h \leq 24$	2.0	1.8	1.6
$24 < h \leq 30$	2.1	2.0	2.1
$30 < h \leq 35$	4.2	5.5	6.8
$35 < h \leq 37$	7.3	11.2	12.4
$37 < h \leq 39$	11.2	10.8	15.9
$39 < h \leq 40$	20.1	26.2	27.6
$40 < h \leq 42$	7.1	5.3	5.4
$42 < h \leq 44$	8.0	7.3	6.1
$44 < h \leq 46$	6.6	6.0	5.0
$46 < h \leq 48$	7.0	5.9	4.3
$48 < h \leq 50$	5.2	4.3	3.1
$50 < h \leq 54$	7.0	5.3	3.7
$54 < h \leq 60$	7.0	4.9	3.4
$60 < h \leq 70$	4.0	2.5	1.8
$70 < h$	2.0	1.0	0.8

Notes: These data cover all men (both manual and nonmanual workers) whose pay for the survey period was not affected by absence. The 1968 data are from Department of Employment and Productivity, *New Earnings Survey 1968*, H.M.S.O., 1970, Table 83, p. 120. The 1977 data are from Department of Employment, *New Earnings Survey 1977, Part A: Report and Key Results*, H.M.S.O., 1977, Table 27, p. A35. The 1981 data are from Department of Employment, *New Earnings Survey 1981, Part A: Report and Key Results*, H.M.S.O., 1981, Table 27, p. A90.

force participation rates and hours worked. Just as men have spent a declining fraction of their lives at work for pay, so have they spent an increasing fraction at school. Some evidence of this is provided by the cohort analyses in Tables 1.14 and 1.15. These data are taken from surveys in 1970 and in 1971 of men of different ages and they document the striking association between the age of the cohort and the years spent at school.[18]

2.2. *Cross-sectional variations in work behavior*

Some important variations in labor force participation across individual men are documented by the linear probability estimates in Table 1.16. These are reproduced from Bowen and Finegan's (1969) monumental work on the 1960 Census of Population. As is evident from Table 1.16, there is a strong positive relation-

[18] They are a biased indicator of the degree to which schooling levels completed have risen over time insofar as mortality rates are associated with years of schooling. On this association, see Grossman (1975).

Table 1.13
United Kingdom: Manual workers' basic paid vacation entitlements as set down in national collective bargaining agreements, 1951–1982.

Year	Percentage of workers with basic vacations of < 2 weeks	2 weeks	Between 2 and 3 weeks	3 weeks	Between 3 and 4 weeks	≥ 4 weeks
1951	31	66	2	1		
1955	1	96	2	1		
1960		97	1	2		
1965		75	22	3		
1970		41	7	49	3	
1975		1	1	17	51	30
1980				2	24	74
1982				2	5	93

Notes: Until 1965, the column given as "3 weeks" is, in fact, "3 weeks and over". In addition to these annual vacations, workers are usually entitled to payment of wages for public or statutory holidays or days in lieu of these payments. The data for 1951–65 are from the Department of Employment and Productivity, *British Labour Statistics Historical Abstract 1886–1968*, London, H.M.S.O., 1971, Table 34, p. 91. Data for 1970 onwards are from various issues of the Department of Employment's *Gazette*.

Table 1.14
United States: Schooling completed by the male population in 1970 by age.

Years of age in 1970	Year of birth	Median years school completed	Percentage of cohort whose highest schooling levels completed were: ≥ 4 years of college	≥ 2 years of college	≥ 4 years of high school	≥ 8 years of elementary school	≥ 5 years of elementary school
≥ 75	≤ 1895	8.3	5.3	8.8	20.9	57.1	79.4
70–74	1896–1900	8.6	6.2	10.1	24.5	64.1	85.3
65–69	1901–1905	8.8	7.4	11.8	27.6	68.1	88.0
60–64	1906–1910	9.6	8.7	13.9	34.7	75.1	91.5
55–59	1911–1915	10.7	9.3	14.9	41.4	79.8	93.4
50–54	1916–1920	12.0	10.8	17.2	49.7	84.7	95.0
45–49	1921–1925	12.2	14.1	21.2	55.6	87.1	95.7
40–44	1926–1930	12.2	16.4	23.7	57.3	88.4	96.4
35–39	1931–1935	12.4	18.6	26.2	64.3	90.2	96.8
30–34	1936–1940	12.5	18.5	26.6	68.9	92.7	97.6
25–29	1941–1945	12.6	19.5	29.6	74.2	94.7	98.2

Notes: These data are constructed from those given in Table 199 of U.S. Department of Commerce, Bureau of the Census, *1970 Census of Population*, Volume I, Characteristics of the Population, Part 1, U.S. Summary, Section 2, June 1973.

ship between participation and schooling: for prime-age males (that is, those aged 25–54 years), a person with 17 or more years of schooling has almost a 9 percent higher probability of being in the labor force than someone with 0–4 years of schooling who is otherwise identical in his observable characteristics. This participation–schooling relationship among older men is especially strong. For prime-age males, ceteris paribus, a white man is almost 2 percent more likely

Table 1.15
Britain: Highest educational qualification attained by male population in 1971 by age.

Years of age in 1971	Year of birth	Percentage of cohort whose highest educational qualifications were at the level of "Higher education"	"Middle education"	"Lower education"
≥ 65	≤ 1906	5.1	14.9	80.0
60–64	1907–1911	7.6	23.6	68.8
50–59	1912–1921	6.1	25.8	68.1
40–49	1922–1931	10.7	27.2	62.1
30–39	1932–1941	14.2	33.2	52.6
25–29	1942–1946	13.6	41.9	44.5

Notes: The level "Higher education" includes university degrees, equivalent professional qualifications, and other qualifications beyond the GCE "A" level standard. "Middle education" includes any subjects passed at the GCE "A" level and "0" level plus clerical and commercial qualifications and apprenticeships. "Lower education" means no qualifications attained. The data are from Office of Population Censuses and Surveys, Social Survey Division, *The General Household Survey 1971*, Introductory Report, H.M.S.O., Table 7.15.

to be in the labor force than a black man. A married man with his spouse present is much more likely to be in the labor force (8 percent more likely for prime-age males, other things equal) than a man with a different marital status. Greater nonwage income is associated with lower participation and participation probabilities form an inverted U-shape with respect to age: they rise until 25 years of age, then remain constant until the middle-to-late fifties at which point they decline rapidly.

Some empirical regularities with respect to the hours worked by men are evident from the ordinary least-squares regression results presented in Table 1.17. These estimates describe the work behavior of 23059 men aged from 25 to 55 years of age at the time of the 1980 Census of Population.[19] The column "weekly hours" relates to the number of hours usually worked during those weeks the person worked in 1979; the column "weeks per year" relates to the number of weeks during 1979 in which a person did any work for pay or profit (including paid vacation and paid sick leave); and the column "annual hours" relates to the product for any person of "weekly hours" in 1979 and "weeks per year" in 1979.

[19] The sample of 23059 men was determined as follows. There are 94025 dwelling units included in the Public Use Sample Tape "C" sample nationwide file. Of these, 8,021 units were rejected because they were vacant, another 25725 units were rejected because no male was listed as household head (or, if a woman was listed as the household head, no husband or live-in partner was listed), another 22198 units were rejected because the male was not aged between 25 and 55 years (inclusive), another 1097 were rejected because the male received some farm income, and another 12933 were rejected because either the male's labor income was truncated (being less than \$ − 9,995 or more than \$75000) or the male's data on labor supply were missing. This yields a sample of 24051 men of whom 992 had zero hours of work in 1979.

Table 1.16
Ordinary least-squares estimates of labor force participation equations fitted to data on individual men from the 1/1000 sample of the 1960 U.S. Census of Population.

	(1)	(2)	(3)	(4)
Age-group	18–24 years	25–54 years	55–64 years	65–74 years
nobs	3095.0	22 415.0	4967.0	3392.0
modv	94.0	96.7	85.2	38.7
Estimates of:				
Intercept	79.3	83.7	73.5	48.4
Years of schooling				
0–4	Reference	Reference	Reference	Reference
5–7	5.3(3.3)	4.1(0.7)	5.4(1.8)	5.9(2.3)
8	9.4(3.2)	5.2(0.7)	10.5(1.7)	14.5(2.3)
9–11	9.2(2.9)	6.3(0.6)	13.4(1.8)	17.9(2.7)
12	10.3(2.9)	6.9(0.6)	13.5(2.0)	20.4(2.9)
13–15	6.5(3.1)	8.1(0.7)	17.2(2.2)	25.1(3.5)
16	11.2(3.5)	8.5(0.7)	18.0(2.7)	31.9(4.5)
≥17	5.6(5.3)	8.8(0.7)	26.6(3.1)	39.7(5.5)
Ethnicity				
Black	Reference	Reference	} Reference	} Reference
Other nonwhite	1.2(4.9)	2.7(1.3)		
White	1.5(1.3)	1.8(0.4)	1.9(1.7)	0.5(2.8)
Marital status				
Never married			Reference	Reference
Separated or divorced	} −6.7(0.9)	} Reference	4.0(2.4)	1.7(4.0)
Widowed			0.7(2.7)	1.2(3.5)
Married spouse present	Reference	7.8(0.3)	12.6(1.7)	12.7(2.9)
Nonwage income				
< $500		Reference	Reference	Reference
$500–999		−4.1(0.5)	−19.0(1.6)	−31.1(2.3)
$1000–1999		−10.1(0.7)	−35.1(1.8)	−39.9(1.9)
$2000–2999		−13.9(1.2)	−34.0(2.6)	−44.8(2.5)
$3000–4999		−7.0(1.2)	−36.7(3.0)	−55.2(3.3)
≥ $5000		−13.2(1.4)	−30.3(3.1)	−40.9(4.3)
Years of age				
18/55/65	Reference		Reference	Reference
19/56/66	7.0(1.9)		−1.6(1.9)	0.5(2.9)
20/57/67	7.5(1.8)		−1.5(2.0)	−1.6(2.9)
21/58/68	10.5(1.8)		−2.0(2.0)	−2.4(3.0)
22/59/69	8.2(1.8)		−1.3(1.9)	−4.7(3.1)
23/60/70	11.3(1.8)		−5.5(2.0)	−6.4(3.1)
24/61/71	8.6(1.8)		−6.0(2.1)	−12.2(3.1)
62/72			−5.6(2.1)	−9.2(3.3)
63/73			−10.8(2.1)	−8.5(3.5)
64/74			−9.1(2.1)	−13.8(3.5)
25–34		Reference		
35–44		−0.4(0.3)		
45–54		−1.2(0.3)		
$\hat{F}$ ratio	10.5	92.2	45.4	40.0

Notes: These estimates are from Bowen and Finegan (1969, Tables A-38, A-1, A-14, and A-15). Standard errors are given in parentheses next to estimated coefficients. The number of observations is given by "nobs" and the mean of the dependent variable is given by "modv". All the variables above are in the form of dummy variables with "Reference" indicating the category omitted from the list of variables. Under the group of variables "Years of age" the first column (18,19,20,etc.) relates to the 18–24 year olds in column (1), the second column (55,56,57,etc.) relates to the 55–64 year olds in column (3), and the third column (65,66,67,etc.) relates to the 65–74 year olds in column (4). The group described as "Separated or divorced" under "Marital status" includes married men with their spouses absent. "Nonwage income" represents the sum of rental income, interest, dividends, alimony, pensions, and welfare payments.

Table 1.17
Ordinary least-squares estimates of male hours and weeks worked equations fitted to data from 1/1000 sample of the 1980 U.S. Census of Population.

Independent variable		Dependent variable		
Mean and standard deviation	Definition	Weekly hours	Weeks per year	Annual hours
	Constant	36.2	34.91	1194.88
9.53	Average hourly earnings	−0.226	−0.107	−13.78
(10.00)	in dollars	(0.006)	(0.005)	(0.36)
0.477	Interest, dividend, and	0.089	0.010	4.62
(2.318)	rental income in thousands of dollars	(0.026)	(0.022)	(1.57)
0.307	Other income of the indi-	−0.214	−1.141	−55.38
(1.502)	vidual in thousands of dollars	(0.039)	(0.034)	(2.40)
5.978	Family income minus male	−0.027	0.001	−1.17
(7.547)	head's in thousands of dollars	(0.008)	(0.007)	(0.52)
37.98	Age in years	0.385	0.471	38.33
(8.89)		(0.072)	(0.062)	(4.40)
1521.3	Age squared in years	−0.005	−0.005	−0.43
(704.0)		(0.001)	(0.001)	(0.06)
0.46	1 = Completed high school	1.098	2.200	132.61
(0.50)		(0.229)	(0.198)	(14.06)
0.46	1 = Completed any college	2.152	3.020	219.20
(0.50)	education	(0.237)	(0.205)	(14.57)
0.43	1 = Any children aged	0.199	0.237	20.70
(0.74)	0–6 years	(0.090)	(0.078)	(5.55)
0.82	1 = Any children aged	0.133	0.044	7.55
(1.06)	7–16 years	(0.062)	(0.054)	(3.83)
0.84	1 = Married and spouse	1.068	1.803	121.47
(0.36)	present	(0.197)	(0.170)	(12.09)
0.02	1 = Married and spouse	1.044	0.112	59.41
(0.15)	absent	(0.424)	(0.366)	(26.04)
0.05	1 = Hispanic	−1.981	−1.711	−160.51
(0.21)		(0.284)	(0.245)	(17.44)
0.07	1 = Black	−2.736	−1.549	−190.34
(0.26)		(0.229)	(0.198)	(14.05)
0.02	1 = Not White nor Black	−1.508	−1.489	−130.32
(0.15)	nor Hispanic	(0.390)	(0.337)	(23.94)
0.06	1 = Self-employed	4.473	−0.260	219.20
(0.24)		(0.246)	(0.213)	(15.12)
0.19	1 = Employed by local, state,	−1.133	0.274	−43.94
(0.39)	or Federal government	(0.152)	(0.131)	(9.30)
0.05	1 = Health disability	−1.342	−5.312	−262.86
(0.22)		(0.262)	(0.226)	(16.05)
0.83	1 = Lived in a metropolitan	−0.518	0.679	2.68
(0.38)	area	(0.160)	(0.138)	(9.79)
0.06	1 = Lived in New England	0.400	0.095	26.98
(0.23)		(0.289)	(0.250)	(17.76)

Table 1.17 continued

Independent variable		Dependent variable		
Mean and standard deviation	Definition	Weekly hours	Weeks per year	Annual hours
0.16	1 = Lived in Mid-Atlantic	−0.434	0.608	4.81
(0.37)	states	(0.213)	(0.183)	(13.04)
0.19	1 = Lived in East North	0.731	0.639	64.99
(0.39)	Central states	(0.205)	(0.177)	(12.60)
0.07	1 = Lived in West North	0.546	0.599	53.11
(0.26)	Central states	(0.267)	(0.231)	(16.40)
0.16	1 = Lived in South Atlantic	0.475	0.897	62.03
(0.37)	states	(0.214)	(0.185)	(13.14)
0.06	1 = Lived in East South	0.065	0.253	20.39
(0.23)	Central states	(0.290)	(0.251)	(17.85)
0.10	1 = Lived in West South	1.492	0.879	112.58
(0.30)	Central states	(0.241)	(0.208)	(14.82)
0.05	1 = Lived in Mountain states	0.143	0.251	21.34
(0.23)		(0.294)	(0.254)	(18.05)
	$\bar{R}^2$	0.096	0.117	0.130

Notes: The mean (and standard deviation in parentheses) of weekly hours is 43.41 (9.26), that of weeks worked is 48.89 (8.08), and that of annual hours is 2131.07 (579.26). There are 23059 observations in each regression equation. Another 992 observations had zero annual hours of work so the labor force participation rate of this group was 95.9 percent. The East North Central states are Ohio, Indiana, Illinois, Michigan, and Wisconsin. The West North Central states are Iowa, Minnesota, Missouri, Kansas, Nebraska, South Dakota, and North Dakota. The East South Central states are Kentucky, Tennessee, Alabama, and Mississippi. The West South Central states are Arkansas, Louisiana, Oklahoma, and Texas. The omitted region consists of California, Oregon, Washington, Alaska, and Hawaii.

The notes to Table 1.17 provide mean values and standard deviations of these variables. According to these estimates, a dollar higher average hourly earnings is associated with 14 fewer hours worked per year, the responsiveness of weekly hours being greater than the responsiveness of weeks per year. The behavioral implications of this negative hours–earnings association are not clear, however: the interviewees are asked their earnings (wage income plus self-employment income) in 1979 and the variable "average hourly earnings" consists of annual earnings divided by annual hours of work; consequently, any errors in measuring hours of work are communicated to the measure of average hourly earnings. Increases in interest, dividend, and rental income are positively (though weakly) associated with hours of work while other income of the individual (mainly public assistance and social security and, as such, it is typically work-related income) is negatively associated with work behavior. The hours–age relationship forms an inverted U-shape with the maximum occurring around 44 years of age. Men with higher schooling levels completed work longer hours as do fathers with

Table 1.18
Percentage distribution of hours worked in 1974 according to hours worked in 1967.

		Hours in 1967							
		0–1499	1500–1849	1850–2149	2150–2499	2500–2999	3000–3499	≥ 3500	
Percent of observations in 1967		5.5	7.0	29.0	23.7	19.3	9.4	6.0	100
Hours in 1974	0–1499	35.7	16.7	10.1	8.7	5.0	8.1	7.4	
	1500–1849	12.4	26.1	10.5	10.7	8.5	2.5	7.6	
	1850–2149	14.7	31.7	49.5	28.2	20.7	17.3	11.9	
	2150–2499	14.1	15.1	18.1	28.6	27.4	15.3	9.7	
	2500–2999	13.5	4.8	7.8	16.1	23.3	30.2	17.1	
	3000–3499	5.9	3.0	2.4	5.8	9.6	18.0	23.1	
	≥ 3500	3.7	2.6	1.6	1.8	5.6	8.6	23.2	
Total		100	100	100	100	100	100	100	

Notes: The underlying data consist of 2209 men all of whom were married in the first year of interview (1968) and all of whom worked no less than 250 hours in both 1967 and 1974.

younger children, married men, non-Hispanic white men, self-employed men, men who claimed a health disability, and men who were not government workers.[20] Some marked regional variations in hours worked are evident. It is important to observe that only between 9.7 percent and 13.0 percent of the sample variation in these measures of work behavior is accounted for by the least-squares combination of variables in Table 1.17. Indeed, the inability of empirical studies of working hours to remove anything more than a relatively small fraction of the observed variation in a large sample's hours is striking.

Notwithstanding the popular notion that, each and every year, virtually all men work 2000 hours per year (40 hours per week and 50 weeks per year), in fact there exists a substantial amount of variation across individuals in their hours of work and also important variations for many individuals from year to year. Some indication of the temporal variations in annual hours of work is provided by the data in Table 1.18 which are taken from a paper by Hill and Hoffman (1977) that also analyzes men from the Michigan Panel. The data in Table 1.18 describe 2209 men all of whom were married in the first year of interview (1968) and all of whom were at work for at least 250 hours in both the years 1967 and 1974. The first column of Table 1.18 shows that 5.5 percent of these men worked 0–1499 hours in 1967; of these men who worked 0–1499 hours in 1967, 35.7 percent also

[20] In the dummy variable categories, the omitted groups are men who did not complete high school, men without any children, unmarried men, non-Hispanic white men, men neither self-employed nor working for the government, men with no health disability, men not living in a metropolitan area, and men living in California, Oregon, Washington, Alaska, and Hawaii.

worked 0–1499 hours in 1974. The main diagonal in Table 1.18 tends to have larger entries than the off-diagonal terms, but this is by no means always the case: thus, of those who worked 2500–2999 hours in 1967, 23.3 percent worked the same hours in 1974, whereas 27.4 percent worked 2150–2499 hours, an indication of some regression towards the mean. The authors describe these changes as "pervasive" and, indeed, 51.1 percent of the variance of the logarithmic change in these men's annual earnings between 1967 and 1974 was attributable alone to the variance of the logarithmic change in hours worked.

3. Conceptual framework

3.1. *The canonical model*

The model that guides most economists' analyses of the determinants of the supply of working hours derives most directly from Hicks' (1946) paragraph 11 of his Mathematical Appendix. According to this characterization, the labor supply function is derived from a general model of consumer demand in which a fixed endowment of a commodity is divided into one part for sale on the market and another part reserved for direct consumption. In this instance, the endowment consists of a fixed block of time, T, that in the simplest of cases is to be divided between hours worked in the market, h, and hours spent in other activities, $l: T = h + l$. The *reservation demand* for hours of "leisure", l, simply consists of what is left over from market sales of h. In this canonical model, there is no savings decision to be made and the individual is fully informed of all the values of the relevant variables and parameters. An individual with personal characteristics A (such as his age or race) possesses a well-behaved (real-valued, continuous, quasi-concave) utility function defined over his consumption of commodities, x, and his hours of work, h:

$$U = U(x, h; A, \varepsilon), \tag{3}$$

where ε stands for the individual's "tastes". Whether ε is called a taste component or an individual's "ability in home production" or whatever, the essential point is that, unlike the variables in A, ε is unobserved to the researcher. In accordance with the empirical findings reported above whereby a substantial fraction of the variation in hours of work across individuals is not removed by variables observed by the economist, the presence of ε in the utility function allows for individuals to differ from one another in ways not observed by the researcher.

The partial derivative of U in eq. (3) with respect to x is assumed to be positive and that with respect to h is assumed to be negative, at least in the neighborhood of the observed hours of work.[21] If throughout the analysis the relative prices of the different commodities do not change, then x represents a Hicksian composite commodity. The individual sells his services to the consumer in the product market either directly when he is "self-employed" or indirectly when he is employed by a firm to contribute towards producing a commodity. In either case, the individual's total compensation, c, for his market work depends positively upon how much of his time is alloted to this activity: $c = c(h)$. In the simplest of cases, each hour of work is rewarded at the same fixed rate, w, and $c(h) = wh$. The average and marginal payment for his work time are now the same and, if p denotes the fixed per unit price of the bundle of commodities x and if y represents income independent of the working decision, then the individual's budget constraint is linear and homogeneous of degree zero in p, w, and y:

$$px = wh + y. \tag{4}$$

The individual is assumed to do the best he can given the constraints he faces. Or, more formally, the individual chooses values of $x > 0$ and $h \geq 0$ that maximize eq. (3) subject to the budget constraint (4).[22] Observe that this problem has been characterized in terms of a single individual's objective function and budget constraint. This is by no means necessary. Suppose this individual's utility depends upon his spouse's market work time (h_2) in addition to his own work time (h_1):

$$U = U(x_1, h_1, h_2;\ A, \varepsilon). \tag{5}$$

If his spouse's utility function contains the same arguments and if the two of them pool their incomes and expenditures,

$$p_1x_1 + p_2x_2 = w_1h_1 + w_2h_2 + y, \tag{6}$$

in the simple case of a linear budget constraint where w_1 and w_2 denote the

[21]Another characterization of the problem involves defining the utility function over activities that are produced by a household production function whose inputs are purchased goods and time. In Becker's (1965) formulation, time at market work does not directly enter the utility function at all and so the question does not arise of whether U is decreasing in h. See Atkinson and Stern (1979) and Chapter 4 by Gronan in this Handbook.

[22]The problem is sometimes written in terms of leisure, l, and the endowment of time, T, by having the individual select $x > 0$ and $l > 0 \leq T$ to maximize $U(x, l;\ A, \varepsilon)$ subject to $px + wl = wT + y = I$, where I is called full income. This formulation in an empirical context poses the problem of what value to assign to T, the results not being invariant to this assignment. I prefer the formulation of the problem in the text that involves variables whose counterparts in the data are more easily defined.

hourly wage rates paid to individuals 1 and 2, respectively, and x_1 and x_2 represent the consumption of commodities by individuals 1 and 2, respectively, then the problem becomes one of selecting x_1, x_2, h_1, and h_2 to maximize the utility functions of the two individuals subject to their joint budget constraint (6). As stated, this is a bargaining problem and typically the solution may be satisfied with many different combinations of x_1, x_2, h_1, and h_2. To determine which of the many possibilities will obtain requires the introduction of particular behavioral postulates that yield specific solutions.[23] The usual method of handling these problems is to assume that the social choice conditions for the existence of a well-behaved aggregate (household) utility function have been met or that the household's utility function is identical with that of the "head" of the household who integrates the welfare of all the household's members [see Samuelson (1956) and Becker (1974)]. Under these circumstances, x_1, x_2, h_1, and h_2 are chosen to maximize eq. (5) subject to the budget constraint (6). Clearly, in these household models, each individual's allocation of his work time depends upon not only his own wage rate, but also the wage rate of his spouse.

Return to the formulation whereby a single individual selects $x > 0$ and $h \geq 0$ to maximize $U(x, h;\ A, \varepsilon)$ subject to a linear budget constraint $px = wh + y$. It is important to distinguish the characteristics of the interior solution for hours of work, $h > 0$, from the corner solution, $h = 0$. In the case of the individual selecting a positive number of hours to supply to the market, the first-order condition for a constrained maximum[24] requires that commodities and hours of work be chosen such that the negative of the marginal rate of substitution (m) of working hours for commodities equals the real wage (w/p):

$$\frac{w}{p} = -m(x, h;\ A, \varepsilon) = -\frac{\partial U/\partial h}{\partial U/\partial x}. \tag{7}$$

The reduced form equations, the commodity demand and working hours supply functions, are derived by solving eq. (7) jointly with the budget constraint (4):

$$\left.\begin{aligned} x &= x(p, w, y;\ A, \varepsilon) \\ h &= h(p, w, y;\ A, \varepsilon) \end{aligned}\right\}, \quad \text{if } h > 0. \tag{8}$$

The properties of this hours of work equation are discussed below. This interior solution for hours of work may be expressed differently by making use of the concept of the individual's reservation wage, w^*. The real reservation wage, w^*/p, is the slope of an indifference curve between commodity consumption and

[23] For instance, Manser and Brown (1979) assume a Nash solution to this bargaining problem.
[24] The assumption that the utility function is quasi-concave ensures the satisfaction of the second-order conditions for a constrained maximum.

hours at work evaluated at $h = 0$ and, for any given individual, typically the value of this reservation wage will vary from one indifference curve to another, i.e. the reservation wage will depend upon x and so indirectly upon y for any given A and $\varepsilon: w^*(y, A, \varepsilon)$. Equivalently, the real reservation wage is equal to the negative of the marginal rate of substitution of working hours for commodities evaluated at $h = 0$: $w^*/p = -m(x, 0;\ A, \varepsilon)$. The reservation wage is the individual's implicit value of his time when at the margin between participating in the labor market and not participating.[25] If, at that margin, the market's valuation of his time, w, exceeds the individual's implicit value of his time, w^*, then he will participate in the labor market and supply a positive number of hours of market work. Then eqs. (7) and (8) will hold enabling us to write:

$$\text{if } w > w^*, \quad \text{then } h = h(p, w, y;\ A, \varepsilon) > 0. \tag{9}$$

On the other hand, if at the margin between participating and not participating in the labor market the individual places a greater value on an extra unit of his time than does the market (that is, if $w^* > w$), then naturally the individual will reserve his entire allocation of time for himself and the solution to the constrained maximization problem will be a corner $h = 0$. Consequently, we may write:

$$\text{if } w \le w^*, \quad \text{then } h = 0. \tag{10}$$

Consider now the properties of the labor supply function $h = h(p, w, y;\ A, \varepsilon)$ derived in eq. (8). The zero homogeneity property that was introduced through the budget constraint carries over to the commodity demand and labor supply functions: a given proportionate change in p, w, and y leaves the optimizing values of x and h in eqs. (8) unchanged. A second property of the labor supply function so derived is manifested when examining the effect of a small increase in w on the supply of h: $\partial h/\partial w$. The Slutsky equation decomposes this effect into a substitution effect, s, and an income effect, $h \cdot \partial h/\partial y$:

$$\frac{\partial h}{\partial w} = s + h\frac{\partial h}{\partial y}. \tag{11}$$

The substitution effect, s, measures the utility-constant (or income-compensated) effect of an increase in the wage rate on the individual's hours of work and the theory of constrained utility maximization outlined above restricts s to be positive: an increase in the wage rate raises the price of an hour not worked in the market and, at the same level of utility, this induces less consumption of non-market time and more time allocated to market work. At the same time, an

[25] Or the real reservation wage, w^*/p is the value of the real wage such that hours of work are zero exactly, i.e. from eq. (8), $h(1, w^*/p, y;\ A, \varepsilon) = 0$.

increase in the wage rate augments the individual's wealth allowing him to consume more of those things that increase his utility and to consume less of those things that generate disutility (such as hours of market work). This is the income effect of a wage increase on hours of market work and it is given in eq. (11) by $h \cdot \partial h/\partial y$. This term is negative provided nonmarket time is a normal commodity. Consequently, the sign of the uncompensated effect of an increase in the individual's wage rate on his hours of work [the left-hand side of eq. (11)] is indeterminate in sign and depends on the relative magnitudes of the substitution and income effects.

As in other constrained maximizing problems where the constraint is linear, the optimizing eqs. (8) possess a symmetry property according to which $(\partial x/\partial w)_{\bar{u}} = -(\partial h/\partial p)_{\bar{u}}$, where the $\bar{u}$ subscript denotes that these derivatives involve "pure" price changes, i.e. they are evaluated with utility held constant. In addition, under these circumstances of an interior solution to the maximization problem where the constraint is linear and the utility function is quasi-concave, the derived hours of work equation will be a *continuous* function of the budget constraint variables.

Frequently, eq. (11) is expressed in terms of elasticities:

$$E = E^* + (mpe), \tag{12}$$

where $E = (\partial h/\partial w)(w/h) \gtrless 0$ is the uncompensated wage elasticity of hours of work, $E^* = (sw)/h > 0$ is the income-compensated wage elasticity, and $mpe = w \cdot \partial h/\partial y$ is the marginal propensity to earn out of nonwage income. The second term on the right-hand side, *mpe*, is often described in the empirical literature as the "total income elasticity".[26] If both commodities and nonworking time are "normal" (i.e. if both $\partial x/\partial y > 0$ and $-\partial h/\partial y > 0$), then the *mpe* is less than zero but greater than minus unity.[27] If nonworking time is "inferior", then a dollar increase in nonwage income increases the consumption of commodities by more than one dollar.

Substitute the optimizing commodity demand and labor supply functions (8) into the utility function (3) to express the individual's maximized utility as an indirect function of commodity prices, the wage rate, and nonlabor income:

$$V = V(p, w, y;\ A, \varepsilon). \tag{13}$$

This indirect utility function also possesses the zero homogeneity property in p, w, and y: because an equiproportionate change in p, w, and y leaves the optimizing x and h unchanged according to (8), so must the maximized value of

[26] It may be written as the product of $(wh)/y$ and $(\partial h/\partial y)(y/h)$.

[27] Differentiating the budget constraint with respect to y (and in so doing recognizing the dependence of x and h on y) yields the Engel aggregation condition $p(\partial x/\partial y) - w(\partial h/\partial y) = 1$.

utility be unaltered. It is straightforward to show that $\partial V/\partial p = -\lambda x$, $\partial V/\partial w = \lambda h$, and $\partial V/\partial y = \lambda$, where λ is the marginal utility of nonlabor income when evaluating the utility function at its optimum so that, combining these results,

$$
\begin{aligned}
-\frac{\partial V/\partial p}{\partial V/\partial y} &= x(p, w, y;\ A, \varepsilon),\\
\frac{\partial V/\partial w}{\partial V/\partial y} &= h(p, w, y;\ A, \varepsilon).
\end{aligned}
\tag{14}
$$

These equations, Roy's Identity, imply that the functional form of the commodity demand and labor supply equations may be derived relatively easily once a particular form of the indirect utility function, eq. (13), has been specified.[28]

3.2. *Aggregation*

The theory outlined above applies to a single individual. It has often been applied to data that have been aggregated across individuals. Thus, some claim to have estimated the income and substitution effects (or the net wage effect $\partial h/\partial w$) of eq. (8) by using data across industries or occupations and by specifying the dependent variable as the average hours worked of individuals in a given industry or occupation. [For instance, Metcalf, Nickell, and Richardson (1976) and S. Rosen (1969).] Others use time-series observations on average hours worked by all employees (both male and female) in the economy to fit eq. (8). [For instance, Abbott and Ashenfelter (1976, 1979), Barnett (1979, 1981), Darrough (1977), and Phlips (1978).]

There are two issues to address. The first one assumes all individuals occupy an interior solution to their constrained maximization problem and enquires into the conditions under which each individual's labor supply function can be aggregated into a macro labor supply function that possesses the properties of eq. (8). The second and more relevant issue looks into the aggregation problem when some individuals are at a corner solution and others are at an interior solution to their maximization problem.

[28]The dual to the budget-constrained utility maximization problem characterizes the individual as selecting x and h to minimize the net cost, $px - wh$, of attaining a prescribed level of utility. The reduced form equations corresponding to this problem are the utility-constant commodity demand and labor supply functions and if these functions are substituted back into the objective function, $px - wh$, the net expenditure function is derived.

The first issue is not identical to the standard problem in the consumer demand literature because in that literature all consumers are assumed to face the same commodity prices whereas in the labor supply context one price, the wage rate, varies across individuals. The papers listed above using aggregate data to estimate labor supply functions have specified as arguments some average of the wage rates of the workers and an average nonwage income. Therefore, consider the case in which the arithmetic mean of these variables is used in a macro labor earnings equation and in which the macro earnings equation is to be derived by aggregating each worker's labor earnings function. In these circumstances, each worker's *mpe* (marginal propensity to earn $= w \cdot \partial h / \partial y$) must be the same and it must be independent of the wage rate and nonwage income. In addition, the commodity demand functions must be linear in both wages and nonwage income. [See Deaton and Muellbauer (1980, pp. 159–161) and Muellbauer (1981).] These are nontrivial restrictions on the form of the labor supply and commodity demand equations although they do not rule out some interesting cases.[29]

The second aggregation problem has more serious implications and to appreciate these difficulties let us invoke a set of extreme assumptions, namely, that a population of individuals is identical in all characteristics observed by the economist (i.e. they have the same y and A and face the same p and w), but they have different values of the unobserved variable ε. Let $f(\varepsilon)$ be the density of ε in the population. These differences in ε generate a distribution of reservation wages across these individuals. Suppose this distribution of reservation wages is described by the density function $\phi(w^*)$ and suppose $\Phi(w^*)$ is the cumulative distribution corresponding to the density function. The cumulative distribution function $\Phi(\bar{w}^*)$ is interpreted as giving for any value $\bar{w}^*$ the probability of the event "$w^* \leq \bar{w}^*$". The proportion of these individuals who offer positive hours of work to the labor market consists of those whose values of w^* satisfy eq. (9), that is, those for whom $w^* < w$. Equivalently, the labor force participation rate (π) of this group is simply the cumulative distribution of w^* evaluated at $w^* = w$:

$$\pi(p, w, y, A) = \Phi(w; p, y, A),$$

where the dependence of the labor force participation rate on the variables assumed to be the same in this hypothetical population (namely, p, w, y, and A) has been made explicit. Because the cumulative distribution function is necessarily a monotone nondecreasing function [i.e. $\Phi(\underline{w}^*) \leq \Phi(\bar{w}^*)$ for $\underline{w}^* < \bar{w}^*$], an increase in the wage rate offered to these individuals cannot reduce the labor force participation rate:

$$\frac{\partial \pi}{\partial w} = \frac{\partial \Phi(w)}{\partial w} = \phi(w) \geq 0.$$

[29]The labor supply equation derived from a Stone–Geary utility function is a special case of the class of permissible functions that aggregate. The more general class accommodates a wider range of substitution possibilities than does the Stone–Geary.

Exactly how much the labor force participation rate increases (if at all) will depend upon the shape of the density function $\phi(w^*)$ in the neighborhood of $w^* = w$.[30]

The variable most often used in studies of labor supply with aggregated data measures the average hours worked per employee. This may be written

$$\mathscr{E}(h|w > w^*) = \frac{\int h(p, w, y; A, \varepsilon) f(\varepsilon)\,\mathrm{d}\varepsilon}{\pi(p, w, y, A)},$$

where the integration is over all those at work and where the hours of work function is that corresponding to the interior solution of the constrained utility maximization problem, eq. (8). Unless the conditioning event $w > w^*$ is satisfied for the entire population, i.e. unless $\pi = 1$, the partial derivatives of $\mathscr{E}(h|w > w^*)$ are not the same as the partial derivatives of eq. (8), $h(p, w, y; A, \varepsilon)$, and it is the latter to which the income and substitution effects outlined in Section 3.1 relate. Studies that regress average hours worked per worker on average wage rates and nonwage income and that interpret the resulting estimates in terms of income and substitution effects are compounding the effects of changes in these variables on (1) the hours worked by those who are at work both before and after these changes with the effects on (2) the composition of the population between workers and non-workers.

These problems of aggregating over individuals some of whom are occupying interior solutions to their constrained utility maximization problem and others corner solutions are likely to be more innocuous for studies restricted to prime-age males (for whom π does not fall far short of unity) than for those relating to young men, older men, and women. The aggregate time-series studies mentioned above, however, are fitted to data describing all workers, male and female, young and old, urban and rural and for the entire adult population, of course, the labor force participation rate has been substantially less than unity (see table 1.5). At this grand level of aggregation, there are the additional problems raised by the fact that the microeconometric evidence suggests differences in the utility functions of men and women even after allowing for differences in the unobserved components ε. So even though during this century the labor force participation rate of all adults in the United States has changed relatively little, the composition of the labor force has changed considerably: according to the U.S. decennial Censuses, whereas in 1900 some 18 percent of the labor force were women, in

[30] Thus, while an increase in w may increase or may decrease hours worked per employee, an increase in w cannot decrease the fraction of the population at work. On this, see Lewis (1967), Ben Porath (1973), and Heckman (1978). The distinction between the labor force and the number employed is not crucial to this argument. Whether hours spent searching for a job is included in the definition of the offer to sell hours or it is excluded (so that π measures the fraction of the population who are employed), this does not affect the substance of the argument.

1970 the figure had more than doubled to 37 percent. These problems of deriving meaningful behavioral parameters from aggregate time-series data are further aggravated by the difficulties that arise when individuals face different *nonlinear* budget constraints (discussed in Section 3.3) and when the conditions are almost certainly not satisfied for the identification from these data of a labor supply function – after all, while some are regressing hours per worker on the average wage rate and interpreting the results in the terms of the income and substitution effects of a labor supply equation, others are taking virtually the same aggregate data, running very similar regression equations, and interpreting the results in terms of the parameters of a structural labor demand function! Both groups of researchers tend to find a negative partial correlation between hours and wage rates: one group interprets this as a negatively-inclined labor supply function while the other group confirms the existence of an inelastic labor demand function![31] The inescapable conclusion is that the equations fitted to aggregate time-series data are not to be regarded as supplying meaningful evidence on the parameters of behavioral hours of work equations and so, in evaluating the empirical work in Section 4 below, I omit a discussion of the estimates from aggregated data.

A somewhat different set of aggregation issues arises in those few studies that use as the measure of labor supply not average hours worked, but the labor force participation rate of different cities. This procedure was employed by Mincer (1962) in his influential work on the labor supply of married women. He cast the wife's decision-making in a family context and he proposed and implemented a specification that distinguished more clearly than had previous researchers between the income and substitution effects operating on the wife's behavior. In his application, he used as his measure of labor supply the labor force participation rates of married women across different metropolitan areas of the United States. This use of aggregate participation rates as the measure of labor supply was followed in a number of subsequent studies, some of them dealing with the labor supply of men.[32] In these papers, the authors have often interpreted the coefficients on the wage rate and nonwage income variables in terms of the derivatives

[31] Some of the labor demand studies use hours per worker [e.g. Nadiri and Rosen (1974)] as the variable to be explained while others use total manhours [e.g. Sargent (1978)]. In either case an identification problem arises. As an example, compare the work of Abbott and Ashenfelter (1976, 1979) with that of Coen and Hickman (1970). Both use highly aggregated annual observations on variables covering a similar period – from 1929 to 1967 in the case of Abbott and Ashenfelter and from 1924 to 1965 (excluding 1941 to 1948) in the case of Coen and Hickman. Abbott and Ashenfelter maintain they are estimating a labor supply equation in a system of consumer demand equations while Coen and Hickman maintain they are estimating a labor demand equation in a system of input demand functions. In fact, both sets of authors seek to explain first-differences in hours worked, in labor earnings, or in manhours worked. Abbott and Ashenfelter (1979) estimate an uncompensated wage elasticity of the supply of hours worked of -0.07 for the linear expenditure system and of -0.14 for their form of the Rotterdam model. Coen and Hickman's preferred estimate of the elasticity of the demand for manhours with respect to wages is -0.19.

[32] See Ashenfelter and Heckman (1974), Bowen and Finegan (1964, 1969), Greenhalgh (1979), and Kosters (1966, 1969).

for the average individual of the hours of work eq. (8) expressed as a fraction of total time available. What justification can be provided for this?[33]

Assume that the period relevant to the constrained utility maximization problem is the individual's lifetime so that the budget constraint variables are defined in terms of their "permanent" values. The individual then determines the proportion of his life to be spent at market work, the particular timing of that participation being determined (it is assumed) by factors orthogonal to the labor supply problem. In this case, among a group of individuals with the same p, w, and y, the probability that one of them is at market work is the same as the proportion of available lifetime hours allocated to market work. What is crucial in this chain of reasoning is that the proportion of his lifetime supplied to market work (equal by assumption to the participation rate) should correspond to an interior solution to the constrained maximization problem for *all* individuals in the relevant population. Otherwise, instead of eq. (8) being applicable to all individuals, it holds for only a subset of the population with the remainder described by a corner solution, namely eq. (10). In fact, virtually all men in the United States are in the labor force at least part of their lives: according to the 1970 U.S. Census of Population, of all men aged 55 years and over who were not in the labor force during the Census week of 1970 and who responded to the question concerning their last year worked, a little over 1 percent had never worked at all. Although all but a tiny fraction of men work in the market at some stage in their life, there remain a number of heroic assumptions in this chain of reasoning – the particular timing of a person's participation is unlikely to be uncorrelated with the permanent budget constraint variables nor in many applications of this procedure do the authors exercise great care in distinguishing permanent budget constraint variables from their currently observed counterparts – such that it is difficult to accept the interpretation of the coefficients on the wage rate and nonwage income variables in cross-city labor force participation rate equations as the parameters on an hours of work function such as eq. (8).

3.3. Nonlinear budget constraint

Now return to the analysis of the individual's allocation of time and consumption. Section 3.1 assumed the simplest form for the budget constraint according to which each and every hour supplied by the individual to the market is rewarded at the fixed rate w. This assumption does not require that each employer does nothing more than specify for each job a fixed wage per hour, leaving the individual employee to choose how many hours he wishes to work. Even if each employer specified not merely the wage rate but also the number of hours each employee is expected to work, provided the wage offer does not vary

[33] The argument that follows is taken from Heckman (1978).

systematically with the stipulated hours and provided the entire range of hours of work is covered by the employers' offers, then a continuous linear budget constraint arises from the aggregation over many employers' wage-hours packages.

Nevertheless, there seem to be important instances in which a continuous, linear budget constraint does not accurately describe an individual's work-income opportunities and as a result the wage rate can no longer be assumed exogenous to the individual. For instance, the presence of quasi-fixed hiring and training costs that are more closely related to the number of employees rather than to their total hours worked encourages firms to offer higher wage rates for longer hours worked per employee [Lewis (1969)]. If this is the case, the wage-hours contract offered by each employer is such that relatively long work hours are tied to relatively high hourly wage rates and consequently the market hours–wage locus facing an individual worker is no longer linear. Even if the employer–employee contract should grant the employee considerable discretion over his hours of work, some payments systems will result in a nonlinear budget constraint. Such is the case when the employee is rewarded (at least in part) by what he produces on the job (such as with piece-rate systems or sales commissions) and this in turn is not a simple linear function of his hours worked. Furthermore, if it is his after tax compensation that is relevant to the individual's allocation decisions[34] and if the tax rates on his income are not independent of the amount of that income, then again the individual is no longer presented with a linear budget constraint. Even if statutory tax rates did not change with income, effective tax rates might vary because of systematic income tax evasion or because of the latitude exercised by administrators in the tax revenue and welfare disbursement agencies. Finally, there are fixed costs and benefits to working, that is, expenditures and compensation that do not vary over all values of an individual's hours of work. As an example of a fixed compensation, some health insurance schemes are available to each individual workers more cheaply when provided to all employees as a group and these benefits take the form of a lump-sum payment that does not depend upon an individual's precise hours worked (although they are sometimes available only if a certain minimum number of hours are regularly worked). Fixed money costs of work arise from travel expenses or necessary expenditures for the performance of the job; these costs must be incurred if any hours are worked in the market, but once the individual is at work they do not change with the number of hours worked.[35]

The modifications required by a nonlinear budget constraint for the theory of the allocation of time in Section 3.1 depend upon the particular form taken by the budget constraint. There are three cases to be considered: the first is when the budget constraint may be assumed to be fully differentiable and it forms a convex

[34] This has been tested in H. Rosen (1976), Hausman and Wise (1976), and Johnson and Pencavel (1984), all of whom could not reject the hypothesis that the relevant variable was after-tax wages not before-tax wages.

[35] Fixed time costs consist of the expenditure of time in travelling to and from work. For an analysis of these, see Moses and Williamson (1963), Oi (1976), and Cogan (1981).

set (so, if taxes are the cause of budget constraint nonlinearities, they are progressive at all levels of income) in which case the techniques of differential calculus may be applied and local comparisons of $-m$ with the slope of the budget constraint identify the individual's optimum allocation of consumption and work; the second is when the budget constraint forms a convex set, but it is piecewise linear with kinks at various levels of income; and the third is when the budget constraint set is nonconvex because of regressive tax rates or "lumpy" fixed costs. Consider each of these three cases in turn.

Where the budget constraint forms a convex set and where it is continuous throughout and fully differentiable, then once again Kuhn–Tucker methods can be applied to determine whether an individual works in the market and, if he works, the number of hours he chooses. In particular, let c be the individual's total compensation for his market work and let c be a positive function of hours worked, $h: c = c(h; B)$ with $c'(h; B) > 0$, $c''(h; B) < 0$, and where B stands for variables that affect the position of the compensation function and that are exogenous to the individual worker. The individual may now be characterized as choosing $x > 0$ and $h \geq 0$ to maximize $U(x, h; A, \varepsilon)$ subject to the budget constraint $px = c(h; B) + y$. For an interior solution, the negative of the marginal rate of substitution of working hours for commodities, $-m$, equals the real marginal rate of compensation:

$$\frac{c'(h; B)}{p} = -\frac{\partial U/\partial h}{\partial U/\partial x} = -m(x, h; A, \varepsilon).$$

An analogous modification is made to the condition that determines whether an individual will work: if $c'(0; B) \leq w^*$, then $h = 0$. For this type of budget constraint, a typical procedure is to replace the true nonlinear constraint with that artificial linear constraint which would induce the same hours of work by the individual. That is, if $\tilde{h}$ denotes the hours of work and $\tilde{x}$ the commodity consumption bundle that solve the constrained utility-maximization problem and if $\tilde{w} = c'(\tilde{h}; B)$, then the linearized budget constraint is the equation $p\tilde{x} = \tilde{w}\tilde{h} + \tilde{y}$, where $\tilde{y}$ is known as "linearized nonwage income" or, sometimes, as "virtual" income [Burtless and Hausman (1978)] (see Figure 1.1). The hours of work eq. (8) may then be written as $h = h(p, \tilde{w}, \tilde{y}; A, \varepsilon)$.[36]

The problem is only slightly less straightforward in the second case when the income tax system is progressive throughout, but the tax rate rises with income in discrete steps so the budget constraint has linear segments connected by kinks. Each segment of the budget constraint is defined by its real after-tax wage rate

[36] Observe that, because hours of work are affected in part by the unobserved variables ε and the artificial budget constraint is linearized around the observed hours of work, $\tilde{w}$ and $\tilde{y}$ are also going to be affected by ε. Consequently, in estimation, $\tilde{w}$ and $\tilde{y}$ cannot be treated correctly as exogenous variables. Hall (1973), Hausman and Wise (1976), and Rosen (1976) calculate the marginal wage rate and linearized nonwage income not in the manner described, but at the same number of working hours for everyone in their sample. This leads to an analogous sort of inconsistency that comes from not instrumenting the marginal wage and linearized nonwage income variables.

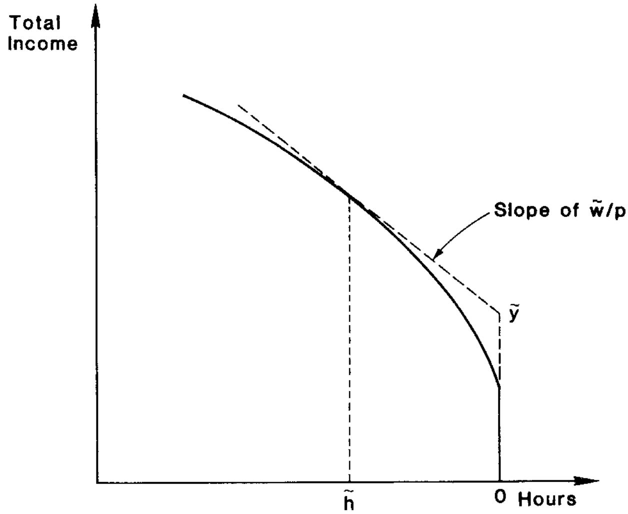

Figure 1.1

and by its real level of linearized nonwage income (i.e. by the height of the nonwage income axis if the slope of the budget constraint is extended to the vertical axis). The familiar tangency condition between the real net wage rate and $-m(x, h; A, \varepsilon)$ holds for any point chosen along one of the linear segments. An individual will locate at any kink if, at this point, his $-m(x, h,; A, \varepsilon)$ lies between the slopes of the budget constraint on either side of this kink. Once again, because the budget constraint is convex, local comparisons of $-m$ with the slope of the budget constraint are sufficient to identify the hours of work corresponding to maximum utility.

Local comparisons of the slope of the indifference curves with the slope of the budget constraint are not sufficient to identify the global utility optimum when the budget set is nonconvex, the third case. Examples of this are provided in Figures 1.2 and 1.3. In Figure 1.2, the income tax system is regressive as is the case when the implicit tax rate on welfare income (received at relatively low levels of total income) exceeds the explicit personal income tax rate. In Figure 1.3, there are fixed money costs of working of the amount ab' so that the budget constraint is $0ab$ if the individual works and $0ab'$ if the individual does not work in the market. For those who work, these fixed money costs are tantamount to a lower level of nonwage income. These fixed costs can be avoided altogether, however, by not working in the market and their lumpiness induces a discontinuity into the hours of work function: if only a relatively small number of hours are worked (relative, that is, to the market wage rate), then insufficient labor income will be earned to offset the fixed money expenditures of working, let alone to compensate for the disutility of market work; once the net wage rate rises sufficiently to

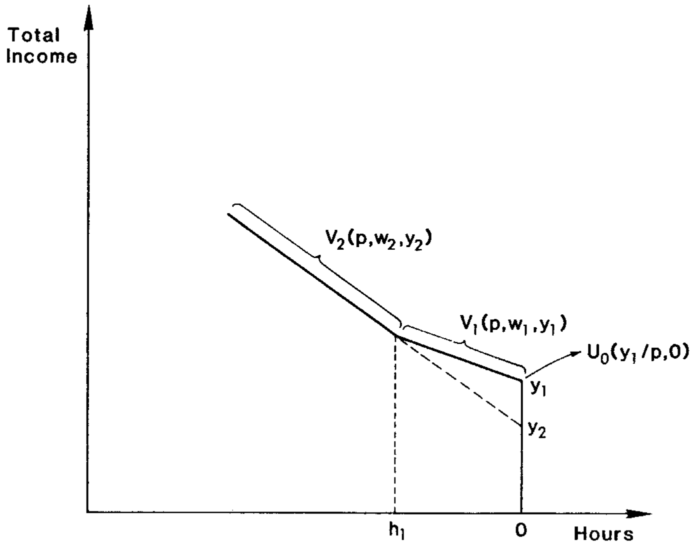

Figure 1.2

induce the individual to work, he works sufficient hours to generate enough labor income to pay the fixed costs of work and to offset the disutility of hours at work. These minimum hours of work are called reservation hours (h_r in Figure 1.3). When the budget constraint is nonconvex, the hours of work function may not be a continuous function of the slope of the budget constraint.

With a nonconvex budget constraint such as that in Figure 1.2, the individual must evaluate his utility at all locations along the frontier of his budget constraint. He is fully capable of doing this because he knows the form of his own utility function, he knows A and ε, and he knows the values of his budget constraint variables. He proceeds by dividing up his utility-maximizing problem into distinct stages, each stage corresponding to a particular corner or segment of his budget constraint. At the first stage he evaluates the utility of not working; in this case his consumption would be y_1/p and his utility would be $U_0(y_1/p, 0; A, \varepsilon)$. At this next stage, he moves to the segment of his budget constraint between 0 and h_1 hours where w_1 is the slope of his budget constraint. Given p, w_1, and y_1 and conditional upon working between 0 and h_1 hours, he could determine whether a tangency condition (a local maximum) obtains between his indifference curve and his budget constraint. It may not, but if it does a maximum level of utility is given by $V_1(p, w_1, y_1; A, \varepsilon)$. He then proceeds to the segment of his budget constraint to the left of h_1 where the net wage is w_2 and linearized nonwage income is y_2. Again, given p, w_2, and y_2 and conditional upon working more than h_1 hours, the individual ascertains whether a tangency condition obtains. If it does, his maximum level of utility is given by

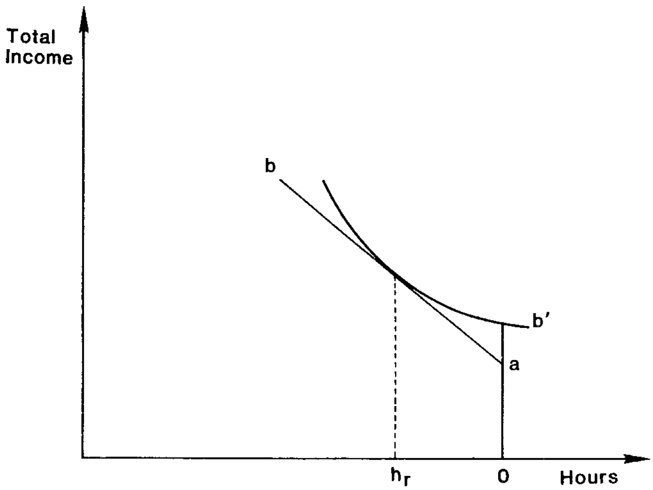

Figure 1.3

$V_2(p, w_2, y_2; A, \varepsilon)$.[37] Having determined the existence of any local maxima in the interior of his budget constraint, if there is more than one, he selects that with the greater utility. He checks to ensure that the utility associated with any interior local maximum exceeds U_0. If no local maximum exists in the interior of his budget constraint, his maximum in Figure 1.2 must be at zero hours of work. If the local maximum in the interior of his budget constraint dominates U_0, then his hours of work are determined by the application of Roy's Identity: $(\partial V_i/\partial w_i)/(\partial V_i/\partial y_i) = h(p, w_i, y_i; A, \varepsilon)$.

Even if the economist knows the form of the individual's utility function, he cannot replicate the individual's procedure exactly unless ε does not exist. This is, in fact, how Wales and Woodland (1979) proceed by presuming full knowledge of each individual's utility function (i.e. they suppress ε) and of his budget constraint, but assuming that there are errors in measuring hours of work, errors that are distributed independently of p, w, y, and A.

3.4. Restrictions on hours of work by employers

The models described to this point are characterized by the fact that an individual faces a budget constraint covering all possible hours of work. As mentioned at the beginning of the previous section, this does not necessarily

[37]If at a higher level of income, another segment of the budget constraint existed with a lower net wage than w_2 (a kink that bent out would exist), the direct utility function at this kink would have to be evaluated.

mean that each employer offers this continuum of possibilities, only that the market as a whole presents this set of opportunities. However, there exists a long tradition in economics of regarding this notion as fanciful and of characterizing the effective choices for the individual as those of working a "normal" or "standard" work schedule (hours per day, days per week, and weeks per year) or of not working at all. The employer may require overtime to be worked during a period of an unusually high level of business activity and may occasionally put his employees on short time when business is unusually slack, but at all times the employee's hours choices (if he works at all) are supplanted by his employer's discretionary actions.

Under these circumstances, the individual's constrained maximization problem consists simply of choosing x and h to maximize $U(x, h; A, \varepsilon)$ subject to the constraints $px = wh + y$, $x > 0$, and h equals either $\bar{h}$ or 0, where $\bar{h}$ denotes the employer's "take-it-or-leave-it" hours. The individual's choice degenerates into a simple comparison between his maximum utility if he works, $\bar{U} = U((w\bar{h} + y)/p, \bar{h}; A, \varepsilon)$ and his utility when not at work $U_0 = U(y/p, 0: A, \varepsilon)$. If it is the case that $\bar{U} > U_0$, $\bar{h}$ could exceed the hours he would choose (given the *same* values of the other exogenous variables) if the employer allowed any hours to be worked. Or, again, if $\bar{U} > U_0$, $\bar{h}$ might fall short of the hours the individual would choose (given the *same* values of the other exogenous variables) if the employer permitted him to work any number of hours the individual wishes. If this is the case, the individual's hours of work do not correspond to a situation in which the slope of the budget constraint is tangent to the individual's indifference curve. This attribute distinguishes this class of models from those in Sections 3.1 and 3.3.

Of course, in any labor market in which these hours of work restrictions are a permanent and regular feature, it is incorrect to specify the other variables constraining an individual's behavior to be the same in the presence of the hours constraints as in their absence. For, in evaluating the pecuniary and non-pecuniary net benefits of alternative jobs, individuals will gravitate towards those employers who fix working hours close to workers' preferences while employers who stipulate unpopular working hours will tend to experience difficulties in recruiting or retaining workers. In this manner, the wage rate will respond to these variations in the supply of workers to different employers and compensating wage differentials will arise. It would be an error, therefore, to estimate market equilibrium models in which workers are characterized as being constrained to work the number of hours mandated by their employers without at the same time treating the wage rate paid to these workers as jointly determined.[38]

Suppose these employer-mandated hours of work restrictions obtain and that an individual determines he is better off by working $\bar{h}$ hours than by not working at all. Let $h_0 = h(p, w, y; A, \varepsilon)$ be the hours this individual would work if the

[38] This is exactly how hours constraints are modelled in Abowd and Ashenfelter (1981).

employer allowed him to work any number of hours. Then the information required to help determine this individual's preferences for work and consumption (given p, w, y, and A) is h_0, but h_0 is not observed and only $\bar{h}$ is available. Under these circumstances, some economists have argued that time spent searching for the desired number of hours should be included in h_0 and they have used the sum of $\bar{h}$ and hours of unemployment, UN, as an estimate of h_0: $\bar{h} + UN = h_0$. [For instance, Cohen, Rea and Lerman (1970), Garfinkel (1973), Greenberg and Kosters (1973), and Hill (1973).] Or when only some unknown fraction, a, of reported hours of unemployment represent the offer to sell labor, observed hours of work ($\bar{h}$) may be expressed as a function of reported hours of unemployment (UN) plus a vector of variables believed to affect the hours an individual would choose to work in the absence of the employer's mandates:[39]

$$\bar{h} = h_0(p, w, y; A, \varepsilon) - a(UN). \tag{15}$$

Stochastic versions of this equation have been estimated by Dickinson (1974), Morgan (1979), Kalachek, Mellow and Raines (1978), Ashenfelter and Ham (1979), and Ashenfelter (1980). Whereas the earlier papers took account of unemployment in this way on the argument that they would measure more accurately or confidently conventional income and substitution effects, the more recent literature has interpreted the stochastic version of eq. (15) as "...a method for testing whether measured unemployment may be thought of as involuntary" [Ashenfelter (1978)]. According to this argument, "if, on the one hand, measured unemployment is simply another name for voluntary non-market time", then a should be zero; "if, on the other hand, measured unemployment is closely related to the extent to which workers face constraints on their labor market choices," then a should be positive. In fact, with cross-section data, a has been estimated as greater than unity [Dickinson (1974)], as almost exactly unity [Morgan (1979)], as 0.92 [Kalachek, Mellow and Raines (1978)], and as about 0.78 [Ashenfelter and Ham (1979)]; with aggregate time-series data [Ashenfelter (1980)],[40] the estimates of a ranged from 0.36 to 0.48 (with an estimated standard error of about 0.18) when the unemployment variable was treated as exogenous and to be equal to 0.04 (with a standard error of 0.23) when the unemployment variable was instrumented.

In view of the reams written on the subject of "voluntary" and "involuntary" unemployment, the proposal of resolving the empirical relevance of the issue simply by determining whether the coefficient a in eq. (15) is estimated to be zero

[39]It is unfortunate that the utility function and budget constraint underlying eq. (15) are not written down explicitly because it is not obvious how UN enters either the objective function or the constraint. Without knowing that, the behavioral interpretation of eq. (15) is difficult to discern.

[40]As Ashenfelter (1978) himself recognizes, the results from the aggregate time-series analysis were never in doubt: as the estimates of eq. (2) in Section 2.1 made clear, hours of work move closely with the unemployment rate over the business cycle whereas wage rates, nonlabor income, and commodity prices [the other right-hand side variables in eq. (15)] display considerably less business cycle variability. It is claimed that eq. (15) provides a structural explanation for this association.

must have considerable appeal to the profession.[41] Unfortunately, it is not so straightforward a matter for, according to the view that "measured unemployment is simply another name for voluntary non-market time", the duration of unemployment represents one part of the individual's optimal allocation of time and income and, as such, is jointly determined with hours of work and commodity consumption. According to this view, given the variations in individuals' hours of work left unaccounted for by the typical variables available to the economist, it is by no means surprising that, even after removing the influence of the variables p, w, y, and A, one object of choice in this allocation problem (hours of unemployment) is correlated with another dimension (hours of work). Would the existence of a partial correlation across households between expenditures on food and expenditures on clothing necessarily imply that clothing is rationed? The more relevant test is not whether a is zero, but rather a test of whether UN is endogenous.[42] However, this test comes up against the serious problem of an appropriate instrumental variable: what is the variable that can be validly excluded from an hours of work equation and that, at the same time, accounts for variations in the duration of unemployment? I do not know of one.[43] If this is so, then we are not capable of discriminating between the two different characterizations of unemployment.[44]

[41] When asked by some surveys, many individuals claim they would like to work a different number of hours from those they are currently working and some economists infer from this that the model in this section is the relevant one. This is surely an incorrect inference. It is not clear how the respondent interprets the question, but it is likely he answers the question assuming all other variables remain constant. In this case this may only mean that employers are not indifferent to the number of hours that their employees work. If the market offers tied wage-hours packages and the worker selects the best combination of wages and hours on his opportunity locus, then the relevant model is that in Section 3.3 above.

[42] Deaton (1982) also makes the argument that the relevant test in this context is an exogeneity test. In his case, he notes that, when commodities and hours of work are weakly separable in the utility function, the commodity demand equations may be written as a function of the prices of each commodity and of total income, $wh + y$, instead of as a function of w and y separately. When h is freely chosen, $wh + y$ is endogenous. Provided commodities and hours are weakly separable, the form of the commodity demand functions is the same whether hours are constrained or not. Using data on 1,617 households from the British Family Expenditure Survey, Deaton estimates such a system of hours-constrained commodity demand equations where an instrument for total income, $wh + y$, is provided by $w\hat{b} + y$, $\hat{b}$ being a parameter of the preference structure as estimated from the unconstrained version of the model. The results are ambiguous though Deaton infers they slightly favor the model characterizing hours of work as unconstrained. As he fully recognizes, there are a number of stringent assumptions in Deaton's application of this procedure and, indeed, the weak separability hypothesis is itself decisively rejected, but future work may be able to relax some of these assumptions and a modification of this methodology may yield some insights.

[43] In Ashenfelter's (1980) aggregate time-series study, the instrumental variables consisted of higher order terms of the wage rate, nonlabor income, and the prices of commodities. As he himself observes, the validity of these variables as instruments leans heavily on having identified correctly the functional form of the hours of work equation and, because we are not at all confident of the appropriate functional form, these variables are not very satisfactory instruments. In his study of individuals, Ham (1983) proposed using industry, occupation, and local unemployment rates as instruments for each individual's unemployment experience. Whether these are valid instruments depends upon the interpretation of the stochastic error term in the hours of work equation. In Ham's analysis (as in the model of labor supply in this survey paper), the error term represents variations in preferences that are unobserved to the researcher. It is unlikely that the distribution of these "tastes

A more fundamental issue that this rationing literature on hours of work does not address is the relevant wage rate at which individuals are being rationed. When dealing with rationed commodities where all consumers face the same prices, the price of a rationed commodity may be well defined. But in the case of individuals facing different wage rates, it is crucial that we identify the wage rate when rationed. In other words, in accounting for observed, rationed, hours of work $\bar{h}$ in eq. (15), what is the relevant wage rate, w, on the right-hand side? In aggregate studies such as Ashenfelter's, the wage rate used is the average wage received by all those at work and not unemployed so the implicit assumption is that rationed individuals and unrationed individuals face the same exogenous wage. In studies of this kind when data on individuals are used, for individuals experiencing some unemployment the wage rate that rations these men when unemployed is assumed to be the same as the wage rate they receive when employed. What is the appropriate rationed wage rate when an individual experiences no spell of employment and is always recorded as unemployed? Such individuals are deliberately excluded from these studies.[45] Because no exchange of labor services takes place while an individual is unemployed, no wage rate is recorded and observationally this is equivalent to the situation that arises when the reservation wage, w^*, exceeds the offered wage, w. In other words, the situation is observationally equivalent to what some economists call "voluntary unemployment".

3.5. *Life-cycle models*

All the models outlined above have been static, one-period descriptions of behavior. An important development in research on labor supply over the past

for work" parameters is independent of the unemployment experiences of these men; that is, those men with greater tastes for leisure will tend to take longer or more frequent spells of unemployment. Then, if industry, occupation, and local unemployment rates are correlated with the unemployment experiences of individual men (as Ham maintains), then these unemployment rates must also be correlated with the utility function parameters imbedded in the error term of the hours of work equation. In other words, these unemployment rates do not serve as appropriate instruments.

[44]A different procedure for testing for the presence of employer-mandated restrictions on hours of work is contained in Ham (1982). For a sample of prime-aged male workers experiencing no unemployment and claiming no underemployment, he estimates a labor supply function that allows for the possibility of sample selection bias resulting from excluding these unemployed and underemployed workers. He then tests whether the estimates that make no adjustment for the exclusion of the unemployed and underemployed differ significantly from those that do make that adjustment. He finds a significant difference and argues that the differences move in the direction suggested by the proposition that these unemployed and underemployed workers are constrained by employers' restrictions on hours of work.

[45]For example, Lundberg (1983) writes: "The sample was restricted to two-head households in which both husband and wife worked at some time.... The exclusion of these households was...to ensure a wage observation for each individual."

ten years has been the specification and estimation of life-cycle, multi-period, models according to which consumption and labor supply decisions in each period are made with regard to prices and wage rates in all periods. Utility is defined over lifetime consumption and lifetime hours of work and similarly the budget constraint incorporates incomes and expenditures in different periods plus the opportunity to reallocate incomes and expenditures across periods by borrowing and lending. Whereas in the static models discussed above interest and dividend income from previous savings decisions was treated as exogenous, in a life-cycle context it becomes endogenous and only inherited assets and unanticipated net returns on capital are genuinely exogenous. The life-cycle counterparts to eqs. (8) in Section 3.1 relate consumption and hours worked at age t to prices and to wage rates at each and every age where future budget constraint variables are appropriately discounted to the present.

The notion that an individual's or a household's consumption and working decisions are made with the future very much in view squares with some basic patterns of life-cycle behavior. The prototype is described by a young married couple starting out with few assets and working long hours, a portion of these hours representing on-the-job training; then moving to a higher asset position, continuing to work long hours (at least for the man) and starting to raise a family with the implied financial responsibilities for the future; and later in life working fewer hours and concomitantly running down their assets. Also, recall from Section 2.2 above that in U.S. cross-section data both male labor force participation probabilities and male hours of work display an inverted-U shape with respect to age. Hourly wage rates also map out an inverted-U shape with age although the peak in hours worked precedes the peak in wage rates.[46] The correspondence of the hours and wage profiles with respect to age conforms to the most basic implication of the life-cycle labor supply model, namely that an individual will supply more hours to the market during those periods when his wage rate is highest; this is the effect of evolutionary wage differences on hours worked. The hours–age and wage–age profiles of black men are flatter than those of white men with the peaks of both profiles occurring at younger ages for black men than for white men. Weekly hours and weekly wages also follow an inverted-U shape with respect to age in British data presented by Browning, Deaton and Irish (1983). They present these graphs separately for manual and nonmanual workers: for manual workers, wages peak a little later than hours; for nonmanual workers, the peaks in the two series are roughly coincident. At all ages, manual workers have higher hours and lower wages than nonmanual

[46]According to the life-cycle interpretation, the fact that the peak in hours worked precedes the peak in wage rates implies that the rate of interest exceeds the individual's rate of time preference. Weiss (1972) expresses this well: "The rate of interest induces an early work effort since labour earnings can be invested at a higher rate of return. The subjective discount rate induces the postponement of work since future effort seems less painful when viewed from the present."

workers. The life-cycle model of labor supply outlined below is an attempt to provide an explicit and formal characterization of these empirical regularities.[47]

The empirical implementation of the life-cycle model would appear to require a great volume of data: to understand an individual's labor supply today, the economist needs information on prices and wages throughout the individual's life! In fact, the empirical work on life-cycle labor supply has proceeded by placing sufficient restrictions on the form of the lifetime utility function that the parameters governing the dynamic allocation of consumption and hours can be estimated with relatively little data. To date, there exist two general approaches to this dynamic allocation problem. One derives from the literature on habit persistence and stock adjustment and specifies the individual's utility function in period t as conditional on the individual's consumption and hours of work in the previous period. The notion that the standards by which individuals gauge their welfare are molded by their prior experiences is, of course, an old one. Preferences displaying this state dependence in the labor supply literature have been estimated at the aggregative level by Philips (1978) and employed in aggregate business cycle simulations by Kydland and Prescott (1982) and they have been estimated with individual panel data by Hotz, Kydland and Sedlacek (1982) and Johnson and Pencavel (1984).[48] Whereas in this specification the lifetime utility function is intertemporally not (strongly) separable,[49] the opposite hypothesis is maintained in the second approach to the individual's life-cycle labor supply problem. Substantially more research has been conducted along the lines of the second approach and so I proceed to outline its central features in a little more detail.

Assume the lifetime utility function is additive over time and write the individual's utility in period t as a strictly concave function of commodities consumed in period t, x_t, and of hours worked in period $t, h_t: U_t(x_t, h_t; A_t, \varepsilon_t)$, where, as before, A_t denotes exogenous variables observed by the researcher while ε_t is a component unobserved by the researcher. Let the rate of time preference be given by ρ and suppose a fixed "lifetime" of $N+1$ periods. Then the individual's utility function is

$$\sum_{t=0}^{N} (1+\rho)^{-t} U_t(x_t, h_t; A_t, \varepsilon_t). \tag{16}$$

[47]The life-cycle model would attribute the greater hours with lower wages of British manual workers compared with nonmanual workers in terms of the greater life-cycle wealth of the latter.

[48]The interpretation of these models in the interesting special case of the Stone–Geary utility function is provided in the papers by Phlips and Spinnewyn (1982) and Pollak (1970).

[49]However, when a consumer fully recognizes the evolution of his tastes as he ages, Spinnewyn (1981) shows that the intertemporal model of consumer behavior with habit persistence can be transformed into a model without such persistence by a suitable redefinition of the cost of consumption and wealth.

The lifetime budget constraint is

$$K_0 + \sum_{t=0}^{N} (1+r)^{-t}(w_t h_t - p_t x_t) = 0, \tag{17}$$

where K_0 denotes initial wealth and r is the rate of interest which for convenience is assumed to be fixed. Bequests have been neglected although it is straightforward to permit a role for them. The individual selects $x_t > 0$ and $h_t \geq 0$ for each period to maximize (16) subject to the constraint (17), the first-order conditions for which are eq. (17) and

$$\frac{\partial U_t}{\partial x_t} = \theta^t \lambda_0 p_t, \qquad t = 0,\ldots,N, \tag{18}$$

$$-\frac{\partial U_t}{\partial h_t} \geq \theta^t \lambda_0 w_t, \quad t = 0,\ldots,N, \tag{19}$$

where $\theta = (1+\rho)/(1+r)$ and where λ_0 is the Lagrange multiplier attached to the budget constraint and is interpreted as the marginal utility of initial wealth when evaluating the utility function at its optimum. If eq. (19) is a strict inequality, the individual does not work in period t; if it is an equality, then some hours of work are supplied to the market. In what follows, given the high labor force participation rates of price–age men, I assume (19) is satisfied by an equality.

Now solve eqs. (18) and (19) for consumption and working hours in any period:

$$x_t = x\left(\lambda_0 \theta^t p_t, \lambda_0 \theta^t w_t; A_t, \varepsilon_t\right), \quad t = 0,\ldots,N, \tag{20}$$

$$h_t = h\left(\lambda_0 \theta^t p_t, \lambda_0 \theta^t w_t; A_t, \varepsilon_t\right), \quad t = 0,\ldots,N. \tag{21}$$

In these equations, λ_0 is endogenous, a function of the lifetime budget constraint variables and of A_t and ε_t. Indeed, it can be shown that $\partial \lambda_0 / \partial K_0 < 0$, $\partial \lambda_0 / \partial w_t \leq 0$, and $\partial \lambda_0 / \partial p_t > 0$ [see Heckman (1974a, 1976a)]. Eqs. (20) and (21) have been called "λ_0-constant" functions or, more felicitously, Frisch demand and supply functions [Browning (1982)] in recognition of Ragnar Frisch's extensive use of additive utility functions. Given the assumed concavity of the utility function, these Frisch demand and supply functions possess many of the properties of conventional demand and supply functions: $\partial h_t / \partial(\lambda_0 \theta^t w_t) > 0$ and $\partial x_t / \partial(\lambda_0 \theta^t p_t) < 0$; there is a symmetry property $-\partial h_t / \partial(\lambda_0 \theta^t p_t) =$

$\partial x_t/\partial(\lambda_0\theta' w_t) \gtrless 0$; and these functions are homogeneous of degree zero in λ_0^{-1}, p_t, and w_t[50] [see Heckman (1974a)].

Because the period-specific utility function represents one branch of the entire lifetime utility function, the Frisch labor supply eq. (21) is not independent of monotonic transformations of U_t. The important feature of these equations for empirical analysis is that they relate consumption and labor supply decisions in any period to variables outside that period only through λ_0 and that otherwise within-period prices and wage rates determine x_t and h_t. The variable λ_0 is a sufficient statistic in that it contains all the information concerning the lifetime budget constraint variables which is relevant to the current choice of consumption and hours of work. Moreover, although λ_0 varies across individuals in accordance with differences in their lifetime budget constraint variables and in other exogenous variables, for a given individual λ_0 is constant over his lifetime when future wages and prices are known with certainty. The derivative of h_t with respect to w_t in eq. (21) shows how an individual's hours respond to *evolutionary* wage changes, i.e. changes in wages along on individual's wage–age profile. As MaCurdy (1982) has emphasized, corresponding to the two classes of variables (the current period variables and the life cycle component, λ_0) in eqs. (20) and (21), the formulation of an empirically tractable model of life-cycle behavior naturally decomposes into two stages: the first is the specification of the Frisch equations and the second is the formulation of an equation to determine λ_0.

At the first stage, the immediate problem is, of course, that λ_0 is not directly observed. Moreover, λ_0 is not a random variable uncorrelated with wages and prices. Because it is not random, it cannot be consigned to some error term. However, as we shall see below, for certain forms of the Frisch equations, λ_0 (or a simple transformation of λ_0) may be expressed as an additive fixed effect that, in estimation with panel data, is easily accounted for by first-differencing the data over time.[51] The second stage of the estimation procedure relates λ_0 to its

[50] Instead of obtaining the Frisch equations by solving the first-order conditions from explicit constrained utility maximization, Browning (1982) shows they may be derived more simply by defining a consumer's within-period profit function as follows:

$$\Pi_t\left(\lambda_0^{-1}, \tilde{p}_t, \tilde{w}_t; A_t, \varepsilon_t\right) = \max_{x_t, h_t}\left\{\lambda_0^{-1} U_t(x_t, h_t; A_t, \varepsilon_t) - \tilde{p}_t x_t + \tilde{w}_t h_t\right\},$$

where $\tilde{p}_t = \theta' p_t$, $\tilde{w}_t = \theta' w_t$, and naturally λ_0^{-1} may be called "the price of utility". Then, by applying the envelope theorem to this profit function, the negative of eq. (20) is derived from $\partial \Pi_t/\partial \tilde{p}_t$ and eq. (21) is derived from $\partial \Pi_t/\partial \tilde{w}_t$. As is the case for a price-taking firm's profit function, this consumer's profit function is increasing the price of output (λ_0^{-1}), is decreasing in the prices of inputs ($\tilde{p}_t$ and $-\tilde{w}_t$), and is convex and linearly homogeneous in λ_0^{-1}, $\tilde{p}_t$, and $\tilde{w}_t$.

[51] If the Frisch hours eq. (21) is written with hours or earnings on the left-hand side (as distinct from some transformation of them such as their logarithms), then for λ_0 to be specified as an additive fixed effect the within-period utility function must be quasi-homothetic in commodities consumed and hours worked. See Browning, Deaton and Irish (1983).

determinants, namely the lifetime budget constraint variables, the rate of time preference, A_t, and ε_t. Observations on the entire budget constraint variables are, of course, not available so lifetime profiles must be simulated by using the observed income and wage data of people of different ages. Moreover, an explicit, closed-form solution for λ_0 is often not possible so instead the expression for λ_0 is approximated. Clearly, this second stage is less cleanly specified and estimated than the first stage, but knowledge of λ_0 is essential to describe an individual's labor supply response to *parametric* wage changes, i.e. wage changes that shift the entire wage–age profile.

A model involving decision-making over time would appear to require allowance for uncertainty about the future values of variables and an important aspect of this life-cycle model is that it accommodates such uncertainty in a tractable form. To see this, first rewrite the certainty model by defining by recursion $\lambda_0(1+\rho)^t/(1+r)^t = \lambda_t$ and so the first-order conditions eqs. (18) and (19) become:

$$\frac{\partial U_t}{\partial x_t} = \lambda_t p_t, \tag{22}$$

$$-\frac{\partial U_t}{\partial h_t} = \lambda_t w_t, \tag{23}$$

$$\lambda_t = \left(\frac{1+r}{1+\rho}\right)\lambda_{t+1}, \tag{24}$$

where λ_t is the marginal utility of wealth in period t. The Frisch demand and supply functions eqs. (20) and (21) are the same with λ_t replacing $\lambda_0\theta^t$:

$$x_t = x(\lambda_t p_t, \lambda_t w_t; A_t, \varepsilon_t), \quad t = 0,\dots,N, \tag{25}$$

$$h_t = h(\lambda_t p_t, \lambda_t w_t; A_t, \varepsilon_t), \quad t = 0,\dots,N. \tag{26}$$

Eq. (24) defines the optimal savings strategy and the lifetime problem decomposes into two levels. At the first level, an individual allocates his wealth over his life such that his marginal utility of wealth evolves as he ages according to eq. (24). At the second level, conditional upon wealth allocated to a given period, the within-period allocation problem is addressed. Strong separability of the lifetime utility function is more than sufficient to decentralize the life-cycle problem in this way.[52]

Now allow for uncertainty in the form of the individual being unsure of real wages or real rates of interest or even his preferences in the future. In these

[52] In fact, weak separability is sufficient and necessary. See Blackorby, Primont and Russell (1975).

circumstances, suppose the consumer revises his plans each period as new information on these variables is revealed and, in particular, suppose he maximizes his current and discounted expected utility subject to his period-by-period budget constraint. The first-order conditions describing the solution to this problem are identical to eqs. (22) and (23), but eq. (24) is now modified to read

$$\lambda_t = (1+\rho)^{-1}\mathscr{E}[(1+r)\lambda_{t+1}],$$

where r in this formula is the rate of return to be paid on each dollar of assets held at the beginning of period $t+1$. Because both r and λ_{t+1} are random, $\mathscr{E}[(1+r)\lambda_{t+1}]$ will typically involve the covariance between these two terms, but if a riskless rate of return exists, say, $\bar{r}$, then the previous equation may be written

$$\lambda_t \frac{1+\rho}{1+\bar{r}} = \mathscr{E}(\lambda_{t+1}) \tag{27}$$

or the expected (at period t) marginal utility of wealth in period $t+1$ is proportional to the marginal utility of wealth in period t, similar to a Martingale stochastic process [MaCurdy (1976)].[53] The consumer's savings policy implies that the means of all future values of λ are revised to account for all forecasting errors at the time they are realized. And because λ_t is a sub-Martingale, through eqs. (22) and (23), $(\partial U/\partial x_t)/p_t$ and $(\partial U/\partial h_t)/w_t$ also follow a sub-Martingale. So, according to this model, at the start of the life-cycle the consumer sets λ_0 so that it takes account of all the information on the future values of variables available at that time. As new information is acquired over time so λ_t is revised according to eq. (27). At each age, in order to satisfy eqs. (22), (23), and (27), the consumer requires knowledge of the variables observed in that period to determine his optimal consumption and hours of work and to update his marginal utility of wealth. Consequently, whereas eqs. (20) and (21) form the basis of empirical work of life-cycle labor supply under the assumption of perfect foresight, eqs. (25) and (26) constitute the analogous equations under conditions of uncertainty.

In the presence of uncertainty, when estimating an equation based on eq. (26), the error term will include forecast errors and, because w_t, A_t, and ε_t contain components unforeseen before their realization, w_t (even if measured without error), A_t, and ε_t will not be distributed independently of the equation's disturbance. Finding variables that are correlated with w_t and A_t and yet are uncorrelated with unanticipated components of these variables (i.e. finding genuine instruments) is difficult.

[53] This result and the conditions underlying it were derived by MaCurdy (1976). That all prices follow a Martingale or sub-Martingale process was conjectured by Alchian (1974).

Certain features of the life-cycle model have considerable appeal. For instance, anyone who has estimated static labor supply functions can testify to the awkward problems in deriving an accurate measure of nonwage income, that is, the income an individual would receive at $h = 0$. The life-cycle model avoids these difficulties. Whereas the static model has to be augmented with explanations in terms of family responsibilities in order to account for the age-pattern of hours of work, the life-cycle model addresses this empirical regularity explicitly.[54] Few would deny that there are circumstances in which the future values of certain variables affect current working decisions. The more pertinent issues are, first, whether these effects are sufficiently important to account for the key variations in male labor supply and, second, whether the particular model sketched above incorporates the essential features of intertemporal decision-making. We shall return to these two issues when the empirical work on life-cycle labor supply is discussed in Section 5 below.

4. Estimation of the static model

4.1. Specification

What guidance has the theory of labor supply outlined in the previous section provided for empirical work? As far as the conventional static model is concerned, I know of no attempts with individual data to specify all of the refutable implications of the theory – the positivity of the substitution effect, the symmetry condition, the zero homogeneity condition – as a series of research hypotheses that are either corroborated or refuted by the data.[55] This is surely surprising in view of the extensive literature that has been concerned with testing the predictions from the consumer's allocation problem (without the hours of work dimension) and that has done so by applying the theory to data aggregated over individuals. The availability of data sets containing observations on the actual decision-making units, the individual or the household, and on the same individu-

[54] The distinctive age-hours of work pattern is apparent in Current Population Survey data organized by Smith (1983). She presents data on annual hours of work by age, by sex, and by race from the four Surveys from 1977 to 1981. For instance, for all men in 1981 (unadjusted for all other characteristics) those aged 16–17 years were estimated to work an average of 715 hours, 18–19 years worked 1209 hours, 20–24 years worked 1634 hours, 25–34 years worked 2016 hours, 35–44 years worked 2126 hours, 45–54 years worked 2108 hours, 55–59 years worked 2037 hours, 60–64 years worked 1839 hours, and those 65 years and over worked 1241 hours.

[55] Occasionally one or other of these implications has been tested. For instance, Wales and Woodland (1976) determined in their husband and wife joint allocation model whether the matrix of compensated wage and price elasticities was correctly signed. For approximately half of their observations it was and for the other half it was not.

als over time means that the observable implications of the theory do not need to be augmented by a series of heroic aggregation assumptions in order to subject the theory to empirical scrutiny. Of course, many other problems remain in implementing the theory, but these turn out not to be specific to labor supply issues and they are rarely resolved except under exceptional circumstances by applying the theory to data aggregated over individuals.

While the implications of the conventional theory of labor supply have rarely been modelled as a series of testable hypotheses, researchers do not seem to be reluctant to treat the qualitative implications of the theory as maintained hypotheses. For instance, Burtless and Hausman (1978) estimate a labor supply model that allows for a distribution across individuals of values for the effect of nonwage income on hours, but in doing so they constrain this effect to be nonpositive. In fact, the estimates of this effect pile up close to zero and one wonders how many individuals would have positive values if the estimation scheme did not prohibit it.[56] In many studies, it seems as if estimates that do not generate positive substitution effects for hours of work or that suggest nonmarket time is an inferior good are not interpreted as refutations of the theory, but as indicating some error in implementing the theory. This is, of course, supposed to be an attribute of a discipline in its "normal science" phase although some would question quite legitimately whether the conventional model of labor supply had earned the right to this status.

Perhaps the primary contribution to date of the theory to empirical research on labor supply has been that of distinguishing the effects on hours of work of changes in wage rates from changes in nonwage income. Although this may appear a trivial contribution, it distinguishes the economist's approach to the topic of market work behavior from that of most other social scientists.[57] Moreover, as Mincer (1963) showed, the distinction may be usefully applied to understanding other patterns of behavior besides hours of work.

Although there have been a number of instances to the contrary, the general procedure has *not* been to specify a particular expression for the direct or indirect utility function (or expenditure function) and then to estimate the implied hours of work function. More often, an hours of work function convenient for estimation has been specified ab initio and the popular choice has been one that is linear in the parameters. That is, eq. (8) has been specified as follows:

$$h_i = \alpha_0 + \alpha_1\left(\frac{w}{p}\right)_i + \alpha_2\left(\frac{y}{p}\right)_i + \alpha_3 A_i + \varepsilon_i, \tag{28}$$

[56] They report that about one-fifth of the sample has an estimated elasticity of hours with respect to nonwage income of between -0.01 and zero. Their restriction on the effect of nonwage income on hours of work arises from the *global* requirement on their estimating technique that the substitution effect be non-negative for all individuals and for all values of the exogenous variables.

[57] For instance, see the interesting sociological study of labor supply in Smith-Lovin and Tickamyer (1978).

where i denotes individual i. In this form, ε_i is a stochastic disturbance term representing individual i's unobserved "tastes for work" and the zero homogeneity condition is a maintained hypothesis. Normalizing p to unity, the uncompensated wage effect is $\alpha_1 \lesseqgtr 0$ while, provided leisure is not an inferior good, $\alpha_2 \leq 0$. Consequently, the substitution effect, s, is given by $\alpha_1 - h\alpha_2$ which should be positive according to the allocation model outlined in Section 3.1 above. Provided $\alpha_1 > h\alpha_2$, eq. (28) implies a larger substitution effect for those who work longer hours.

Because any labor supply equation possessing all the properties of utility-maximizing hours of work functions implies a particular expression for the direct utility function, one may derive the form of the utility function when a linear hours of work equation such as (28) is specified:

$$U(x, h; A, \varepsilon) = \left(\frac{\alpha_2 h - \alpha_1}{\alpha_2^2}\right)\exp\left\{\frac{\alpha_2(\alpha_0 + \alpha_2 x + \alpha_3 A + \varepsilon) - \alpha_1}{\alpha_2 h - \alpha_1}\right\},$$

where $\alpha_1 > \alpha_2 h$.[58] Although x and h do not appear symmetrically in this unfamiliar utility function and although the error term occupies an unintuitive role, these will be small considerations if it is important to have a convenient hours of work estimating equation.

Questions concerning the form of the utility function, however, have received little attention compared with the research investigating the consequences of the error term, ε_i. The reason for this concern is that eq. (28) describes only those men whose optimizing problem is solved by working a positive number of hours; for others, the individual's problem is solved by setting h to zero. In other words, letting αX_i stand for the deterministic part of the right-hand side of eq. (28), the correct specification is as follows:

$$h_i = \alpha X_i + \varepsilon_i, \quad \text{if } w_i > w_i^*(p_i, y_i, A_i, \varepsilon_i), \tag{29}$$

$$h_i = 0, \quad \text{if } w_i \leq w_i^*(p_i, y_i, A_i, \varepsilon_i), \tag{30}$$

where the dependence of the reservation wage, w_i^*, on p_i, y_i, A_i, and ε_i has been made explicit. Clearly, if observations on only those men for whom $h_i > 0$ are used to estimate (28) by ordinary least squares, then $h_i > 0$ implies $\alpha X_i + \varepsilon_i > 0$ or $\varepsilon_i > -\alpha X_i$. Thus, when restricting the estimation of (28) to the sample of working men, ε_i is not distributed independently of X_i even though ε_i may be distributed randomly in the population; because $\mathscr{E}(\varepsilon_i | X_i) \neq 0$, one of the conditions under which ordinary least-squares provides a consistent estimator is violated. Expressed differently, when eq. (28) is fitted to the sample of working

[58] This is derived in Deaton and Muellbauer (1981, p. 96) and in Hausman (1981). Deaton and Muellbauer (1981) consider the case when the composite commodity theorem does not hold and the different components of x are identified.

men, observations are not selected from the population randomly, but systematically according to the requirement $\varepsilon_i > -\alpha X_i$ and a *sample selection bias* results.[59] The magnitude of the bias is likely to be less serious for those samples from populations for which most observations satisfy the criterion $w_i > w_i^*$. In other words, the least-squares selection bias is likely to be more important in describing the hours of work behavior of older and younger men than of prime-age males.[60]

An alternative and insightful characterization of this sample selection problem [attributable to Heckman (1976b)] recasts the issue as a conventional case of omitting a term from a least-squares regression equation. Define $\Delta w_i = w_i - w_i^*$ and observe that $\Delta w_i > 0$ if the individual works in the market [so that eq. (29) holds] while $\Delta w_i \leq 0$, if $h_i = 0$. Denote the determinants of Δw_i by Z_i which will include p_i, y_i, A_i, and ε_i as well as the variables influencing the offered wage rate:

$$\Delta w_i = \delta Z_i + u_i,$$

where u_i is a random variable assumed to have expectation zero and finite variance. Then, the regression of h_i given X_i over the sample of workers (i.e. over the sample for whom $\Delta w_i > 0$) is

$$\begin{aligned} \mathscr{E}(h_i | X_i, \Delta w_i > 0) &= \alpha X_i + \mathscr{E}(\varepsilon_i | u_i > -\delta Z_i) \\ &= \alpha X_i + \phi(\delta Z_i, \xi_i), \end{aligned} \tag{31}$$

where ξ_i denotes the parameters governing the joint density of ε_i and u_i. Because Z_i incorporates the effects of ε_i, the expected value of ε_i given $u_i > -\delta Z_i$ will not be zero. Applying ordinary least squares to (31) is equivalent to omitting the term ϕ, the conditional mean of ε_i, from the regression and thus the bias that results may be understood in terms of conventional omitted-variable bias arguments.

For instance, consider a variable such as nonwage income, y, that appears in both X_i and Z_i. A least-squares regression of h_i on X_i for a sample of workers that omits the conditional mean of ε_i, ϕ, results in estimates of the coefficient on nonwage income, say $\hat{\alpha}_2$ from eq. (24), that may be written approximately as

$$\hat{\alpha}_2 = \alpha_2 + \partial\phi/\partial y.$$

[59]The sample selection bias is not solved by Hall's (1973) procedure of fitting eq. (28) to workers and nonworkers together (setting h to zero for nonworkers). This procedure requires that eq. (29) hold not for $w > w^*$, but for $w \gtrless w^*$, a requirement that contradicts the theoretical structure.

[60]In the labor supply case, the sample selection problem is further complicated by the absence of observations on one of the independent variables, the wage rate facing (and not being accepted by) nonworkers. In his study of married women, Heckman (1974b) proposed and implemented a model that combines an equation determining wage rate offers with an equation determining the marginal rate of substitution of hours for commodities. Both equations were characterized by errors that were correlated with the exogenous variables because of sample selectivity problems.

The coefficient α_2 measures the effect of nonwage income on hours worked on the part of those already working and this is the derivative that figures in the analysis of interior solutions to the individual's constrained utility-maximization problem. This analysis suggests that, provided leisure is not an inferior good, $\alpha_2 \leq 0$. The term $\partial\phi/\partial y$ measures the effect of nonwage income in changing the sample of observations, i.e. the sample who work from the population. Suppose that those with greater nonwage income have tastes for work that are less inclined against work than those with little nonwage income (after controlling for the other determinants of work behavior). Then, as y is increased, so the composition of the sample is altered towards those with less aversion to work. Consequently, $\partial\phi/\partial y > 0$, $\hat{\alpha}_2 > \alpha_2$, and the estimated effect of nonwage income on hours of work will be biased in such a way as to indicate a less negative income effect than is really the case.

The sample selection bias can be addressed in a number of different ways. Perhaps the most common procedure is Heckman's (1976b) two-step estimator which replaces $\phi(\cdot)$, the conditional mean of ε_i, in eq. (31) with its value predicted from a previously-estimated equation. Although our understanding of the issues has been greatly enhanced by the large literature that has arisen on the subject of sample selection bias, I know of no evidence from empirical studies of male labor supply (whether old, young, or prime-age men) that documents grievous biases from a strategy of restricting estimation to the sample of workers and of not making any correction for this deliberate nonrandom selection of the observations.[61]

The following section presents the empirical results from fitting static labor supply functions. It is impossible for me to graph each fitted hours of work equation as a function of the observed values taken by the variables of interest. Yet this is exactly what is needed for a full understanding of the implications of any given set of estimates. Unfortunately, only rarely are such graphs presented. The normal substitute is to present the implied values of the behavior responses calculated at sample mean values or, less frequently, the average of the behavioral responses calculated for each observation.[62] Some papers do not even do this nor do they provide sufficient information for such calculations to be made by an interested reader. It is high time the editors and referees of all journals required that every empirical paper considered for publication present descriptive statistics on their samples analyzed.

[61] The paper by Wales and Woodland (1980) provides a convenient list of alternative methods. Also they report some sampling experiments with different estimators.

[62] These two methods of summarizing the behavior responses – either calculating the behavioral responses at the mean values of the variables or calculating the implied responses for each observation and then forming the average – may yield quite different values depending upon the form of the function and the distribution of the values of the variables. Although the latter may well be a preferable procedure, it is well nigh impossible to simulate all the studies to perform the calculations required.

The summary estimates I shall concentrate on are those measuring, first, the effect of a proportional increase in wage rates on the proportional change in hours worked and, second, the effect of a small increase in nonwage income on hours worked and, given wages, on earnings. The former is, of course, the uncompensated elasticity of hours of work with respect to wages (E) and the latter I call the marginal propensity to earn (*mpe*) out of nonwage income. Following eq. (12), the income-compensated elasticity of hours of work with respect to wages (E^*) is simply the difference between E and the *mpe*:

$$E = \frac{\partial h}{\partial w}\frac{w}{h}; \qquad mpe = w\frac{\partial h}{\partial y} \quad \text{and} \quad E^* = E - mpe.$$

Being independent of the units in which the budget constraint variables are measured, estimates of elasticities are more conveniently compared across different studies than are changes in hours worked over a given period of time (a year maybe or a week) per dollar or pound change in the wage rate. From the value of the *mpe* may be inferred how much of an increase in nonwage income is spent on the consumption of commodities. The consumption literature provides information on the marginal propensity to consume out of nonlabor income,[63] but this research focuses upon the division of an additional dollar of nonlabor income between consumption and saving holding labor income fixed, an issue involving intertemporal considerations. By contrast, the static model of time and consumption outlined in Section 3.1 takes such savings decisions as being determined at a prior stage of the individual's allocation problem and the question that arises from this model is the within-period division of an additional dollar of nonlabor income between the consumption of commodities and of leisure. Most of the estimates of this *mpe* come from the labor supply research to be surveyed shortly, but some educated guesses about the probable magnitude of this can be formed from measured effects of nonwage income on commodity consumption. Such estimates have been presented by Deaton (1982) using data on 1617 households from the British Family Expenditure Survey of 1973. In straightforward least-squares linear regressions that impose little prior structure on the data, he relates household expenditures on nine different categories of consumer goods to the husband's wage rate, nonwage income,[64] the number of children, the number of workers in the family, and a home ownership dummy variable. Nonwage income exerts a positive effect on the consumption of each category of goods and the sum of these marginal propensities to consume $(\sum p_i\, \partial x_i/\partial y)$ is about unity implying

[63] See, for instance, Holbrook and Stafford (1971).

[64] The husband's wage rate is defined as the ratio of "normal" weekly earnings to "normal" hours worked per week and then adjusted for income taxes. Nonwage income is, in fact, the net income of the household minus the husband's earnings.

a zero value for the *mpe*.[65] I know of no comparable study with U.S. data, but insofar as one may generalize from these results then a value of the *mpe* not far from zero is to be expected.

When comparing estimates of these behavioral responses from different research, it should be remembered that the points of evaluation differ across studies and, moreover, that for any given study these behavioral responses themselves vary from observation to observation. The manner in which these behavioral responses differ across observations is determined once the functional form for the estimating equation has been chosen. For example, when a linear hours of work equation is estimated both E and the *mpe* will necessarily be greater for individuals with relatively high wages. There is no strong prior reason to believe either that this should be true or that it should not be. Therefore, in specifying hours of work estimating equations, some economists feel more comfortable working with utility functions familiar from the research on consumer behavior. In research on labor supply, most of the (direct) utility functions posited have been additive in commodity consumption and in each individual's hours of work. The additivity assumption will necessarily bring with it restrictions on the relationship between E and the *mpe* and, in particular, analogous to Deaton's (1974) reasoning, additivity of the direct utility function can be shown to imply

$$E = (mpe) + \omega^{-1}\mu^{-1}[1+(mpe)], \tag{32}$$

where $\mu = (wh)/y$ and $\omega = (\partial\lambda/\partial y)(y/\lambda) < 0$ is the elasticity of the marginal utility of nonwage income with respect to nonwage income.[66] In other words for someone for whom nonwage income is a very small fraction of total income (i.e. for someone whose value of μ^{-1} is very small), additivity of the direct utility function will restrict the estimated value of his *mpe* to be similar to his estimated value of the uncompensated elasticity of hours of work with respect to wages (E) and for this individual the compensated elasticity, $E^* = \omega^{-1}\mu^{-1}(mpe)[1+(mpe)]$, will tend to be a small number. Of course in some data nonwage income appears for a number of people not to be such a small part of total income so for such individuals E will not approximate the *mpe*, but nevertheless eq. (32) shows that

[65] In fact, the estimates of the *mpe* after imposing more structure on the data are similar to these least squares regressions. See Atkinson and Stern (1980) and Deaton (1982).

[66] More generally, direct additivity of the household utility function $U = \phi[f_0(x) - f_1(h_1) - f_2(h_2)]$ implies the following relationships for the elasticity of hours of work of individual 1:

$$E_{1j} = \mu_1^{-1}(mpe)_1\left[\mu_j + \omega^{-1}(mpe)_j\right] + \delta_{1j}\omega^{-1}\mu_1^{-1}(mpe)_1, \quad j = 0,1,2,$$

where $\delta_{1j} = 1$ if $j = 1$ and $\delta_{1j} = 0$ otherwise, where $(mpe)_0 = p\,\partial x/\partial y$, and where E_{10} must be interpreted as the negative of the uncompensated elasticity of hours of work with respect to commodity prices. Note that, because a part of income is endogenous, ω here is different from the usual concept of Frisch's money flexibility.

additivity builds in restrictions among the behavioral responses that the data are unlikely to conform to.

It is useful as a reference for our discussion below to illustrate eq. (32) with a utility function (or a variant of it) that has been used relatively often in labor supply analysis. Abstracting from variations in personal characteristics A and in individual tastes ε, consider the following additive (strongly separable) utility function described by the parameters b, c, B, and ρ:

$$U(x,h)=\left[(1-B)(x-c)^{\rho}+B(b-h)^{\rho}\right]^{1/\rho}, \tag{33}$$

where $0<B<1$, $x>c$, $b>h$, and $\rho<1$. This utility function goes by different names – sometimes the nonhomothetic constant-elasticity-of-substitution function, sometimes the one-branch utility tree – but I shall refer to it as the generalized Stone–Geary utility function [Pollak (1971)]. This function conveniently nests some special cases that have frequently been used in fitting labor supply functions.[67] The optimizing hours of work function from eq. (33) is

$$h=b-\frac{B^{\zeta}w^{-\zeta}(y+bw-cp)}{\left[(1-B)^{\zeta}p^{1-\zeta}+B^{\zeta}w^{1-\zeta}\right]}, \tag{34}$$

where $\zeta=(1-\rho)^{-1}>0$ and the *mpe* and the uncompensated elasticity of hours of work with respect to wages (E) are as follows:

$$\begin{aligned} mpe &= -\frac{B^{\zeta}w^{1-\zeta}}{(1-B)^{\zeta}p^{1-\zeta}+B^{\zeta}w^{1-\zeta}} \\ E &= (1-bh^{-1})\left\{\frac{[\zeta b+(1-\zeta)h]w}{(y+bw-cp)}-1\right\} \\ &= -1+[1+(mpe)]\left[\zeta bh^{-1}+(1-\zeta)\right]. \end{aligned} \tag{35}$$

The behavioral responses corresponding to the Stone–Geary utility function are obtained by letting ζ equal unity, whereas the conventional constant-elasticity-of-substitution function is obtained in eq. (33) by setting the "reference" parameters, c and b, to zero and replacing the term $B(b-h)^{\rho}$ with $-B^{*}h^{\rho}$. Chipman's (1965) "weakly homothetic" utility function results when $\rho\to-\infty$. With utility function (33), $\omega=-\zeta^{-1}y(y+bw-cp)^{-1}$, so that with the definitions of the *mpe* and E above, eq. (32) is easily derived.

[67] Within the class of empirical work making use of nonexperimental data on individual workers, eq. (33) covers the functional forms used by Betancourt (1971), Blundell and Walker (1981, 1983), Brown, Levin, Rosa, Ruffell and Ulph (1982–83), Hurd and Pencavel (1981), Rosen (1978), Wales (1973), and Wales and Woodland (1979). In addition, the hours of work equation derived from eq. (29) is similar to that estimated by Atkinson and Stern (1980, 1981).

In short, whether derived explicitly from a particular utility function or simply written down ab initio, the hours of work estimating equation involves selecting a specific functional form and the choice of this function inevitably embodies some assumptions about the differences in the behavioral responses (i.e. the differences in E and the *mpe*) across individuals. Unfortunately, at present an assessment of these assumptions is difficult because so little is known about these variations.

In most cases, the static model has been estimated by fitting a regression equation such as eq. (28) to cross-section data collected from a sample survey of households or of individuals. The precise questions asked vary from survey to survey, but normally an individual (or his spouse) is asked about his hours worked (and his weeks worked) in a given week (year) or in a typical week (year), his labor earnings during a specified period of time or his usual hourly earnings, and his income from other sources. The response to these questions form the basis of the observations on the purported labor supply function.

In an econometric exercise associating quantities (hours of work) and prices (wage rates), prior to estimation it is appropriate to enquire whether what is being estimated is a supply function, a demand function, or some hybrid. Suppose that a worker with a specific set of characteristics valued by firms faced a horizontal demand curve for his services, i.e. the worker may choose any hours to work at a given wage rate. Workers with different characteristics of varying values to firms would face horizontal demand curves at different levels of real wages. Provided some of these characteristics were not at the same time associated with these workers' preferences for income or leisure,[68] then in a cross-section of individuals the revealed wage–hours combinations would reflect the intersection of different horizontal demand curves with a fixed (for a given set of variables determining preferences) labor supply function. This provides one rationalization of the common presumption that a regression of hours worked on wage rates and other variables maps out a labor supply function.

As noted in Section 3.3 above, most firms appear not to be indifferent to the hours worked by each of their employees: the presence of quasi-fixed hiring and training costs that are more closely tied to the *number* of employees than to their hours worked encourages firms to offer higher wage rates for longer hours worked [Lewis (1969)]. If this is the case, the worker faces a wage–hours locus such that shorter hours of work are renumerated at a lower hourly wage rate. Once again, across workers with the same preferences, their labor supply function is traced out by a series of different (nonhorizontal) labor demand schedules, each demand curve indexed by a particular quality of labor. Provided identifying variables exist, the labor supply function can be estimated by a regression of hours worked

[68] What are the (identifying) variables that appear in the demand function for hours by employers and that do not enter the supply function for hours of work? Perhaps the most obvious candidates for such variables are indicators of the level of local labor market activity.

on wage rates, but now of course account must be taken of the fact that the wage offered by employers is no longer independent of each worker's own decisions.

4.2. *Empirical results from U.S. nonexperimental data*

A brief chronology of the major phases of modern empirical research on male labor supply may be listed as follows. Kosters' (1966, 1969) analysis of the hours worked of married men aged between 50 and 64 years old ranks as the first modern empirical study of this topic both by virtue of its close attention to its theoretical underpinnings and by virtue of his use of a sample of observations on individuals;[69] there soon followed many studies [a number of them being brought together in Cain and Watts (1973)] whose methods were similar to Kosters', but which analyzed other groups in the labor force; in response to the diversity of results from these studies and in an attempt to account for them, the next phase of research [as best illustrated by DaVanzo, DeTray and Greenberg (1973, 1976)] was the application of a variety of different procedures to a single body of data; the 1970s also saw increasing attention to the econometric implications of nonrandom sample selection [Heckman (1974b, 1976b)] and nonlinear budget constraints [Burtless and Hausman (1978), Wales and Woodland (1979)]; meanwhile, from the mid-1970s, new sources of information were becoming available, namely the results from the various negative income tax experiments and the estimates from British research; finally, the 1970s witnessed increasing attention to the life-cycle models of labor supply and, at the time of writing, this seems to be the most active area of male labor supply research.

In order to trace this chronology a little more closely, return to Kosters' original analysis of the hours worked by employed married men aged 50–64 years. His observations were drawn from the 1 in 1000 sample of the 1960 Census of Population and he estimated to these data ordinary least-squares equations linear in the logarithms of the variables. One such equation is the following which was estimated with 8467 observations:

$$\ln h_i = -\underset{(0.0044)}{0.094} \ln w_i - \underset{(0.0015)}{0.0073} \ln y_i + \cdots + \hat{\varepsilon}_i, \quad R^2 = 0.10,$$

where estimated standard errors are in parentheses beneath coefficients and where the dots indicate that 16 other variables were included in the regression equation. The income-compensated wage elasticity of hours of work (E^*)

[69]A number of studies preceded Kosters' that examined the issues at an aggregate level – Douglas (1934) had measured the association between hours worked and earnings at the industry level, Finegan (1962) at the occupational level, Winston (1966) at the national level – but Kosters appears to have been the first to apply the theory to the unit whose behavior it is meant to describe.

implied by the estimates is +0.041 when evaluated at the (geometric) mean values of the observations. The estimate of −0.094 for the uncompensated wage elasticity was robust with respect to changes in equation specification and, moreover, accorded well with previous estimates – with Douglas's (1934) preferred estimate "in all probability somewhere between −0.1 and −0.2" and with Winston's (1966) estimates of −0.07 to −0.10, though less so with Finegan's (1962) estimates of −0.25 to −0.35. On the other hand, the estimate of −0.0073 for the nonwage income elasticity of the supply of working hours appeared to be sensitive to changes in functional form and in the precise definition of nonwage income.

Kosters' procedures with relatively minor modifications were soon being applied by other researchers to different samples. A stimulus to this research was provided by the prominent public policy debate over the costs of welfare reform which were intimately tied to the labor supply effects of taxes and transfers. In part as a consequence of this emphasis on welfare reform, a number of studies that reported in early 1970s restricted their empirical work to samples of the relatively poor. In constructing such samples, observations were discarded on the basis of values taken by a variable (income) that is clearly related to the endogenous variable of interest (hours of work). This induces an analogous sort of sample selection bias as that discussed in Section 4.1 above.[70]

This feature of male labor supply studies of the early 1970s – that observations on relatively high income individuals or households were eliminated from their samples – represented only one dimension in which the various research papers differed from one another. They also differed in the precise definitions of the variables, the particular functional relationship posited, the assumptions made about commodity prices, and the set of nonbudget constraint variables included in the hours of work regression equations. These differences in the implementation of the labor supply model yielded sufficiently disparate estimates as to provide little practical assistance to questions of public policy. In view of these differences, it was important to address the question: "With respect to which set of assumptions and procedures are the hours of work estimates sensitive and with respect to which are they robust?" This was taken up by DaVanzo, DeTray and Greenberg (1973) who applied many different procedures to a single body of data, namely, 5294 white, married, male heads of households aged 25–54 years drawn from the 1967 Survey of Economic Opportunity (SEO). Their Rand report is full of valuable information for anyone embarking on his own labor supply

[70] For an elaboration of this point in the labor supply context, see Cain and Watts' (1973) lucid statement. For a more general treatment of the issue, see Goldberger (1981). Studies that imposed some sort of income criterion in defining their analysis sample included those of Boskin (1973), Fleischer, Parsons, and Porter (1973), Greenberg and Kosters (1973), Hall (1973), Hill (1973), Kalachek and Raines (1970), Kurz et al. (1974), and Rosen and Welch (1971).

study.[71] The same question was addressed by Masters and Garfinkel (1977) in their extensive analysis of data from the 1967 SEO and from the 1972 Michigan Panel Study of Income Dynamics (PSID). The differences in procedures among the studies and the consequences of these different procedures may be summarized as follows.

1. Problems in measuring the hours and wage rate variables. In studies based on data from the 1960 Census of Population or the 1967 Survey of Economic Opportunity, the hours of work variable combined one dimension of work behavior (namely, hours per week) in one year (in 1960 for the Census and in 1967 for the SEO) with another dimension of work (namely, weeks worked per year) in a different year (in 1959 for the Census and in 1966 for the SEO).[72] Then this dependent variable often appeared in the construction of the wage rate variable (i.e. for the Census data, annual labor income in 1959 was divided by this estimate of hours worked) so that any errors in measuring true hours worked in 1959 or in 1966 will appear in the wage rate variable inducing a spurious negative correlation between hours worked and wage rates. What contribution, if any, was this making to the frequent finding of a negatively-sloped labor supply curve? The answer, it seemed, was that the slope of the male ordinary least-squares estimated hours of work function was more negative when such a wage variable was used than when an alternative wage rate variable (such as an instrumented wage rate) was constructed. Evidence on this is contained in Bloch (1973), DaVanzo, DeTray and Greenberg (1973), Masters and Garfinkel (1977), and Borjas (1980). Nevertheless, even after trying to rid the wage variable of this spurious correlation, most studies found a negative (uncompensated) own-wage elasticity of hours of work at sample mean values: for instance, DaVanzo, DeTray and Greenberg (1973) report estimates between -0.15 and -0.09,[73] Masters and Garfinkel (1977) "best estimate" is -0.110, and Ashenfelter and Heckman's (1973) is -0.156.

[71] Much of their analysis was conducted with a sample of 2012 men who reported being unaffected by unemployment and by poor health and who received no work-related transfer payments. They then considered the consequences of adding to the original sample 3282 men who reported these characteristics.

[72] Many other definitions of the hours worked variable have been used. A common one is the product of the number of weeks worked in a given year and the average number of hours worked per week during those weeks in which the individual worked. Some studies add an estimate of hours spent unemployed to the number of hours worked.

[73] DaVanzo, DeTray and Greenberg's estimates reported here are derived from Tables 11 and 12 of their Rand study where the dependent variable is measured as annual hours of work and where the wage rate and nonwage income variables are instrumented. The sample in this case consists of those 2012 men who reported no unemployment or health disability nor receipt of any work-related transfer payments. Other variables included in these equations are age, age squared, schooling, household size, number of children less than six years of age, various variables denoting location of residence, the spouse's annual earnings, and the annual earnings of other family members.

2. The measurement of nonwage income. This variable was particularly difficult to measure accurately. Koster's procedure was to form this variable by deducting the husband's earnings from total household income, but this meant y included transfer income that was not independent of the husband's hours of work. Also, y excluded income in the form of the service flow from durable goods and housing. Moreover, this definition of nonwage income incorporated the earnings of the wife and of other members of the household and, therefore, it is not exogenous with respect to the husband's labor supply behavior if the work decisions of each member of the household are made jointly.[74] In other studies [e.g. Ashenfelter and Heckman (1973)], y is explicitly measured by aggregating the responses to the survey's questions about the net income received in the form of rents, dividends, interest, private transfers, and alimony payments. Another procedure [e.g. Fleisher, Parsons and Porter (1973)] is to assume that y is proportional to the household's net worth (where the factor of proportionality is given by the relevant rate of return). These different procedures generate markedly different estimates of the effect of nonwage income on hours of work. For instance, the *mpe* (i.e. $w \cdot \partial h / \partial y$) at sample mean values is estimated at -0.27 in Ashenfelter and Heckman (1973), -0.06 in Bloch (1973), -0.08 in Fleisher, Parsons, and Porter (1973), approximately -0.32 in Kalachek and Raines (1970), and -0.047 in Masters and Garfinkel (1977). However, these estimates are sensitive to the particular specification of the estimating equation and, indeed, it is by no means uncommon for a positive (partial) association to exist between nonwage income and hours of work. For instance, of the 57 different estimated coefficients on net worth reported in Tables 6, 9, 11, and 12 of DaVanzo, DeTray and Greenberg's Rand study, only 16 would be judged as significantly different from zero on conventional two-tailed t-tests and, of these 16, exactly one-half is positive and one-half is negative. Positive (partial) correlations between male hours worked and nonwage income are reported in Cohen, Rea and Lerman (1970), Dickinson (1974), Garfinkel (1973), Hill (1973), Kniesner (1976), and Masters and Garfinkel (1977) and they would probably have been calculated in Burtless and Hausman (1978), Hausman (1981), and Hurd and Pencavel (1981) if the estimation procedure had not prohibited it. In view of these widely varying estimates on nonwage income, when an equation such as eq. (24) is fitted and the substitution effect is calculated residually as $\alpha_1 - h\alpha_2$, given the negative (un-

[74] This raises another class of differences among the various empirical studies, namely, the treatment of the wife's labor earnings. Sometimes her earnings are incorporated into nonwage income in which case the tacit assumption is that these earnings produce an income effect on the husband's hours of work, but no substitution effect. On other occasions, the wife's wage rate is included as a separate independent variable, but often its estimated coefficient is insignificantly different from zero by conventional criteria. This was DaVanzo, DeTray and Greenberg's finding and, moreover, their estimates for the coefficient on the husband's wage rate were affected only trivially by different ways of specifying the wife's earnings.

compensated) effect of wages on hours of work that is typically estimated (i.e. given $\alpha_1 < 0$), it is by no means unusual for the implied substitution effects for male workers to be negative at the sample mean values of h. Such negative effects appear in the empirical work of, for instance, Cohen, Rea and Lerman (1970), DaVanzo, DeTray and Greenberg (1973), Fleisher, Parsons and Porter (1973), Hall (1973), Kniesner (1976), Kosters (1966), and Masters and Garfinkel (1977). This hardly constitutes a resounding corroboration of the conventional static model of labor supply.

3. The treatment of taxes. Sometimes, as in Kosters' study and in Ashenfelter and Heckman's (1973) study, no allowance was made for personal income taxes either in forming the wage rate or the nonwage income variable. On other occasions, as in Boskin (1973) and in Hall (1973), the budget constraint was assumed to be continuous and to form a convex set and budget constraint variables net of taxes were constructed, but then the joint determination of all these budget constraint variables with hours of work was ignored. There have been few instances [one is Kurz et al. (1974)][75] in which the budget constraint variables were adjusted for taxes and, in addition, they were treated as endogenous. In order to assess the effects of adjusting the budget constraint variables for taxes, we should like to see from the same body of data estimates of hours of work equations based on pre-tax budget constraint variables and instrumental variable estimates based on post-tax budget constraint variables. I know of no study that presents this information for men though Mroz (1984) has undertaken such a comparison for married women and found relatively small differences between the two sets of estimates.

Is the assumption that the after-tax budget constraints for most men are continuous and form a convex set an important departure from the truth? Some think so. Therefore, they have proposed and applied more elaborate algorithms that are designed to search over each segment of a piecewise-linear budget constraint in order to determine the parameters describing the utility-maximizing hours of work. For instance, Wales and Woodland (1979) assume they know without error each individual's net wage rate and nonwage income and they use these budget constraint variables together with the unknown parameters of the individual's constant-elasticity-of-substitution utility function (posited to be the same and nonstochastic for all individuals) to impute each individual's hours of work along each segment of his piecewise-linear budget constraint. For each individual, therefore, there is a relationship between the different possible values of the utility function's parameters and his imputed hours of work, given the values of his budget constraint. Among many possible values of the parameters

[75] However, the procedures of Kurz et al., do not yield a consistent estimator because nonlinear transformations of the imputed wage rate and nonwage income variables were used in the hours of work equations.

of the utility function, those are selected that minimize the sum over all individuals of the squared difference between the imputed hours and the actual hours. The only sources of error in their model are errors in maximization or the effects of random variables (examples of which, write Wales and Woodland, are unanticipated expenditures or illness) that cause the individual to work different hours from those given by his budget constraint variables and utility function. They applied their algorithm to a sample (from the Michigan PSID) of 226 married men whose wives did not work in the labor market and their estimates of the utility function parameters implied values of the (uncompensated) wage elasticity of hours of work of 0.14 and of the marginal propensity to earn of -0.70. This wage elasticity lies above the central tendency of estimates while the marginal propensity to earn is an even more noticeable outlier and one might be inclined to wonder whether the more conventional estimation methods have seriously misestimated these behavioral parameters. However, Wales and Woodland derived similar estimates when they applied the more conventional approach of linearizing the budget constraint around the observed hours of work for each man so that the more elaborate algorithm did not appear to be responsible for the estimates of the relatively high wage elasticity and aberrant marginal propensity to earn.

Other studies using these sorts of algorithms have also yielded odd estimates. For instance, Hausman's (1981) work is a generalization of Wales and Woodland's to allow for stochastic variation in preferences across individuals, but otherwise he proceeds on similar lines.[76] With a sample of 1085 married men from the 1975 Michigan PSID, Hausman has the benefit of almost five times as many observations as Wales and Woodland.[77] Fitting a linear hours of work function, Hausman estimated an (uncompensated) wage elasticity of male working hours of zero and a marginal propensity to earn of approximately -0.77.[78] Although this latter estimate is not without precedent, it differs sharply from the implications of estimates of nonwage income on consumption. Hausman's estimate implies that an additional dollar of nonwage income induces such a reduction in working hours that (at sample means) labor earnings fall by 77 cents and the consumption of commodities increases by only 23 cents. Income effects in consumption could be this small, but the prevailing evidence suggests the contrary.

[76]An excellent exposition of Hausman's work [and that of Burtless and Hausman (1978)] is contained in Heckman and MaCurdy (1981) and Heckman, Killingsworth and MaCurdy (1981).

[77]This increase in the size of the sample is not achieved costlessly, however. Whereas Wales and Woodland examined only those men whose wives did not work in the labor market, Hausman made no distinction between men whose wives were working and those who were not.

[78]This value is derived as follows. Hausman reports a mean gross wage rate of \$6.18 and predicted mean hours of 2181. This implies labor income of \$13 479. Suppose someone with this income faces a marginal tax rate of 25 percent. Then the mean net wage rate is approximately \$4.64 (= \$6.18 × 0.75). Given his estimate of $\partial h/\partial y$ of -0.166, the *mpe* for such an individual is -0.77.

In short, these studies, using more elaborate computational algorithms, yield estimates of the key behavioral parameters that diverge from the central tendency of estimates and that are somewhat implausible. Because these studies pay greater attention to some issues (especially the piecewise-linear nature of the budget constraint and perhaps also its nonconvexity) at the cost of the neglect of others (e.g. they treat wage rates and nonwage income as exogenous and not measured with error), it is by no means evident that their estimates of the male labor supply function are to be regarded as preferable to those derived from more prosaic and perhaps more robust estimating methods.[79]

4. Assumptions about commodity prices. In most cross-section studies it was assumed that all individuals face the same prices for commodities so that variations in the money wage rate and money nonlabor income correspond to variations in the real values of the variables. There were a few studies [e.g. Bloch (1973), Boskin (1973)] that made use of some Bureau of Labor Statistics information on the cost of living in different regions and cities. If such geographic cost-of-living adjustments are not made, then this rationalizes the presence of region and city size dummy variables that often appear in estimated labor supply equations. When this BLS information on cost-of-living differences by city size and by region was used to deflate the wage rate variable, both DaVanzo, DeTray and Greenberg (1973) and Masters and Garfinkel (1977) report small changes in the estimated coefficient in the wage rate.

5. Issues of functional form. Kosters' linear-in-the-logarithms specification reported above is unusual in this literature. More frequently, as discussed in Section 4.1 linear equations along the lines of eq. (28) have been estimated. Occasionally the following semi-logarithmic specification in wage rates has been posited:

$$h_i = \alpha_0 + \alpha_1 \ln\left(\frac{w}{p}\right)_i + \alpha_2\left(\frac{y}{p}\right)_i + \alpha_3 A_i + \varepsilon_i,$$

which restricts the uncompensated wage effect to be smaller (in absolute value) for high wage individuals. There is, of course, no a priori reason to believe that the data will naturally conform to the restrictions on the behavioral parameters implied by these functions. In view of the prominent role occupied in introductory texts by the so-called backward-bending labor supply curve, it was natural for researchers to determine the empirical relevance of such a phenomenon. Normally this has been effected by adding quadratic terms in the wage rate to

[79] This conjecture about the robustness – that methods such as Hausman's and Wales and Woodland's are less robust with respect to small departures from the assumptions that underlie them as compared with the more conventional estimation methods – is also contained in Heckman (1983).

equations such as eq. (28) [e.g. Bloch (1973), DaVanzo, DeTray and Greenberg (1973), Hill (1973), Rosen and Welch (1971)] or by estimating a free form whereby the efficient α_1 is allowed to vary across different wage intervals [e.g. Cohen, Rea and Lerman (1970), DaVanzo, DeTray and Greenberg (1973), Garfinkel (1973), Greenberg and Kosters (1973), Hall (1973)]. There have been instances in which evidence for such a backward-bending hours of work function for males has been reported [e.g. Cohen, Rea and Lerman (1970)], but forward-bending curves have also been estimated [e.g. Hurd (1976), Kurz et al. (1974)], and from an overview of the empirical results, there does not appear to be powerful evidence for nonlinearities in the wage-hours relationship for men. However, most of this research on functional form has been incidental to other issues and a systematic empirical investigation of the variation of income and substitution effects across individuals has yet to be undertaken in labor supply research.[80]

6. Nonbudget constraint variables included in the hours of work equation. The various studies on male labor supply differ from each other in the set of control variables entered in the hours of work regression equation. For instance, some studies include a measure of the individual's educational attainment [e.g. Cohen, Rea and Lerman (1970), Garfinkel (1973), Hill (1973), Kniesner (1976), Kosters (1966), Rosen and Welch (1971)] while other studies exclude it [e.g. Ashenfelter and Heckman (1973), Bloch (1973), Boskin (1973), Hausman (1981), Hurd (1976), Masters and Garfinkel (1977)]. When such a variable is included, its estimated coefficient is almost always positive and significant by conventional criteria suggesting that, other things equal, more formally educated men work longer hours. Moreover, DaVanzo, DeTray, and Greenberg's investigation found that the size and sign of the wage coefficient was extremely sensitive to the presence of years of schooling in the estimated hours of work equation.[81] As another example, a measure of the number of dependents in the household is sometimes included in an equation accounting for variations in the working hours of men [e.g. Bloch (1973), Boskin (1973), Cohen, Rea and Lerman (1970), Hausman (1981), Masters and Garfinkel (1977)] and it is sometimes excluded [e.g. Ashenfelter and Heckman (1973), Fleisher, Parsons and Porter (1973), Garfinkel (1973) Rosen and Welch (1971)]. When a variable of this kind is included, it tends to reveal a significantly positive (partial) association with hours of work. In general, researchers have been somewhat cavalier in their choice of nonbudget constraint variables to be included in an hours of work equation, but unfortunately DaVanzo, DeTray, and Greenberg's experiment with their school-

[80]A start is contained in Dickinson (1979, 1980).

[81]Some researchers may well be seduced into omitting schooling from the hours of work regression equation because then they may claim it as an instrument for wage rates.

ing variable indicates that the presence or absence of certain nonbudget constraint variables may profoundly affect the inferences about the wage elasticity of hours of work. It is not unusual for no explicit reason to be given for the presence in the hours of work regression equation of these nonbudget constraint variables. Most researchers seem to have in mind that variables such as education or family size are systematically associated with differences in tastes for work (or, equivalently, differences in nonmarket productivity) so that they correspond to what I have denoted as the variables A in the description of the contrained maximization problem above. Nevertheless, as I have emphasized in Section 2, in addition to these taste variations that are believed to be associated with variables (such as education and family size) observed to the researcher, there is also a very important unobservable taste component (as represented by ε in Section 2). Usually this unobserved taste component is simply tacked on as the stochastic term to the hours of work equation, but there exist other ways of addressing the issue of variation in observed tastes. For instances, Greenberg and Kosters (1973) constructed a variable designed to represent differences in preferences for asset accumulation by measuring the difference between an individual's actual net assets and those net assets predicted on the basis of his age and wage rate from a prior regression equation and then expressing this difference as a fraction of total imputed wealth. This inclusion of this so-called preference variable changed their estimated coefficient on nonwage income in an hours of work regression equation from positive to negative. The problem with this variable, as Cain and Watts (1973) note, is that its construction makes use of information about the wage rate and nonwage income and thus it is natural to wonder whether it incorporates some part of the conventional wage and income effects of the budget constraint.

A number of the estimates from U.S. nonexperimental data of the static model's behavioral responses are brought together in Table 1.19 Although the major studies are included, this table is not exhaustive. In several cases [such as Wales and Woodland (1976, 1977)] insufficient information is provided in the publications with which to calculate the compensated wage-elasticities or the *mpe*. In other cases [e.g. Hall (1973)] many different estimates are presented and I gave up the attempt to summarize them adequately with a few numbers. I have also excluded studies such as those of Hausman (1981) and Hurd and Pencavel (1978) that in estimation restricted the effect of nonwage income on hours to be nonpositive. In drawing inferences from Table 1.19, the caveats given in Section 4.1 above should be kept in mind. These estimates are drawn from different estimating equations and from different functional forms and evaluating the estimated parameters at sample mean values of the variable provides only a very rough and inexact method of comparing behavioral responses. Table 1.19 reveals that, of the estimates presented, Wales and Woodland's (1979) are considerably different from the rest, a result I attribute both to the restriction between E and

Table 1.19
Estimates from U.S. nonexperimental data of behavioral responses for men.

	E	mpe	E^*
Ashenfelter and Heckman (1973)	−0.16	−0.27	0.12
Bloch (1973)	0.06	−0.06	0.12
Boskin (1973)	−0.29	−0.41	0.12
DaVanzo, DeTray and Greenberg (1973)	−0.15	−0.004	−0.14
Dickinson (1974)	−0.11	0.08	−0.19
Fleisher, Parsons and Porter (1973)	−0.19	−0.23	0.04
Garfinkel (1973)	0	0	0
Greenberg and Kosters (1973)	−0.09	−0.29	0.20
Ham (1982)	−0.16	−0.11	−0.05
Hausman and Ruud (1984)	−0.08	−0.63	0.55
Kniesner (1976a)	−0.17	−0.01	−0.16
Kosters (1966)	−0.09	−0.14	0.04
Masters and Garfinkel (1977)	−0.11	−0.05	−0.06
Wales and Woodland (1979)	0.14	−0.70	0.84

Notes: The estimates reported for DaVanzo, DeTray and Greenberg (1973) correspond to those given on the last line of Table 11 of their Rand report where both the wage rate and nonwage income variables were instrumented. Those for Ham (1982) correspond to those given in column (1) of Table IV of his paper. Those for Kniesner (1976a) apply to those men whose wives were not at work for pay. For Masters and Garfinkel (1977), I took what they described as their "best estimates" of E and the *mpe* even though the coefficients reported did not derive from the same regression equation. Boskin's (1973) results are those for white men only. Dickinson's (1974) *mpe* is calculated from his estimate coefficient on "other (nontransfer) family income". Hausman and Ruud's estimates are calculated for a household with an assumed marginal tax rate of 25 percent so the husband's net wage rate is \$4.31 and the wife's net wage rate is \$2.63.

the *mpe* implicit in their use of the CES function[82] and to their estimating method which may well not be robust with respect to small departures from the assumptions underlying its use. Of the remaining studies, the largest estimate of E is 0.06 [Bloch (1973)] and the smallest is -0.29 [Boskin (1973)]. The central tendency of estimates of E lies between -0.17 and -0.08 and a simple average of all the estimates of E in Table 1.19 (excluding Wales and Woodland's) is -0.12. Table 1.19's estimates of the *mpe* (again excluding Wales and Woodland's) range from a low of -0.63 [Hausman and Ruud (1984)] to a high of 0.08 [Dickinson (1974)]. The estimates of the *mpe* are more disparate than those for E and I hesitate to infer its value from such a varied set of estimates. Certainly, the large negative numbers seem very unlikely. In five cases in Table 1.19, the compensated wage elasticity of hours of work, E^*, is negative. Of the six positive

[82]As equation (35) makes clear, in the CES case when $b=0$, $E=-\zeta+(1-\zeta)(mpe)$.

values of E^* (excluding Wales and Woodland's and Hausman and Ruud's), the mean is 0.11. If E is -0.12 and E^* is 0.11, the *mpe* is -0.23.

4.3. *Empirical results from British data*

Modern British research on male hours of work got under way in the 1970s and from the beginning the work has consistently been concerned with the implications of the taxation of income on the supply of labor and so the studies invariably adjust each individual's budget constraint variables for such taxes.[83] The first papers were those of Brown, Levin and Ulph (1976) and Layard (1978). The data analyzed in the former study came from a survey conducted at the end of 1971 by a private market research firm. With a relatively small and perhaps unrepresentative sample[84] of 284 married men whose wives were not at work in the labor market, Brown, Levin and Ulph (1976) estimated (with a conventional ordinary least-squares regression linear in parameters but nonlinear in the budget constraint variables) an (uncompensated) own-wage elasticity of hours of work of between -0.085 and -0.131 at sample mean values.[85] This was derived from a curious specification in which both linearized nonwage income and a measure of "other income" were included.[86] Subsequent work by Brown (1981) and his associates using similar procedures yielded comparable wage elasticities and marginal propensities to earn of between -0.31 and -0.35. Other methods were also applied to these data including a study by Ashworth and Ulph (1981) that, independently of the work of Wales and Woodland (1979) and Burtless and Hausman (1978), proposed and implemented the procedure of searching over each individual's entire piecewise linear budget constraint to determine the utility-maximizing hours of work. With a generalized constant-elasticity-of-substitution indirect utility function applied to 335 married men, Ashworth and Ulph (1981) derived estimates that implied an uncompensated wage elasticity of hours of work of between -0.07 and -0.13 and a marginal propensity to earn of between -0.36 and -0.57.

Layard's (1978) study involved a much larger sample of 2700 married men from the General Household Survey of 1974 and, with a linear specification along

[83] The British studies of male workers always use weekly hours of work as the dependent variable.

[84] The authors themselves were aware of both the small size and possible nonrandom nature of their sample. A very informative discussion of these data is contained in Brown (1981).

[85] Although this wage elasticity is estimated to be negative at sample mean values, it becomes less negative as the wage rate rises and, indeed, it eventually takes on positive values. In other words, they estimate an hours of work function that is a mirror-image of the textbook backward-bending function.

[86] In a later study [Brown (1980, p. 60)], this is justified on the argument that "other income" is, in fact, dependent upon the male's labor supply. Of course, if this is the case, then it should be included in calculating the wage slope of the budget constraint.

the lines of eq. (24), he estimated an uncompensated wage elasticity of -0.13 and a small (in absolute value) marginal propensity to earn of -0.04. Indeed, with such an income effect, his implied compensated wage effect on hours of work was negative.

In Britain the availability of cross-section information from the Family Expenditure Survey (FES) on both hours of work and expenditures on different groups of commodities has permitted the joint estimation of labor supply and commodity demand equations as implied by eq. (8). Provided the allocation model underlying eq. (8) is correct, estimating such a system of equations has the advantage of generating much more efficient estimates. The greatest potential for these data is to *test* that allocation model, but curiously they have not been used for this purpose to date. Nevertheless, some indications of how these tests would fare are provided in the papers making use of these data. Consider, for instance, the work of Atkinson and Stern (1980, 1981) who specified a generalized Stone–Geary utility function where that generalization is the novel one involving explicit use of Becker's (1965) particular formulation of the household production approach to the allocation of time. In fact, when all the commodities may be aggregated into one composite, their hours of work function closely resembles eq. (30). They select a sample from the 1973 FES consisting of 1617 households with a male head employed full-time (not self-employed) and whose earnings placed him within the (fairly wide) range in which the slope of the after-tax budget constraint was approximately constant. They identify nine different categories of household consumption expenditures plus the hours of work of the men.[87] Their results suggested uncompensated wage elasticities (evaluated at their sample pre-tax mean values) ranging from -0.15 to -0.23 although, as in Brown, Levin and Ulph (1976), the estimated hours of work function is a forward-falling curve and at relatively high wages they estimate a positive wage elasticity. They tend to find that leisure is an inferior commodity and ultimately they impose the constraint that pure leisure is not valued for its own sake (i.e. it has value only insofar as it contributes to the production of utility-generating activities). As is often the case with the Stone–Geary specification, the extent of nonconvexity implied by the estimates is considerable and, in particular, many men work more hours than permitted by the estimates of the maximum amount remaining after allocating time to other activities.

Another study estimating a system of commodity demand and labor supply equations is Blundell and Walker's (1982). From the 1974 FES, they select a sample of only 103 households in which both the husband and the wife work, a term being included to account for this deliberate nonrandom selection of female

[87] The earnings of the wife and of others in the household are included in nonwage income.

workers. They also specified a generalized Stone–Geary utility function[88] in which six groups of commodities and the hours of work of the husband and of the wife appear as arguments. Their parameter estimates implied (at sample mean values) an uncompensated wage elasticity of male hours of work of -0.23 and a marginal propensity to earn of -0.36 with, therefore, an implied (compensated) wage elasticity for men of 0.13. Because the wife's marginal propensity to earn was estimated to be -0.22, their results implied that an additional dollar of nonwage income would raise consumption by only 42 cents (i.e. $1-0.36-0.22$). Blundell and Walker do not indicate how many of their 103 husbands and wives are working more hours than permitted by the estimated parameters describing the maximum feasible hours of work, but there are surely some although probably a smaller proportion of their sample than of Atkinson and Stern's. They test and reject the hypothesis that the husband and wife's time allocation decision is weakly separable from the household's decisions about the consumption of commodities, but maintained throughout the analysis is the hypothesis that expenditures on housing are separable from all other decisions. In a subsequent study of 308 working married couples drawn from the 1977 FES and specifying four categories of consumer goods (but excluding alcohol, tobacco, housing, and other durable goods (expenditures), Blundell and Walker (1983) report an uncompensated wage elasticity of male hours of work of -0.004 (evaluated at 39.6 weekly hours of work, the mean value for their earlier sample) and an *mpe* of -0.203.

The preliminary results from another British project financed by H. M. Treasury are becoming available at the time of writing this survey paper [Brown, Levin, Rosa, Ruffell and Ulph (1983)]. This new project involved both a new survey (conducted in late 1980) and a new sample of 3307 households who provided sufficient information for analysis. In an initial investigation of 810 one- and two-worker households, the researchers applied a similar algorithm to that used by Ashworth and Ulph (1981) and Wales and Woodland (1979) to search over each individual's entire piecewise linear budget constraint. Unfortunately, this algorithm did not identify a well-defined maximum of the likelihood function although the estimates of the parameters of the nonstochastic generalized Stone–Geary function [identical to eq. (33)] are described as being "in the right area". At the sample mean values of the wage rate and nonwage income, the worker in single-worker families is estimated to have an uncompensated wage elasticity of hours of work of -0.32, a compensated wage elasticity of 0.18, and a marginal propensity to earn of -0.50. In two-worker families, the husband

[88] The generalization takes the form of specifying the "subsistence" or "reference" quantities not as parameters, but as functions of commodity prices and of household structure. In fact, because all households are assumed to face the same prices for commodities, the only effective generalization is one which allows the subsistence quantities to vary across households with different numbers and ages of children.

Table 1.20
Estimates of the behavioral responses for British males.

		E	mpe	E^*
Ashworth and Ulph (1981)		−0.13	−0.36	0.23
Atkinson and Stern (1980)		−0.16	−0.07	−0.09
Blundell and Walker (1982)		−0.23	−0.36	0.13
Blundell and Walker (1983)		−0.004	−0.20	0.20
Brown, Levin, and Ulph (1976)		−0.13	−0.35	0.22
Brown et al. (1982–83)	Single worker	−0.33	−0.50	0.17
	Two workers	−0.14	−0.44	0.30
Layard (1978)		−0.13	−0.04	−0.09

Notes: The estimates for Brown, Levin and Ulph (1976) are those where the wife does not work for pay. The estimates for Brown et al. (1982–83) are those for a family with two children.

possesses an uncompensated wage elasticity of between −0.14 and −0.06 (the former estimate for husbands with two children and the latter for husbands with no children), a compensated wage elasticity of between 0.30 and 0.39, and a marginal propensity to earn to between −0.45 and −0.42. In these two-worker families, the wife's marginal propensity to earn is estimated at approximately −0.15 so that together these estimates imply a family's marginal propensity to earn of about −0.60 or, expressed differently, only 40 percent of a small increase in exogenous nonwage income is spent on the consumption of commodities.

A summary of these British estimates is contained in Table 1.20. All the estimates of the uncompensated wage elasticity of hours of work are negative and a simple average of the eight estimates is −0.16. Five of the eight estimates are between −0.16 and −0.13. As was the case with the studies with U.S. males, the variations in the *mpe* and in E^* among the studies is considerably greater than the variation in E. Of the six positive estimates of E^*, the average is 0.21.

4.4. *Empirical results from U.S. experimental data*

The fundamental implication of the allocation model outlined in Section 3.1 is that, for a population of individuals at a given time or for a given individual over time, other things equal, exogenous movements in budget constraints should induce movements in the supply of labor. This most basic proposition stood an excellent opportunity of being tested by the various negative income tax (NIT) experiments that were conducted in the United States in the decade from 1968 to 1978. With the laboratory sciences as a conscious example, these experiments selected a sample of households in a given locality and then introduced to a fraction of this sample (the experimental households) a different budget constraint while continuing to observe the other households (the controls). The consequences of changes in the budget constraint for the supply of labor could be inferred by contrasting the behavior of the experimental households with that of the control households during the experiment and/or by contrasting the behavior of the experimental families during the experiment with their behavior before (or after) the experiment.[89]

In fact, inferences from the experiments were much more difficult to draw. There were several reasons for this. First, the sample of (experimental and control) households studied was drawn selectively from the low-income population. This was a natural decision in view of the concern with welfare reform, but its effect was to introduce problems deriving from the truncation of a variable (income) directly related to the major variable of interest (labor supply). Second, this low-income sample of households was then not allocated randomly between the experimental and the control groups, but rather the allocation design was a more complicated one that partly tried to mitigate the budgetary costs of the experiment. Third, during each experiment, changes took place outside the experiment's control that affected the budget constraints of the participating households and that may have affected the control and experimental households differentially. For instance, in the middle of New Jersey's experiment, the state's welfare program was reformed in such a way that, for a number of experimental households, it now offered a more generous opportunity than the experiment's and so these households opted out of the experiment. As another example, the first NIT payments in Seattle were made (in November 1970) at a time when the area was experiencing a drastic and unprecedented rise in unemployment arising from the extensive layoffs in its aircraft industry and it was feared that an idiosyncratic labor market situation existed from which it was hazardous to extend inferences about the effects of a negative income tax to more typical labor market settings. Fourth, even if the sample of experimental households and the sample of control households had been the same at the outset, greater attrition of controls subsequently from the experiment rendered the two samples different from one another.[90] Fifth, as in all welfare and tax programs, incentives existed for individuals to misreport their incomes so that statutory and actual tax rates diverged. Indeed, it has been conjectured thai the particular incentives created by the NIT experiments operated to exaggerate the magnitude of true labor supply

[89] This is the method prescribed in Orcutt and Orcutt's (1968) classic statement of the case for social experiments.

[90] Of course, the problem of attrition exists in all panel data, not merely in the NIT experimental data. A frequently-cited paper on the subject of attrition is that by Hausman and Wise (1977) who claimed that in the Gary experiment attrition bias was less with a "structural" model of earnings than with an analysis-of-variance (AOV) model. However, this inference was drawn from a comparison between, on the one hand, a "structural" model that included almost all the determinants of attrition in the earnings equation and, on the other hand, an AOV model that excluded many of the determinants of attrition from the earnings equation. The implied constraint in the AOV model was clearly not warranted and their comparison was thereby quite invalid.

effects.[91] Sixth, because most of the experimental households were eligible to receive NIT payments for three years,[92] it has been argued that the labor supply effects should be interpreted as those induced by *temporary* changes in net wage rates and nonwage income.[93] Seventh, because only a relatively small fraction of an area's population had their budget constraints altered by the experiments and because these changes were temporary, the inducements to make institutional adjustments in work schedules were considerably less than would be the case for a national and permanent NIT program. For instance, approximately two-thirds of the husbands in the Gary experiment worked in the steel mills on work schedules that permitted them little flexibility in working hours in their existing jobs. There would have been greater pressures on the employers and the unions to renegotiate different hours of work schedules if it had not been the case that only a relatively small fraction of all employees in these steel mills were enrolled in the experiment and if the experiment had lasted for more than three years. In this sense, the experimental–control differences would tend to understate the adjustments that would occur if the budget constraint changes were not confined to a relatively small population over a relatively short space of time. All these issues certainly impede drawing straightforward inferences from the experimental data although, given the size of the differences in NIT payments between experimental and control households in some of the experiments, it is unlikely that these problems entirely nullify simple experimental–control comparisons.

The NIT experiments were conducted on 1357 households in New Jersey and Pennsylvania from 1968 to 1972, 809 households in rural areas of North Carolina and Iowa from 1969 to 1973, 1800 households in Gary, Indiana, from 1970 to 1974, and 4800 households in Seattle and Denver from 1970 to 1980. Not only were many more households analyzed in the Seattle–Denver experiment compared with the others, but also it involved more generous NIT payments. For the typical male, in each case the experimental treatment meant changing his budget constraint from $0a_1a_2$ to $0b_1b_2a_2$ in Figure 1.4.[94] In other words the NIT experiment paid a grant (or support) of G dollars regardless of the household's income and then applied a relatively higher tax rate τ on all income in excess of G. The breakeven level of income, b_2 in Figure 1.4, occurred when the household's receipts in the form of the grant, G, equalled tax payments, $\tau(wh + y)$. For any individual located to the right of b_2 both before and after the introduction of the

[91] See Ashenfelter (1978), Greenberg, Moffit and Friedman (1981), and Welch (1978).

[92] Some households in Seattle and Denver experiment were eligible to receive payments for five years.

[93] The original work investigating this issue is Metcalf's (1973, 1974).

[94] Figure 1.4 assumes a pre-experimental budget constraint characterized by a continuously rising marginal tax rate. For some households (especially single heads of households), the non-experimental welfare programs (such as AFDC) generate budget constraints similar to the experimental budget constraint $0b_1b_2a_2$.

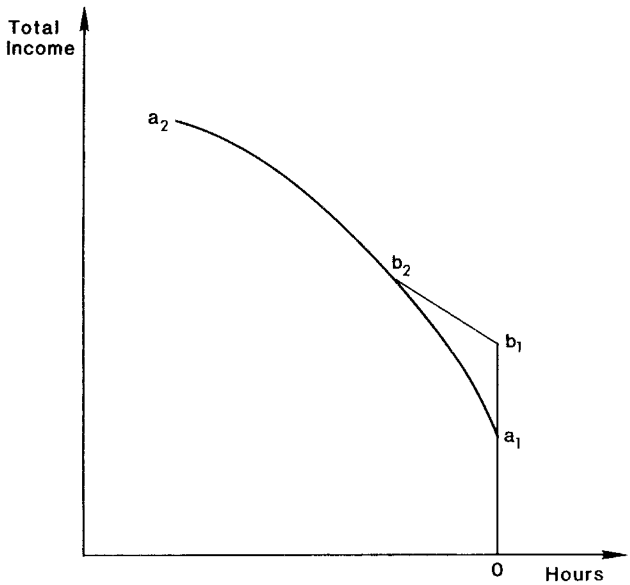

Figure 1.4

NIT experiment, his opportunities were enhanced and his pre-experimental budget constraint $px = wh + y$ became $px = G + (1 - \tau)(wh + y)$, where τ is the differentially higher tax rate applied on income by the NIT experiment. There were also some individuals who were located on their pre-experimental budget constraints to the left of b_2, but who determined upon the introduction of the experiment they would be better off by so reducing their hours of work as to locate on $b_1 b_2$ and become eligible for NIT payments. Other things equal, the flatter an individual's indifference curve (i.e. the greater an individual's elasticity of substitution between income and leisure), the greater the probability of his moving from above the breakeven level of income pre-experimentally to below the breakeven level of income during the experiment.[95] The values of G and τ differed across and within the four experiments[96] and once again the assignment of experimental households among the different NIT programs (each described

[95] Expressed differently, consider an experimental individual who is indifferent between a point on his budget constraint to the left of b_2 and a point to the right of b_2. For this individual, the experimentally-induced change in the budget constraint involves no income effect, only a substitution effect. This is essentially Ashenfelter's (1983) insight that the substitution effect can be measured from estimating the relationship between the fraction of individuals below the breakeven level of income and the slope of the arm $b_1 b_2$.

[96] In the Seattle–Denver experiment, there were some "treatments" in which τ itself was not a constant, but instead fell as income rose.

by a particular combination of G and τ) was not random. In particular, there was a tendency for households with relatively low pre-experimental incomes to be assigned to the less generous NIT programs (i.e. those with relatively low G and high τ) thereby reducing the expected budgetary cost of the experiment.[97] This implies that the particular experimental parameters, G and τ, applied to each household depended upon its pre-experimental earnings and these in turn were not independent of its experimental labor supply insofar as there is correlation over time in a household's work behavior. In short, contrary to some claims, the experimental treatments were not genuinely exogenous both because each household decided whether to received NIT payments by being below the breakeven level of income b_2 and because the particular program parameters it faced were not assigned to it randomly.[98]

To determine whether the data collected by the NIT experiments conform to the basic notion that differences in work behavior are associated with differences in budget constraints, the following ordinary least-squares regression equation was estimated:

$$L_i = \beta_0 X_i + \beta_1 E_i + u_i, \tag{36}$$

where L_i stands for a dimension of individual i's work behavior[99] (such as his weekly hours of work or whether or not he was employed in the labor market), E_i takes the value of unity for an individual allocated to the experimental sample and of zero for an individual in the control sample, X_i measures other characteristics of the individual (and, in the Seattle–Denver research, X_i also includes the variables determining the assignment of individuals to different treatments), and u_i is a stochastic disturbance term. Sometimes eq. (36) was estimated with data

[97]An excellent analysis of the implications of the assignment process is contained in Keeley and Robins (1980) who advise including the variables determining the assignment of households to different NIT programs in equations designed to infer the labor supply effects of the experiments.

[98]This fact vitiates many of the original arguments in support of undertaking such social experiments. These arguments claimed that conventional income and substitution effects would be much easier to measure with experimental data because the experiment induced *exogenous* changes in the budget constraints of experimental households. As noted earlier, because of the nonrandom assignment of households between the control sample and the experimental sample and because of nonrandom assignment within the experimental sample of households to different treatments, the changes in the budget constraint were not truly exogenous to the households. Moreover, the nonlinearity of the budget constraints creates a further reason for the budget constraint variables to be endogenous. What appears to be a more convincing argument in defense of the experiments is that the within sample variations in the budget constraint variables (and especially in nonwage income) tend to be larger than in nonexperimental data and this holds out the hope of measuring the parameters associated with these budget constraint variables more precisely.

[99]Instead of work behavior, a few studies [such as Ashenfelter's (1978)] focus upon the experimental effect on net earnings. There is good reason for this in view of the fact that the NIT-induced change in earnings is proportional to the excess transfer cost of the program over the cost calculated on the basis of pre-experimental incomes alone.

drawn from midway during the experiment in which case X_i usually included some measure of an individual's pre-experimental labor supply. On other occasions, data were pooled from the experimental and the pre-experimental period and experimental–control differences during the experiment were distinguished from such differences before the experiment. The estimates of β_1 in eq. (36) were consistently (though not invariably) negative: for white husbands in the New Jersey–Pennsylvania experiment, the experimental group averaged 5.6 percent fewer hours of work per week than the control group [Rees (1974)]; for black husbands in Gary, the experimental group averaged 6.5 percent fewer hours of work per month than the control group [Moffitt (1979)]; and in the Seattle–Denver experiment, husbands in the experimental group worked 2.2 percent fewer hours per week than those in the control group [Keeley et al. (1978a)]. The differences among the NIT experiments in the point estimates of β_1 were less marked than the differences in their estimated standard errors: the experimental–control differences measured in the New Jersey–Pennsylvania, North Carolina–Iowa, and Gary experiments were often insignificantly different from zero by conventional criteria while those in the Seattle–Denver experiment were clearly significantly different from zero, a consequence of the substantially greater size of the Seattle–Denver experiment. Experimental husbands were also less likely than controls to be employed at any moment midway through the experiment – a 2.6 percent difference for white husbands in New Jersey–Pennsylvania [Rees (1974)], a 4.9 percent differential in Gary [Moffitt (1979)], and a 2.3 percent differential in Seattle–Denver [Pencavel (1982)]. These estimated experimental–control differences tend to understate the magnitude of the experimental labor supply response because the experimental–control dummy E_i in eq. (36) measures the effect of the experiment averaged over those experimental families who receive NIT payments by being below the breakeven level of income and those whose incomes place them above the breakeven level. In other words, β_1 in eq. (36) understates the experimental effects conditional upon being below the breakeven level of income.

The results reported in the previous paragraph were designed to answer the question of whether changes in budget constraints result in changes in work behavior. The evidence suggests that, beyond any pre-experimental differences, the changes introduced by the NIT experiments did induce differences between the experimental and control husbands' work behavior. Of course, the allocation model of Section 3.1 has implications beyond the simple one of maintaining that changes in budget constraints cause changes in work behavior; in any changes in budget constraints, this model distinguishes the effects of changes in wage rates from those attributable to changes in nonwage income. It is natural to determine, therefore, whether the experimentally-induced changes in net wage rates and in net nonwage income each generated effects on work behavior that are compatible with the neoclassical static allocation model. Moreover, distinguishing the effects on work behavior of the NIT tax rate from the effects of the guarantee level is essential if the purpose is to draw inferences from these experiments about how other welfare programs (with different program parameters) would operate.

There have been many different forms of specifying the net wage and nonwage income effects on work behavior induced by the NIT experiments although, as others have observed [e.g. Ashenfelter (1978)], many of the models used by analysts of the experimental data (especially those in New Jersey–Pennsylvania and in North Carolina–Iowa) were specified in ways that make it difficult to recover income and substitution effects from them. In those studies where these behavioral responses (or their transformations) are identified, most of the important differences among the studies have turned on the way in which each household's budget constraint has been measured. As Figure 1.4 makes clear, for both experimental and control households, the nonlinearity of the budget constraint renders the net wage rate endogenous to the labor supply decision. Hausman and Wise (1976) addressed this problem by measuring the budget constraint variables for each individual at the same number of working hours (namely, 1500 hours per year), but this has the effect of simply assigning the wrong budget constraint to all individuals except those who happen to be working 1500 hours. A procedure not so different from this is applied by Keeley et al. (1978a, 1978b) who take up Ashenfelter and Heckman's (1973) proposal of measuring the budget constraint variables in the second year of the Seattle–Denver experiment as those that would obtain at each individual's pre-experimental hours of work. Of course, these measures can only be correct if, in fact, each individual did not change his work behavior as a consequence of the NIT experiment or for any other reason. Johnson and Pencavel (1982) also measure the change in the budget constraint along these lines, but then they treat these variables so constructed as measured with error and apply an instrumental variable estimator. Moffitt (1979) measures the tax rate by averaging each individual's marginal tax rate over the entire length of his budget constraint. Johnson and Pencavel (1984) and MaCurdy (1983) linearize each individual's budget constraint around his observed hours of work during the Seattle–Denver experiment and then treat these budget constraint variables as endogenous by replacing them with their values predicted from a prior regression. Burtless and Hausman (1978) use a generalization of Wales and Woodland's (1979) procedure described in Section 4.3 where the generalization takes the form of permitting each individual's utility function to contain a component that is unobserved to the researcher and that varies (according to a specified distribution) across the population. Wales and Woodland's (1979) method of determining the unknown parameters of the hours of work function by searching over all segments of the piecewise linear budget constraint must now be specified such that each individual's location on a particular segment (given his net wage rate and net nonwage income) is known only probabilistically. As in Wales and Woodland's study,

Table 1.21
Estimates of the behavioral responses for men from the NIT experiments.

		E	mpe	E^*
Ashenfelter (1978a)		0.21	0.02	0.19
Ashenfelter (1978b)		0.17	−0.01	0.18
Burtless and Greenberg (1982)	3 Year	0.08	−0.04	0.12
	5 Year	−0.12	−0.18	0.06
Hausman and Wise (1977)		0.10	−0.01	0.11
Johnson and Pencavel (1982)		−0.16	−0.29	0.13
Johnson and Pencavel (1984)		0.02	−0.17	0.19
Keeley and Robins (1980)		−0.09	−0.14	0.05

Notes: Ashenfelter's estimates are from the North Carolina–Iowa rural experiment and Hausman and Wise's are from the New Jersey–Pennsylvania experiment. All the other estimates make use of data from the Seattle–Denver income maintenance experiment and all these estimates have been evaluated at the same number of hours of work (namely, 1880.97) and the same net wage rate ($2.293). These are the mean values of working experimental husbands in the pre-experimental year whose incomes in that year would have placed them below the breakeven level and they are taken from the sample analyzed by Keeley and Robins (1980). The earlier work by Keeley, Robins, Spiegelman and West (1978a, 1978b) uses the same estimating procedure as in Keeley and Robins (1980), but in the later study the sample includes Chicanos, unlike the earlier work. The difference between Ashenfelter's (a) and (b) estimates is explained in footnote 100.

Burtless and Hausman assume they measure each individual's budget constraint without error.

Table 1.21 summarizes the estimates of the *mpe* and the wage elasticities of hours of work from a number of the analyses of the hours of work of husbands in the NIT experiments. This table does not list every study that claims to be measuring these behavioral responses, but only those studies that satisfy two conditions: first, they provide sufficient structure on the estimated relationships that the results have some claim to correspond to the behavioral responses; and, second, they impose sufficiently few prior estimating restrictions as to supply an opportunity for the data to reveal whether they really conform to the implications of the static allocation model. This second condition implies that I have omitted, for example, Horner's (1977) paper with the New Jersey–Pennsylvania experimental data that measures the parameters of a Cobb–Douglas utility function and Burtless and Hausman's (1978) paper on the Gary experimental data that constrains no individual to have a positive *mpe*. The first condition means that I have excluded studies such as Watts' (1974) and Moffitt's (1979) that involved specifications in which the income and wage effects on hours of work were supposed to be gleaned from the estimated coefficients on the experimental tax rate and the guarantee level in an hours of work equation and where other variables also incorporating measures of wages and nonwage income were included in the regression.

Of the studies listed in Table 1.21, Hausman and Wise's (1977) makes use of the New Jersey–Pennsylvania experimental data, Ashenfelter's (1978) makes use of the North Carolina–Iowa experimental data,[100] and the rest make use of the Seattle–Denver experimental data. All the summary estimates in Table 1.21 relating to the Seattle–Denver experiment have been calculated at the same values of working hours and wage rates, as the notes to the table make clear. The point estimates of the uncompensated wage elasticity, E, range from a low of -0.159 [Johnson and Pencavel (1982)] to a high of $+0.015$ [Ashenfelter (1978)]. The point estimates of the *mpe* range from a low of -0.290 [Johnson and Pencavel (1982)] to a high of $+0.015$ [Ashenfelter (1978)]. The estimates of E^*, the compensated wage elasticity, in the different studies range from a low of 0.050 [Keeley and Robins (1980)] to a high of 0.192 [Ashenfelter (1978)]. This relatively narrow range of estimates of E^* comes about through offsetting values of E and the *mpe*, the range of estimates of E and *mpe* being considerably greater. The tendency is for the uncompensated hours of work function to be positively sloped with respect to wage rates at sample mean values. By estimating

[100] The two sets of estimates for Ashenfelter's analysis of the data from the rural experiment correspond to two different parameterizations of the experimentally-induced change in earnings. He posits an hours of work function for family member 1 as $h_1 = h_1(w_1, w_2, y)$, where w_2 is the net wage rate of family member 2. The experimentally-induced change in this person's earnings is

$$w_1\,\mathrm{d}h_1 = w_1\frac{\partial h_1}{\partial w_1}\mathrm{d}w_1 + w_1\frac{\partial h_1}{\partial w_2}\mathrm{d}w_2 + w_1\frac{\partial h_1}{\partial y}\mathrm{d}y$$
$$= -\left[w_1\left(w_1\frac{\partial h_1}{\partial w_1} + w_2\frac{\partial h_1}{\partial w_2}\right)\right]\tau + \left[w_1\frac{\partial h_1}{\partial y}\right](G - \tau y),$$

where $\mathrm{d}w_1$ (the change in wage rates induced by the experiment) is given by $-\tau w_1$, $\mathrm{d}w_2 = -\tau w_2$, and $\mathrm{d}y$ (the change in nonwage income induced by the experiment) is $G - \tau y$. So what is designated in Table 1.21 as scheme (a) regresses the change in earnings on the tax rate τ and on $G - \tau y$. Observe that the coefficient on τ incorporates any cross-wage effects. The estimates under E for Ashenfelter (a) in Table 1.21 sets these cross wage effects to zero. Ashenfelter's second parameterization makes use of the Slutsky decomposition to write the previous equation

$$w_1\,\mathrm{d}h_1 = -w_1\left[w_1\left(\frac{\partial h_1}{\partial w_1}\right)^* + w_2\left(\frac{\partial h_1}{\partial w_2}\right)^*\right]\tau + w_1\frac{\partial h_1}{\partial y}[G - \tau(w_1h_1 + w_2h_2 + y)],$$

where the term $G - \tau(w_1h_1 + w_2h_2 + y)$ corresponds to the NIT payments and where the asterisk denotes compensated wage effects. Here the change in earnings is regressed on the tax rate τ and on NIT payments. Again the coefficient on τ reflects a cross-wage effect and again the estimates under E for Ashenfelter (b) in Table 1.21 sets these cross-wage effects to zero. This second parameterization is similar to that used by Keeley et al. (1978a, 1978b). They set cross-wage effects to zero and divide the last equation in this footnote by w_1.

the labor supply parameters separately for households on the three year experimental program from those on the five year program, Burtless and Greenberg (1982) derive values for E^* and the *mpe* that diverge in the manner that Metcalf (1973, 1974) conjectured: the compensated wage-elasticity is larger and the *mpe* is smaller (in absolute value) for the three year experimental husbands compared with the five year experimental husbands.

It is important to point out that the responses whose point estimates are presented in Table 1.21 are not normally estimated with much precision. For example, Hausman and Wise's point estimate of E of 0.095 comes with an estimated standard error of 0.043 so that a 95 percent confidence interval ranges from 0.001 to almost 0.180. Or the largest of the point estimates of E in Table 1.21, Ashenfelter's 0.207, has an estimated standard error of 0.122 so that a 95 percent confidence interval spans a range from -0.032 to $+0.446$. It is difficult to draw the inference from estimates such as these that the NIT experiments have permitted the relevant behavioral responses to have been measured with much precision.

4.5. *Conclusions*

If the estimates from Tables 1.19, 1.20, and 1.21 are put together, it appears that the estimates of E, the uncompensated wage elasticity of hours of work, from the American nonexperimental data tend to be more negative than those from the data collected in the NIT experiments. This difference between the estimates from the experimental and those from the nonexperimental data conforms to Metcalf's (1973, 1974) conjecture: the temporary nature of the NIT experiments will tend to cause the estimate of the *mpe* to be smaller (in absolute value) and the estimate of the compensated wage elasticity, E^*, to be larger when estimated from experimental data than their "permanent" values. If this is the case, then indeed we should expect the estimates of E to be larger when fitted to experimental then to nonexperimental data. British men appear to be similar to American men in their value for E, although this E decomposes into a more negative *mpe* and a larger E^* for the British. If a single number has to be attached to each of the behavioral responses, then for American prime-age men the (uncompensated) wage elasticity of hours of work is -0.10 and their *mpe* is -0.20.

The inferences in the previous paragraph are drawn from a comparison of the central tendency of the point estimates in Tables 1.19, 1.20 and 1.21. It would be misleading to present these summaries without at the same time emphasizing both the diversity of estimates and the imprecision with which these point estimates are measured. Moreover, if the estimates are interpreted as tests of the static model of labor supply (and no doubt some would not want to take this step), then the frequency of negative values for the income-compensated wage elasticity of hours of work casts serious doubt on its empirical relevance.

5. Estimation of the life-cycle model

The discussion of the Life-Cycle Models in Section 3.5 concentrated on those assuming strong separability of the lifetime utility function and the presentation of empirical work in this section restricts itself to this class of models. Also, as in the discussion of the empirical work on static models, I omit discussion of the estimation of life-cycle labor supply behavior at the macroeconomic level as in the work of Lucas and Rapping (1969) and others. The reason is in part because of major aggregation problems: such work normally seeks to explain movements in aggregate manhours worked and confuses individuals occupying a corner solution to their allocation problem with those at interior solutions. Indeed, the larger part of the movement in aggregate manhours over the business cycle is attributable to movements in the numbers of workers employed and not to movements in the hours worked of those continuous employed.[101] Because the microeconomic evidence reported below is restricted to individuals at interior solutions to their constrained optimization problem, it is not straightforward to go from these estimates to draw implications about corresponding parameters estimated with macroeconomic data.[102] It is not surprising, therefore, that as Altonji (1982) has shown the estimates of the macro parameters are by no means robust with respect to small changes in the assumptions underlying their calculation.

In the microeconomic research described in this section, it should be remembered that, although the life-cycle model has important refutable implications (for instance, the Frisch demand and supply functions possess symmetry, homogeneity, and sign properties), there has been virtually no work testing the

[101] For evidence on this, see Coleman (1983). By contrast, Hall (1980, p. 95) claims: "Both recessions of the 1970's saw pronounced reductions in average hours of work." As Coleman shows, Hall's inferences are in error. His index of aggregate hours is calculated using both the hours per worker and the number of workers series from the BLS establishment surveys. His series on total employment is from the household Current Population Survey. The ratio of aggregate hours from the establishment survey to numbers employed from the CPS yields a variable hours per worker series, but it does not correspond to anything observed in the U.S. economy. When Coleman uses either the ratio of hours to employment both from the establishment surveys or the ratio of hours to employment both from the CPS, the hours per worker series displays little annual variability. In other words, most of the cyclical variability in aggregate manhours is attributable to changes in the number of workers employed and not to changes in hours worked per employee.

[102] Often these macroeconomic models are described as if economic agents operate under uncertainty. As MaCurdy (1982) shows, this further aggravates the problems of identifying from aggregated data the effect on labor supply of parametric wage changes (which is what Lucas and Rapping maintain they are measuring).

empirical relevance of these implications. The life-cycle model has been characterized as the maintained hypothesis and empirical work has taken the form of gauging the parameters describing the presumed life-cycle allocation. Of course, the measurement of the parameters of well-specified models is a necessary ingredient of any science, but such information is not the same as that derived from offering the model good opportunities of being refuted and discovering it has survived such tests of its validity.

As we shall see, these life-cycle models are most convincingly estimated when the research makes use of successive observations over time of the same individuals (i.e. panel data). Because an important component of this work involves regressing changes (over time) in the hours worked of individuals on corresponding changes in their wage rates, it might be noted that the simple correlation between these two variables is negative at least in the U.S. data. For instance, Abowd and Card (1983) report that when changes in the logarithm of hours worked are regressed on changes in the logarithm of wages rates (controlling for no other variables) the estimated coefficient is -0.36 for 1531 prime-age male heads of households in ten years of the Michigan panel and it is -0.28 for 1321 men aged less than 65 years in 1975 in six survey years of the National Longitudinal Survey of Older Men.[103] However, in view of the problems documented in Section 4.2 in measuring hours and wages accurately, there is every reason to wonder how much of this negative correlation between the observed values of the variables is attributable to measurement errors and how much to an association between the true values of the variables. After all, often the wage rate variable is formed by dividing the respondent's annual earnings by hours worked so any error in measuring hours will produce a spurious negative correlation between hours and wage rates and this negative correlation will normally persist when taking first-differences in the variables.

In addition, both the measured hours and the measured wage rate variables do not precisely correspond to their counterparts in the economic model. That is, with respect to wage rates, there are all the problems described in Section 3.3 concerning nonlinear budget constraints (taxes, nonlinear compensation schedules, etc.) while hours of work are often computed as the product of two variables (average hours worked per week and weeks worked per year) and therefore are unlikely to correspond exactly to the true value of the variable. Also, according to one influential model, labor supply should exclude time spent in on-the-job training yet the hours reported rarely deduct such human capital investment. These problems concerning measurement error in wage rates and hours worked may well be exacerbated by first-differencing the variables because permanent components in these variables are thereby eliminated and "noise" components account for a relatively larger part of the measured total. Therefore, if only for

[103] Similar inferences can be drawn from data in Altonji's (1983) paper.

Table 1.22
Estimates of the intertemporal substitution elasticity, γ, for prime-age men.

Research	Equation estimated	Type of data	Other variables in equation	Instrumental variables used	$\hat{\gamma}$
MaCurdy (1981a)	Eq. (38) first-differenced	Panel data on individuals	None	Father's education, mother's education, parents' socio-economic status, schooling, age, interactions, year dummy variables	From 0.14 to 0.35 (0.07) (0.16)
MaCurdy (1981a)	As above	As above	Year dummy variables	As above	From 0.10 to 0.45 (0.13) (0.29)
Becker (1975)	Eq. (41)	Synthetic cohorts	Age, log of nonwage income, log of other family income, log of family size	Group means	0.45 for white men (0.11) 0.10 for nonwhite men (0.04)
Becker (1975)	Eq. (41) first-differenced	Synthetic cohorts	Changes in variables listed above	Group means	0.17 for white men (0.11) −0.06 for nonwhite men (0.08)
Smith (1977)	Eq. (41)	Synthetic cohorts	Age, log of wife's wage, number of young children	Group means	0.32 for white men (0.05) 0.23 for black men (0.11)
Browning, Deaton, and Irish (1983)	Eq. (42)	Synthetic cohort over time	Number of children, long-term interest rate, manual–nonmanual dummy, cohort and year dummies	Group means	About 0.05
Browning, Deaton, and Irish (1983)	Eq. (42) first-differenced	As above	Changes in above variables plus year dummies	Group means used plus age, age squared, lagged wages and prices	About 0.03
MaCurdy (1982)	Eq. (38) replacing λ_0 with age and age-invariant variables	Individual cross-section	Father's education, mother's education, parents' socio-economic status, age, own education	Father's education mother's education, parents' socio-economic status, age, own education, $(\text{age})^2$, $(\text{education})^2$	From −0.07 to 0.28 (0.23) (0.47)

Notes: Estimated standard errors are given in parentheses beneath coefficients. No standard errors are available for Browning, Deaton and Irish's estimates. MaCurdy's work uses the Michigan PSID. Becker makes use of the 1 in 1000 sample from the 1960 Census of Population. Smith uses the 1967 Survey of Economic Opportunity. Browning, Deaton and Irish use the British Family Expenditure Surveys.

purely statistical reasons, it would seem essential in this work to address explicitly the problems of measurement error in wages and hours. In fact, these reasons are compounded by economic considerations arising out of behavior under uncertainty. Consequently, the research surveyed below is restricted to that work taking explicit account of measurement error in these variables.[104] Table 1.22 contains a summary of estimates of the intertemporal substitution elasticity and other features of the research.

The archetypal study of male life-cycle labor supply was MaCurdy's (1981). He specified individual i's utility function at age t to be the addilog:

$$U_i(x_{it}, h_{it}; A_{it}, \varepsilon_{it}) = b_{it}(x_{it})^{\gamma_1} - \tilde{\gamma}^{-1}c_{it}(h_{it})^{\tilde{\gamma}}, \tag{37}$$

where

$$\tilde{\gamma} = \gamma^{-1}+1, \quad \gamma > 0, \quad 0 < \gamma_1 < 1, \quad b_{it} > 0, \quad c_{it} = \exp\left[\gamma^{-1}(-\beta A_{it} - \varepsilon_{it})\right].$$

The objective function is thus not merely additive over time, it is also additive in consumption and hours within any given period. The Frisch hours of work equation for individual i at age t is

$$\ln h_{it} = \psi_i + \gamma \ln w_{it} + \beta A_{it} + \delta t + \varepsilon_{it}, \tag{38}$$

where $\psi_i = \gamma \ln \lambda_{0i}$ and $\delta = \gamma \ln \theta$. The first term on the right-hand side is invariant for a given individual over time and is different from individual to individual. The parameters of eq. (38) supply information on how an individual's hours of work differ over time in response to anticipated, evolutionary, wage changes, i.e. wage changes along a worker's wage–age profile. The proportional change in hours of work induced by a proportional increase in wage rates as a worker ages is measured in eq. (38) by $\gamma > 0$, the intertemporal substitution elasticity.[105]

To estimate eq. (38), MaCurdy used ten annual observations on 513 white, continuously-married, men from the Michigan PSID who were aged 25–46 years in 1967 and who were observed in each of the ten years from 1967 to 1976. The variables in A_{it} could be any whose values did not change over this ten-year period. The estimates of γ from first-differencing eq. (38) ranged from 0.14 to

[104] Abowd and Card (1983) do allow for measurement error in wages, but on the other hand they assume ε_{it} in eq. (16) to be zero, i.e. that the researcher knows each individual's utility function exactly. If their model is augmented to allow for unmeasured characteristics of individuals, then once these ε_{it} are permitted to be correlated for each individual over time Abowd and Card's variance components procedure no longer identifies the intertemporal substitution elasticity.

[105] Because the utility function has been assumed to be additive over time, the intertemporal substitution elasticity is equivalent to the specific substitution elasticity.

0.35 with standard errors on these coefficients of 0.07 and 0.16, respectively.[106] When yearly dummy variables were included in the first-differenced form of eq. (38), γ was estimated much less precisely although its point estimate changed little: the point estimates now ranged from 0.10 to 0.45 with standard errors of 0.125 and 0.29, respectively.[107] By including yearly dummy variables in the first-differenced equation, the coefficient on the wage rate cannot be interpreted as the response of labor supply to changes in wages induced by business cycle forces. So these point estimates implied that, as a male worker ages, a doubling of his wage rates induces a proportional increase in his hours worked of from ten percent to 45 percent.

These general inferences from the Michigan PSID have been confirmed by Altonji (1983) and by Ham (1983). The sample analyzed by Altonji is slightly different from MaCurdy's[108] and he also considers the consequences of using different sets of instrumental variables for the change in $\ln w_{it}$. The consequences for the estimated intertemporal substitution elasticity of the change in the sample are small: when Altonji uses the same variables MaCurdy used as instruments, his estimates of γ center around 0.27 with standard errors about two-thirds of this value. As an instrument for the change in $\ln w_{it}$ (where w_{it} is computed by dividing total earnings by hours worked), Altonji also uses an alternative measure of the wage variable derived by asking workers paid on an hourly basis about their hourly wage rate. Because this information is available for only a subset of workers, the use of this variable reduces his sample size by about 60 percent. The estimates of γ are now around 0.04 with estimated standard errors even larger than this. Similar results are derived when the lagged value of this alternative wage variable is used as an instrument. Ham (1983) uses eight years of data from the Michigan PSID from 1971 to 1979 (including men from the poverty subsample) to estimate a different functional form for the Frisch equation, namely, that postulated by Browning, Deaton and Irish (1983) in eq. (42) below. Evaluated near the mean values of wages and working hours, Ham's estimates of the intertemporal substitution elasticity are around 0.04.[109] In short, Altonji's and

[106] The higher of these estimates of γ come from adding $\gamma \ln h_{it}$ to both sides of eq. (38) and regressing changes in hours on changes in earnings.

[107] The coefficient on t is given by $\gamma \ln \theta \approx \gamma(\rho - r)$ so the coefficients on these yearly dummy variables (after division by γ) may be interpreted as the difference between the rate of time preference and the rate of interest. MaCurdy's estimates imply that on average r exceeds ρ by two to four percentage points.

[108] Altonji uses data from the 12 years of the panel from 1967 to 1978 on continuously married (to the same spouse) men aged 25–48 years in 1967. He includes observations even if they did not work in all 12 years, he includes nonwhites as well as whites, and he includes households from the more heavily sampled low income areas. The result is an increase in the total number of observations from 5,130 to over 8,000.

[109] Ham does not provide information on the mean wage and hours of work of the men in his sample. I have evaluated his point estimates in Table A1 of his paper at 2100 hours of work and at a wage of \$6.00 per hour. These are approximately the average values of these variables for the Michigan panel in 1975, the midpoint in Ham's longitudinal data.

Ham's research with the Michigan PSID underscore MaCurdy's findings of an intertemporal substitution elasticity whose point estimate is less than 0.45 and that is not estimated with precision.[110]

Section 3.5 maintained that the Frisch labor supply equation may also be used as the basis for empirical work when agents make decisions under uncertainty. This is an important point and it is convenient to illustrate this by making use of the particular utility function (37) above. [The argument here draws liberally on MaCurdy (1982).] In this case, the first-order condition corresponding to eq. (28) may be written

$$\ln h_{it} = \gamma \ln \lambda_{it} + \gamma \ln w_{it} + \beta A_{it} + \varepsilon_{it}. \tag{39}$$

It can be shown that $\ln \lambda_{it}$ follows a stochastic process with drift and may be represented as

$$\ln \lambda_{it} = \sum_{j=0}^{t} \tilde{a}_j + \ln \lambda_{i0} + \sum_{j=1}^{t} \tilde{v}_{ij},$$

where $\tilde{v}_{ij}$ is the individual's forecast error at age j that arises from the values of variables at age j diverging from the values expected (at age $j-1$) to obtain at age j. Substituting this expression for $\ln \lambda_{it}$ into eq. (39) and first-differencing yields:

$$\Delta \ln h_{it} = a^* + \gamma \Delta \ln w_{it} + \beta \Delta A_{it} + \varepsilon_{it} - \varepsilon_{it-1} + v_{it}, \tag{40}$$

where $a^* = \gamma \tilde{a}$ and $v_{it} = \gamma \tilde{v}_{it}$. Compared with the equation derived by first-differencing eq. (38) (i.e. the certainty case), it is evident that under uncertainty assumptions have to be made about the nature of the forecast error v_{it}. Now the marginal utility of income in period t will depend upon wages, wealth, and the individual's characteristics in period t and also upon the future path of expected wages. So suppose $\gamma \ln \lambda_{it}$ in eq. (39) may be expressed as

$$\gamma \ln \lambda_{it} = \tilde{b}_{1t} A_{it} + \sum_{j=t}^{N} \tilde{c}_{tj} \mathscr{E}_{it}(\ln w_{ij}) + \tilde{b}_{2t} K_{it} + \xi_{it},$$

where K_{it} is the real value of the consumer's wealth at the start of period t and

[110]Altonji also tried to measure the intertemporal substitution elasticity from the within-period marginal rate of substitution between hours and food consumption. Because substitution within a branch of the lifetime utility function is being estimated (no essential use is made of intertemporal data), his period-specific preferences are estimated only up to a positive monotonic transformation and thus the degree of intertemporal substitutability cannot be inferred. This same problem exists (as they fully recognize) with Blundell and Walker's (1984) research: only the sign of the intertemporal substitution elasticity may be inferred, not its numerical magnitude.

the coefficients $\tilde{b}_{1t}$, $\tilde{c}_{tj}$, and $\tilde{b}_{2t}$ change as individuals age. The revision in $\gamma \ln \lambda_{it}$ at age t is

$$v_{it} = \tilde{b}_{1t}\left[A_{it} - \mathscr{E}_{it-1}(A_{it})\right] + \sum_{j=t}^{N} \tilde{c}_{tj}\left[\mathscr{E}_{it}(\ln w_{ij}) - \mathscr{E}_{it-1}(\ln w_{ij})\right]$$
$$+ \tilde{b}_{2t}\left[K_{it} - \mathscr{E}_{it-1}(K_{it})\right] + \xi_{it} - \mathscr{E}_{it-1}(\xi_{it}),$$

where $\mathscr{E}_{it-1}$ denotes individual i's expectations at age $t-1$ of the associated variables. It is implausible to assume that the economist knows each individual's expectations perfectly and consequently this further restricts the set of variables that may serve as instruments for $\Delta \ln w_{it}$ in eq. (40). These must be variables, of course, that are uncorrelated with unanticipated changes in wage rates, wealth, and preferences and yet that are associated with $\Delta \ln w_{it}$. Appropriate instruments are lagged values of wages and prices, variables known by the individual with certainty at the time that forecasts are made.

Now let us compare MaCurdy's estimates of the intertemporal substitution elasticity with those derived earlier by Becker (1975) and Smith (1977) who proceeded by constructing synthetic cohorts from individual observations drawn from the 1960 Census and the 1967 Survey of Economic Opportunity, respectively. That is, to say, they grouped individuals by age and averaged observations over individuals at the same age so that eq. (38) reads

$$\overline{\ln h_t} = \bar{\psi} + \gamma \overline{\ln w_t} + \beta \bar{A}_t + \delta t + \bar{\varepsilon}_t, \tag{41}$$

where t denotes each age and the bars indicate means. If the value of $\bar{\psi}$ is the same at all ages (i.e. there are no cohort effects), then group means act as instruments and the ordinary least-squares estimator applied to (41) yields consistent estimates. In Becker's work, λ was estimated for white men to be 0.448 (with an estimated standard error of 0.105) and for nonwhite men to be 0.098 (with a standard error of 0.040).[111] When eq. (41) was estimated in its level form to individuals sorted by years of schooling, there was a tendency for γ to fall with years of schooling. This tendency was not apparent when eq. (41) was estimated by first-differencing the variables between successive ages. Becker did not invariably estimate positive values for γ although, when negative effects were estimated, they tended to be small (in absolute value) relative to their estimated standard errors. In Smith's research the logarithm of the wife's wage rate (again averaged over individuals at the same age) was included on the right-hand side of eq. (41). This is consistent with preferences being defined over the hours worked of the wife as well as over the husband's hours and commodity consumption and with period-specific utility not being additive in the hours worked by the husband

[111] These results of Becker's correspond to his use of three-year moving averages of the underlying data. The estimates from the original observations are similar. Smith's results (to be reported shortly) also derive from forming three-year moving averages of all variables.

and by the wife. His estimates of γ for white married men were 0.322 (standard error of 0.047) and for black married men were 0.231 (standard error of 0.107). The estimate of the logarithm of the wife's wage rate (so the effect of an evolutionary increase in the wife's wage rate on the husband's hours of work) was negative though typically it was estimated very imprecisely. The effect of the logarithm of the husband's wage on the wife's hours of work was also negative though larger (in absolute value and also in relation to its estimated standard error). A formal test of the symmetry condition of the Frisch male and female labor supply equations was not conducted.

The most stringent assumption required for data on synthetic cohorts to identify the intertemporal substitution elasticity is that $\bar{\lambda}_0$ (or $\bar{\psi}$) be constant for all age groups or, if it is not, that it be distributed independently of $\overline{\ln w_t}$. In fact, if after controlling for other effects $\bar{\lambda}_0$ is lower for those age groups with currently lower average wage rates (e.g. if younger workers have greater lifetime wealth, but at present are facing lower wage rates than older workers), then the coefficient on $\overline{\ln w_t}$ in eq. (41) will not identify the intertemporal substitution elasticity, but will incorporate vintage effects. This cohort bias can be addressed if synthetic cohorts are constructed in several different calendar years and λ_0 is allowed to have a different value for each cohort. In this event, the variables in eq. (41) would bear a subscript t for the cohort and a subscript k for the calendar year that the cohort mean was observed. This was precisely how Browning, Deaton and Irish (1983) proceeded by constructing synthetic cohorts from successive British Family Expenditure Surveys. In other words, instead of one observation on each cohort that would derive from a single cross-section of individuals, Browning, Deaton and Irish had seven observations on each cohort starting with the tax year 1970/71 and ending with 1976/77. Their cohorts were categorized in 1970/71 into five-year age-groups from 18–23 years old to 54–58 years old (so there were eight cohorts in all) and for each cohort (and for manual and nonmanual workers separately) they formed averages for married men. The hours variable measured weekly hours worked and it was the response to the Survey's question concerning "normal hours". The wage variable was defined as the ratio of "normal" wage and salary income per week (after the payment of income taxes) to "normal hours", the left-hand side variable.

The particular form specified by Browning, Deaton and Irish for the Frisch labor supply equation was[112]

$$\bar{h}_{tk} = \gamma_0' \overline{\ln \lambda_0} + \gamma_1' C_t + \gamma_2' Y_k + \gamma_3' \overline{\ln w_{tk}} + \gamma_4' \overline{(p_k/w_{tk})^{1/2}} + \cdots + \bar{\varepsilon}_{tk}, \tag{42}$$

[112] This Frisch labor supply equation is derived by differentiating the consumer's profit function, Π, with respect to w where Π is given by

$$\Pi(\lambda, p, w) = a_0\lambda^{-1} - a_1 p + a_2 w + 2\gamma_4'(pw)^{1/2} - \delta_2' p \ln(p\lambda) + \gamma_3' w \ln(w\lambda),$$

and where a_1 and a_2 are permitted to depend upon variables other than p, w, and λ.

where the bar indicates the average value of the variable for cohort t in calendar year k. C denotes a vector of cohort dummy variables, Y a vector of calendar year dummy variables, and the dots indicate other variables included in the equation (see Table 1.22). All cohorts were assumed to face the same commodity prices, p_k, in any calendar year k. If γ_4' is zero, commodities and male labor supply are additive within periods. The intertemporal substitution elasticity [the derivative of the logarithm of $\bar{h}_{tk}$ in eq. (42) with respect to the logarithm of w_{tk}] is given by $\bar{h}_{tk}^{-1}[\gamma_3' - (1/2)\gamma_4'(p_k/w_{tk})^{1/2}]$.

Their estimate of γ_3' in eq. (42) was 17.2 with an estimated standard error of 5.5 and their estimate of γ_4' was 26.0 with an estimated standard error of 10.5. Evaluated at approximate mean values, an intertemporal substitution elasticity of 0.05 was implied.[113] Unlike MaCurdy's results, these estimates were sensitive to the omission of the calendar year dummies. The estimate of γ_4' implies that, within each period, leisure time and commodity consumption are complements. The first-differenced (over calendar time) version of eq. (42) where $\overline{\Delta \ln w_{tk}}$ and $\overline{\Delta (p_k/w_{tk})^{1/2}}$ were instrumented yielded similar point estimates to those from fitting eq. (42) in level form although standard errors were larger and the test statistics fell slightly short of standard threshold levels. (See Table 1.22 for the instruments used.) Again the estimates were sensitive to the omission of the calendar year dummy variables.

Browning, Deaton and Irish's survey data also provided information on consumption expenditures though, unlike working hours, these represented actual and not "normal" consumption. They reported the consequences of estimating the Frisch commodity demand equation corresponding to eq. (42):

$$\bar{x}_{tk} = \delta_0' \overline{\ln \lambda_0} + \delta_1' C_t + \delta_2' \overline{\ln p_k} + \delta_3' \overline{(w_{tk}/p_k)^{1/2}} + \cdots + \bar{\varepsilon}_{tk}', \tag{43}$$

where symmetry would require δ_3' to equal $-\gamma_4'$ in eq. (42).[114] The estimated consumption intertemporal substitution elasticity, the effect of a proportional increase in p over the life cycle, is measured to be -1.38.[115] In the estimates of eq. (43) and of its first-differenced version, the value of δ_3' implied that within-period commodity consumption and leisure time are substitutes, a result contradicting the estimates of γ_4' in eq. (42). In the first-differenced equation, however, the estimate of δ_3' is less than its estimated standard error.

[113] These estimates are evaluated at mean weekly hours of work of 43.6 and the approximate mean of $\overline{(p_k/w_{tk})^{1/2}}$, namely 1.15.

[114] Because all cohorts are assumed to face the same commodity prices, the vector of year dummies (Y_k) and the term $\ln p_k$ cannot both be included in this equation. Including $\ln p$ and excluding Y is equivalent to including Y and restricting the coefficients on all the elements of Y to be the same. In fact, an F-test did not reject that constraint. Equation (43) may be derived from the consumer's profit function in footnote 112 by taking the derivative of the negative of Π with respect to p.

[115] The consumption intertemporal substitution elasticity is given by $x^{-1}[\delta_2' - (1/2)\delta_3'(w/p)^{1/2}]$. The statement in the text is derived by evaluating their estimated eqs. 5.10 and 6.5 at $x = 53.3$ and $(w/p)^{1/2} = 0.87$.

As shown in the discussion of Becker's and Smith's research above, a single cross-section of individuals may be used to compute the intertemporal substitution elasticity if synthetic cohorts are constructed from these data. However, under a string of exacting assumptions, the individual observations from a cross-section may be used more conventionally to estimate this elasticity. The essential idea here starts by recognizing that the unobserved variable λ_{0i} is a function of an individual's lifetime wage path and his initial wealth and it continues by noting that, if lifetime wages and initial wealth can be expressed as a function of age and age-invariant characteristics, then λ_{0i} in the Frisch labor supply equation may be replaced by these variables. In particular, MaCurdy (1982) replaced ψ_i in eq. (38) by variables measuring each individual's father's education, his mother's education, the socio-economic status of his parents, and the individual's own education and then fitted the resulting equation using observations on 561 white, continuously married, prime-age men from the Michigan PSID. He estimated this equation with each year's observations from 1967 to 1975 so there were nine separate estimates for the coefficient on $\ln w_{it}$, estimates of γ according to eq. (38). The estimates of γ ranged from a low of -0.07 in the 1975 cross-section to a high of 0.28 in the 1974 cross-section with estimated standard errors of 0.23 and 0.47, respectively. The simple average of these nine estimates of γ was 0.15. Only the age squared and education squared variables are identifying the variation in predicted wages, so it is not surprising that none of these nine coefficients passed the conventional thresholds of being significantly different from zero. These imprecise estimates are not very encouraging with respect to the use of individual observations from a single cross-section to measure the intertemporal substitution elasticity in this way.[116]

To summarize, the estimates to date of the male intertemporal substitution elasticity, γ, range from -0.07 to 0.45 with a central tendency of 0.20 (see Table 1.22). This means that evolutionary changes in wage rates generate relatively small changes in the hours worked of men aged from about 25 to 65 years: a 10 percent increase in his wages will induce about a 2 percent increase in his hours worked. The estimated standard errors surrounding these point estimates are also worthy of note: as often as not, the null hypothesis that life-cycle changes in wages have no effect on hours worked by prime-aged men cannot be rejected at conventional levels of significance. There is ample support here for someone whose research ignores the effects of evolutionary changes of wages on male hours worked.

It is important to note that the research described in the preceding paragraphs is directed towards only one part of the life-cycle characterization; it supplies

[116] MaCurdy (1983) uses cross-section consumption data from the Denver Income Maintenance Experiment to estimate the within-period marginal rate of substitution between commodity consumption and hours of work and then proceeded to the longitudinal dimension of the data to estimate a particular monotonic transformation of the utility function. The 121 men studied appear to display implausibly large wage elasticities though the reasons for the peculiar results are not apparent.

information on how an individual will allocate his working hours as he ages in response to evolutionary changes in his wage rates. In addition, there is the question of the response of labor supply at any age to changes in the entire wage profile. That is, two individuals both at age t' and facing the same wages at t' will supply different hours of work at t' (and at all other ages) if their entire life-cycle wage profiles differ (i.e. if their wages at ages other than t' differ). Answering this question requires relating each individual's marginal utility of wealth variable, λ_{0i}, or its transform such as ψ_i in eq. (38) to each individual's lifetime budget constraint variables, his rate of time preference, A_{it}, and ε_{it}. For male workers this second step seems to have been undertaken only by MaCurdy (1981) who relates his estimated fixed effects for different workers in eq. (38) to exogenous, age-invariant variables that determine each individual's lifetime budget constraint. These variables consist of family background characteristics, terms in the individual's own schooling, and estimated parameters describing the life-cycle growth in wage rates and initial nonwage income. His estimates suggest that, if a consumer experiences a ten percent increase in wage rates at all ages, he will increase his hours of work at all ages by between 0.5 and 1.3 percent. Again, the supply schedule of male hours of work is relatively inelastic with respect to the life-cycle wage profile.

Empirical research at the microeconomic level on male life-cycle labor supply is barely a few years old so surely it is premature to offer a confident evaluation of its performance. Some provisional judgments can be made, however. Does the extensively-used intertemporally additive model incorporate the essential features of life-cycle decision-making? The capacity of the model to take account of many aspects of intertemporal decision-making is really quite impressive. Not merely can it, in principle, be set in a context of uncertainty, but it can be generalized to allow for human capital investment, transactions costs associated with the purchase of consumer durables, and a variety of capital market imperfections (such as differential borrowing and lending rates of interest or transactions costs in financial capital markets). See MaCurdy (1981b). These are all prevalent features of the economy so their tractability within this life-cycle model adds to its appeal.

At the same time, the empirical implementation of this model already makes great demands on available data and augmenting the model to allow for these additional features probably exceeds the capacities of current data sets. If this is the case, then one response is to embark on the collection of more and more detailed information. Perhaps this should be done, but it should not proceed without some assessment of whether this extraordinary effort and expense will yield sufficiently high returns and this, in turn, requires some evaluation of whether the relationships emphasized in the life-cycle literature are important enough to account for the key variations in male labor supply.

At this stage of the research, the focus of the life-cycle research has been upon the labor supply responses to evolutionary movements in wages. The evidence to date indicates that these labor supply responses for prime-age men are very inelastic with respect to life-cycle changes in wages. Similarly, across male workers, the labor supply responses to differences in entire wage profiles appear to be small. In other words, the greater part of the variations in male labor supply across workers and over time is left unexplained by this research.[117] A great deal of effort has been brought to bear on what appears to be relationships of second-order of importance.

6. Conclusions

A great deal of research, much of it careful and some of it ingenious, has been undertaken on male labor supply during the past two decades. The vast proportion of that work – both that based on the static model and that based on the life-cycle model – indicates that the elasticities of hours of work with respect to wages are very small. In other words, the focus of most economists' research has been on behavioral responses that for men appear to be of a relatively small order of magnitude. In the case of applications of the static model of labor supply, there are a number of instances in which the income-compensated wage elasticity of hours of work is estimated to be negative. This, of course, violates an important (some would judge it to be "the" important) implication of that model and consequently it casts doubt on the empirical relevance of the model.

Of course, the static model can always be rescued from such a conclusion by arguing that what is at fault is not the allocation model itself, but rather the string of auxiliary hypotheses (assumptions about functional forms, measurement of the variables, etc.) that are required to apply the theory. Logically, this is a fully defensible position: that the theory's implications are at variance with observation means that at least one (and perhaps no more than one) of the hypotheses associated with the theory and its application is refuted. The problem with this defence is that, if the auxiliary hypotheses are continually being called upon to "save" the theory, then this comes close to denying the theory can ever be tested. It is not as if the model has already survived many different attempts to refute it. If this were the case, a few instances of its apparent failure might be attributed to the nonsatisfaction of the auxiliary assumptions. But, with this model, few scholars have conducted their research with the aim of testing the theory; most have been interested in quantifying a relationship whose existence is presumed to be true. As a by-product of this concern with measurement, they

[117] Some indication of this is provided by the consequences of fitting the hours of work equation whose estimates are reported in Table 1.17 above to the sample of 23,059 men stratified by years of age. In other words, I estimated 31 ordinary least-squares regressions, each one fitted to the hours worked and other data for men at each of the 31 years of age from 25 years to 55 years. All the right-hand side variables listed in Table 1.17 (except, of course, the age variables) were used as regressors. The size of the samples ranged from 514 men for those aged 55 years to 1,154 men for those aged 32 years. As illustrative of the poor explanatory power of the estimated linear combination of the right-hand side variables, the central tendency of the R^2s in these equations was 20% with a range extending from a high of 0.307 for men aged 45 years to a low of 0.135 for men aged 48 years.

have turned up a number of instances in which the behavioral responses take on values that violate the theory's predictions. Under these circumstances, the scientific procedure is surely to regard the theory as it has been formulated and applied to date as having been refuted by the evidence.

This does not mean that budget constraints have nothing to do with male hours of work. On the contrary, evidence from the Negative Income Tax experiments strongly suggests that changes in male work behavior are not independent of changes in their budget constraints. So prices and wages affect work decisions, but perhaps not in the particular way described by the familiar constrained utility-maximizing model. Or this model may be an apt description of some of the population, but a different characterization of behavior may be more appropriate for others. In this case, no *single* model of labor supply is adequate to account for the behavior of all individuals.

There is still much more work to be done with the canonical model. My severe judgments about its empirical relevance will have to be revised if it is shown that its apparent shortcomings to date are, in fact, the consequence of the manner in which it has been applied. If this is the case, then I hope more research with individual or household data will be conducted into the model's implications for the consumption of commodities and for savings. Consumption and savings behavior is supposed to be part of the same allocation process as hours of work and yet the empirical work on these issues has only recently explicitly recognized this. Also, I hope more will be done to integrate time spent in unemployment with decisions concerning hours of work. Current research treats unemployment in different ways: sometimes unemployment is classified as a state indistinguishable from being out of the labor force; sometimes time spent in unemployment is simply added to hours worked in the belief that both activities represent the supply of time to market activities; and sometimes time spent in unemployment is characterized as part of the optimal allocation of an individual's scarce resources, but as behaviorally distinct from hours worked. Little research has been directed towards determining which of these different treatments is the correct one. Furthermore, given the substantial resources that have already been directed towards measuring the effects of wages on work behavior and given the relatively small responses to wages that have been estimated for men, it would be useful if economists redirected some of these efforts into accounting more satisfactorily for variations in labor supply that are associated with other variables. In particular, because only a relatively small proportion of the variation in hours of work of prime-age men in the population is removed by the set of variables on which information is collected in most surveys, we need to know more about what this "unobserved heterogeneity" represents. Are these differences attributable to differences in the particular forms of the employment contracts under which individuals work? Are they associated with differences in discount rates among individuals? Are they attributable to attitudes and values that seem to be acquired from parents? There is a great deal that we do not know and that is waiting to be discovered.

References

Abbott, M. and O. Ashenfelter (1976) "Labour supply, commodity demand, and the allocation of time", *Review of Economic Studies*, 43(3):389–411.

Abbott, M. and O. Ashenfelter (1979) "Labour supply, commodity demand, and the allocation of time: correction", *Review of Economic Studies*, 46(3):567–569.

Abowd, John M. and Orley Ashenfelter (1981) "Anticipated unemployment, temporary layoffs, and compensating wage differentials", in: Sherwin Rosen, ed., *Studies in labor markets*. National Bureau of Economic Research, Univ. of Chicago Press, 141–170.

Abowd, John M. and David E. Card (1983) "Intertemporal substitution in the presence of long term contracts", Working Paper No. 166, Industrial Relations Section, Princeton University, September.

Ainsworth, R. B. (1949) "Earnings and working hours of manual wage-earners in the United Kingdom in October 1938", *Journal of the Royal Statistical Society, Series A*, 112:35–58.

Alchian, Armen A. (1974) "Information, martingales, and prices", *The Swedish Journal of Economics*, 76(1):3–11.

Altonji, Joseph G. (1982) "Intertemporal substitution model of labour market fluctuations: an empirical analysis", *Review of Economic Studies*, 49(159) Special Issue:783–824.

Altonji, Joseph G. (1983) "Intertemporal substitution in labor supply: evidence from micro data", Hoover Institution Conference on Labor Economics, January 1983.

Ashenfelter, Orley (1978) "The labor supply response of wage earners", in: John L. Palmer and Joseph A. Pechman, eds., *Welfare in rural areas: the North Carolina–Iowa income maintenance experiment*. Washington, D.C.: Brookings, 109–138.

Ashenfelter, Orley (1978), "Unemployment as a constraint on labour market behaviour", in: M. J. Artis and A. R. Nobay, eds., *Contemporary economic analysis*. The Association of University Teachers of Economics, 149–181.

Ashenfelter, Orley (1980) "Unemployment as disequilibrium in a model of aggregate labor supply", *Econometrica*, 48(3):547–564.

Ashenfelter, Orley (1983) "Determining participation in income-tested social programs", *Journal of the American Statistical Association*, 78(383):517–525.

Ashenfelter, Orley, and John Ham (1979) "Education, unemployment, and earnings", *Journal of Political Economy*, 87(5):99–116.

Ashenfelter, Orley and James Heckman (1973) "Estimating labor supply functions", in: G. G. Cain and H. W. Watts, *Income maintenance and labor supply*. Chicago: Markham, 265–278.

Ashenfelter, Orley and James Heckman (1974) "The estimation of income and substitution effects in a model of family labor supply", *Econometrica*, 42(1):73–85.

Ashworth, J. S. and D. T. Ulph (1981) "Endogeneity I: estimating labour supply with piecewise linear budget constraints", in C. V. Brown, ed., *Taxation and labour supply*. London: George Allen and Unwin, 53–68.

Atkinson, A. B. and N. H. Stern (1979) "A note on the allocation of time", *Economics Letters*, 3:119–123.

Atkinson, A. B. and N. H. Stern (1980) "On the switch from direct to indirect taxation", *Journal of Public Economics*, 14(2):195–224.

Atkinson, A. B. and N. H. Stern (1981) "On labour supply and commodity demands", in: Angus Deaton, ed., *Essays in the theory and measurement of consumer behaviour in honour of Sir Richard Stone*. Cambridge: Cambridge University Press, 265–296.

Bancroft, Gertrude (1958) *The American labor force: its growth and changing composition*. New York: Social Science Research Council.

Barmby, Tim, Richard Blundell and Ian Walker (1983) "Estimating a life cycle consistent model of family labour supply with cross-section data", unpublished paper, University of Manchester.

Barnett, William A. (1979) "The joint allocation of leisure and goods expenditure", *Econometrica*, 47(3):539–564.

Barnett, William A. (1981) *Consumer demand and labor supply*. Amsterdam: North-Holland.

Becker, Gary S. (1965) "A theory of the allocation of time", *Economic Journal*, 75:493–517.

Becker, Gary S. (1974) "A theory of social interactions", *Journal of Political Economy*, 82(6):1063–1093.

Becker, Gary S. (1975) "The allocation of time over the life cycle", in: G. R. Ghez and G. S. Becker, *The allocation of time and goods over the life cycle*. New York: National Bureau of Economic

Research, Columbia University Press, 83–132.
Ben-Porath, Yoram (1973) "Labor force participation rates and the supply of labor", *Journal of Political Economy*, 81(3):697–704.
Betancourt, Roger R. (1971) "The estimation of price elasticities from cross-section data under additive preferences", *International Economic Review*, 12(2):283–292.
Bloch, Farrell (1973) "The allocation of time to market and non-market work within a family unit", unpublished Ph.D. dissertation, Department of Economics, Stanford University.
Blundell, Richard and Ian Walker (1982) "Modelling the joint determination of household labour supplies and commodity demands", *Economic Journal*, 92(366):351–364.
Blundell, Richard and Ian Walker (1983) "Limited dependent variables in demand analysis: an application to modelling family labour supply and commodity demand behaviour", Discussion Paper ES126, Department of Econometrics and Social Statistics, University of Manchester.
Blundell, Richard and Ian Walker (1984) "A life-cycle consistent empirical model of family labor supply using cross-section data", unpublished paper, March.
Borjas, George J. (1980) "The relationship between wages and weekly hours of work: the role of division bias", *Journal of Human Resources*, 15(3):409–423.
Boskin, Michael J. (1973) "The economics of labor supply", in: G. G. Cain and H. W. Watts, *Income maintenance and labor supply*. Chicago: Markham, 163–181.
Bowen, William G. and T. Aldrich Finegan (1965) "Labor force participation and unemployment;, in: Arthur M. Ross, ed., *Employment policy and the labor market*. University of California Press, 115–161.
Bowen, William G. and T. Aldrich Finegan (1969) *The economics of labor force participation*. Princeton University Press.
Brown, C. V., E. Levin and D. T. Ulph (1976) "Estimates of labour hours supplied by married male workers in Great Britain", *Scottish Journal of Political Economy*, 23(3):261–277.
Brown, C. V. (1980) *Taxation and the incentive to work*. Oxford: Oxford University Press.
Brown, C. V., ed. (1981) *Taxation and labour supply*. London: George Allen and Unwin.
Brown, C. V., E. J. Levin, P. J. Rosa, P. J. Ruffell and D. T. Ulph (1982–83) "Direct taxation and short run labour supply", H. M. Treasury Project, Working Papers Nos. 1 to 12, Department of Economics, University of Stirling.
Browning, M. J. (1982) "Profit function representations for consumer preferences", mimeo, Bristol University.
Browning, Martin, Angus Deaton and Margaret Irish (1983) "A profitable approach to labor supply and commodity demands over the life-cycle", mimeo.
Bry, Gerhard (1959) *The average workweek as an economic indicator*, Occasional Paper 69, National Bureau of Economic Research.
Burtless, Gary and David Greenberg (1982) "Inferences concerning labor supply behavior based on limited duration experiment", *American Economic Review*, 72(3):488–497.
Burtless, Gary and Jerry A. Hausman (1978) "The effect of taxation on labor supply: evaluating the Gary negative income tax experiment", *Journal of Political Economy*, 86(6):1103–1130.
Cain, Glen G. and Harold W. Watts, eds. (1973) *Income maintenance and labor supply*. Chicago: Markham.
Chipman, John (1965) "A survey of the theory of international trade: Part 2, the neo-classical theory", *Econometrica*, 33(4):685–760.
Coase, R. H. (1937) "The nature of the firm", *Economica*, 4(16):386–405.
Coen, Robert M. and Bert G. Hickman (1970) "Constrained joint estimation of factor demand and production functions", *Review of Economics and Statistics*, 52(3):287–300.
Cogan, J. F. (1981) "Fixed costs and labor supply", *Econometrica*, 49(4):945–964.
Cohen, Malcolm S., Samuel A. Rea and Robert I. Lerman (1970) "A micro model of labor supply", BLS Staff Paper 4, U.S. Department of Labor.
Coleman, Thomas S. (1983) "Development and estimation of a dynamic model of labor supply", unpublished mimeo, University of Chicago.
Corry, B. A. and J. A. Roberts (1970) "Activity rates and unemployment: the evidence of the United Kingdom 1951–66", *Applied Economics*, 2(3):179–201.
Corry, B. A. and J. A. Roberts (1974) "Activity rates and unemployment. The U.K. experience: some further results", *Applied Economics*, 6(1):1–22.
Darrough, M. N. (1977) "A model of consumption and leisure in an intertemporal framework: a

systematic treatment using Japanese data", *International Economic Review*, 18(3):677–96.
DaVanzo, Julie, Dennis N. DeTray and David H. Greenberg (1973) "Estimating labor supply response: a sensitivity analysis", R-1372-OEO, The RAND Corporation.
Deardorff, A. V. and F. P. Stafford (1976) "Compensation of cooperating factors", *Econometrica*, 44(4):671–684.
Deaton, Angus (1974) "A reconsideration of the empirical implications of additive preferences", *Economic Journal*, 84(334):338–348.
Deaton, Angus (1982) "Model selection procedures, or, does the consumption function exist?", in: G. C. Chow and P. Crosi, eds., *Evaluating the reliability of macroeconomic models*. New York: John Wiley, 43–65.
Deaton, Angus and John Muellbauer (1980) *Economics and consumer behavior*. Cambridge: Cambridge University Press.
Deaton, Angus and John Muellbauer (1981) "Functional forms for labour supply and commodity demands with and without quantity restrictions", *Econometrica* 49(6):1521–1532.
Dickinson, Jonathan (1974) "Labor supply of family members", in: James N. Morgan, et al., eds., *Five thousand American families — patterns of economic progress*. Michigan, vol. I, 177–250.
Dickinson, Jonathan G. (1979) "Revealed preferences, functional form, and labor supply", unpublished mimeo.
Dickinson, Jonathan G. (1980) "Parallel preference structures in labor supply and commodity demand: an adaptation of the Gorman polar form", *Econometrica*, 48(7):1711–1725.
Douglas, Paul H. (1934) *The theory of wages*. New York: Macmillan.
Duncan, Beverly (1965) "Dropouts and the unemployed", *Journal of Political Economy*, 73(2):121–134.
Durand, John D. (1948) *The labor force in the United States, 1890–1960*. New York: Social Science Research Council.
Ehrenberg, Ronald G. and Paul L. Schumann (1981) "The overtime pay provisions of the fair labor standards act", in: Simon Rottenberg, ed., *The Economics of Legal Minimum Wages*. American Enterprise Institute, 264–295.
Finegan, T. A. (1962) "Hours of work in the United States—a cross-sectional analysis", *Journal of Political Economy*, 70(5):452–70.
Fleisher, Belton M., Donald O. Parsons and Richard D. Porter (1973) "Asset adjustments and labor supply of older workers", in: G. G. Cain and H. W. Watts, *Income maintenance and labor supply*. Chicago: Markham, 279–327.
Frain, LaRue (1929) "The relation between normal working time and hourly and weekly earnings", *Quarterly Journal of Economics*, 43(2):544–505.
Garfinkel, Irwin (1973) "On estimating the labor supply effects of a negative income tax", in: G. G. Cain and H. W. Watts, *Income maintenance and labor supply*. Chicago: Markham, 205–264.
Goldberger, Arthur S. (1981) "Linear regression after selection", *Journal of Econometrics*, 15(3):357–366.
Greenberg, David H. and Marvin Kosters (1973) "Income guarantees and the working poor: the effect of income-maintenance programs on the hours of work of male family heads", in: G. G. Cain and H. W. Watts, *Income maintenance and labour supply*. Chicago: Markham, 14–101.
Greenberg, David, Robert Moffitt and John Friedman (1981) "Underreporting and experimental effects on work effort: evidence from the Gary income maintenance experiment", *Review of Economics and Statistics*, 63(4):581–589.
Greenhalgh, Christine A. (1979) "Male labour force participation in Great Britain", *Scottish Journal of Political Economy*, 26(3):275–286.
Grossman, Michael (1975) "The correlation between health and schooling", in: Nestor E. Terleckyj, ed., *Household production and consumption*. New York: National Bureau of Economic Research, Columbia University Press, 147–211.
Hall, Robert E. (1973) "Wages, income, and hours of work in the U.S. labor force", in: G. G. Cain and H. W. Watts, *Income maintenance and labor supply*. Chicago: Markham, 102–162.
Hall, Robert E. (1980) "Employment fluctuations and wage rigidity", *Brookings Papers in Economy Activity*, 1:91–123.
Ham, John C. (1982) "Estimation of a labor supply model with censoring due to unemployment and underemployment", *Review of Economic Studies*, 49(157):335–354.
Ham, John C. (1983) "Testing whether unemployment represents life-cycle labor supply behavior", unpublished paper, Nov. 1983.
Hausman, Jerry A. (1981) "Labor supply", in: Henry J. Aaron and Joseph A. Pechman, eds., *How*

taxes affect economic behavior. Washington, D.C.: The Brookings Institution, 27–72.
Hausman, Jerry and Paul Ruud (1984) "Family labor supply with taxes", *American Economic Review, Papers and Proceedings*, 74(2).
Hausman, J. A. and D. A. Wise (1976) "The evaluation of results from truncated samples: the New Jersey income maintenance experiment", *Annals of Economic and Social Measurement*, 5(4):421–46
Hausman, Jerry A. and David A. Wise (1977) "Social experimentation, truncated distributions, and efficient estimation", *Econometrica*, 45(4):919–938.
Hausman, Jerry A. and David A. Wise (1979) "Attrition bias in experimental and panel data: the Gary income maintenance experiment", *Econometrica*, 47(2):455–474.
Haveman, R. H. and B. L. Wolfe (1984) "The decline in male labor force participation: comment", *Journal of Political Economy*, 92(3):532–41.
Heckman, James J. (1974a) "Life cycle consumption and labor supply: an explanation of the relationship between income and consumption over the life cycle", *American Economic Review*, 64(1):188–194.
Heckman, James J. (1974b) "Shadow prices, market wages, and labor supply", *Econometrica*, 42(4):679–694.
Heckman, James J. (1976a) "A life-cycle model of earnings, learning, and consumption", *Journal of Political Economy*, Part 2, 84(4):11–44.
Heckman, James J. (1976b) "The common structure of statistical models of truncation, sample selection, and limited dependent variables and a simple estimator for such models", *Annals of Economic and Social Measurement*, 5(4):475–492.
Heckman, James J. (1978) "A partial survey of recent research on the labor supply of women", *American Economic Review Proceedings*, 68(2):200–207.
Heckman, James J. (1983) "Comment [on Hausman's paper]", in: Martin Feldstein, ed., *Behavioral simulation methods in tax policy analysis*. Chicago: National Bureau of Economic Research, University of Chicago Press, 70–82.
Heckman, James J., Mark Killingsworth and Thomas MaCurdy (1981) "Empirical evidence on static labour supply models: a survey of recent developments", in: Z. Hornstein, J. Grice and A. Webb, eds., *The economics of the labour market*. London: Her Majesty's Stationery Office, 74–122.
Heckman, James J. and Thomas E. MaCurdy (1981) "New methods for estimating labor supply functions: a survey", *Research in Labor Economics*, 4:65–102.
Hill, C. Russell (1973) "The determinants of labor supply for the working urban poor", in: G. G. Cain and H. W. Watts, *Income maintenance and labor supply*. Chicago: Markham, 182–204.
Hill, Daniel and Saul Hoffman (1977) "Husbands and wives", in: Gregg J. Duncan and James N. Morgan, eds., *Five thousand American families — patterns of economic progress*. Michigan, vol. V, 29–70.
Holbrook, Robert and Frank Stafford (1971) "The propensity to consume separate types of income: a generalized permanent income hypothesis", *Econometrica*, 39(1):1–22.
Horner, David (1977) "Labor supply of husbands", in: Harold Watts and Albert Rees, eds., *The New Jersey income maintenance experiment, Vol. II: labor supply responses*. New York: Academic Press.
Hotz, V. J., F. E. Kydland, and G. L. Sedlacek, "Intertemporal substitution and labor supply", unpublished paper, December 1982.
Hurd, Michael D. (1976) "The estimation of nonlinear labor supply functions with taxes from a truncated sample", Institute for Mathematical Studies in the Social Sciences, Technical Report No. 217, Stanford University.
Hurd, Michael D. and John Pencavel (1981) "A utility-based analysis of the wage subsidy program", *Journal of Public Economics*, 15:185–201.
Jevons, W. Stanley (1888) *The theory of political economy*. London: Macmillan.
Johnson, Terry R. and John Pencavel (1982) "Forecasting the effects of a negative income tax program", *Industrial and Labor Relations Review*, 35(2):221–234.
Johnson, Terry R. and John Pencavel (1984) "Dynamic hours of work functions for husbands, wives, and single females", *Econometrica*, 52(2):363–390.
Jones, Ethel B. (1961) "Hours of work in the United States, 1900–1957", unpublished Ph.D. dissertation, University of Chicago.
Jones, Ethel B. (1974) *An investigation of the stability of hours of work per week in manufacturing, 1947–1970*. College of Business Administration, University of Georgia.
Kalachek, Edward, Wesley Mellow and Frederic Raines (1978) "The male labor supply function

reconsidered", *Industrial and Labor Relations Review*, 31(3):356–367.

Kalachek, Edward D. and Frederic Q. Raines (1970) "Labor supply of lower income workers and the negative income tax", *Technical studies, the President's commission on income maintenance programs*. Washington, D.C.: U.S. Government Printing Office, 159–186.

Keeley, Michael C. (1981) *Labor supply and public policy: a critical review*. New York: Academic Press.

Keeley, Michael C. and Philip K. Robins (1980) "The design of social experiments: a critique of the Conlisk-Watts assignment model", in: R. G. Ehrenberg, ed., *Research in labor economics*. Vol. 3, 293–333.

Keeley, Michael C., Philip K. Robins, Robert G. Spiegelman and Richard W. West (1978a) "The labor supply effects and costs of alternative negative income tax programs", *Journal of Human Resources*, 13(1):3–36.

Keeley, Michael C., Philip K. Robins, Robert G. Spiegelman and Richard W. West (1978b) "The estimation of labor supply models using experimental data", *American Economic Review*, 68(5):873–887.

Killingsworth, Mark R. (1981) "A survey of labor supply models: theoretical analyses and first-generation empirical results", *Research in Labor Economics*, 4:1–64.

Killingsworth, Mark R. (1983) *Labor supply*. Cambridge: Cambridge University Press.

Klein, L. R. and R. F. Kosobud (1961) "Some econometrics of growth: great ratios of economics", *Quarterly Journal of Economics*, 75(2):173–198.

Kniesner, Thomas J. (1976a) "An indirect test of complementarity in a family labor supply model", *Econometrica*, 44(4):651–669.

Kneisner, Thomas J. (1976b) "The full-time workweek in the United States, 1900–1970", *Industrial and Labor Relations Review*, 30(1):3–15.

Kosters, Marvin H. (1966) "Income and substitution effects in a family labor supply model", P-3339, The Rand Corporation.

Kosters, Marvin H. (1969) "Effects of an income tax on labor supply", in: Arnold C. Harberger and Martin J. Bailey, eds., *The taxation of income from capital*. Washington, D.C.: Studies of Government Finance, Brookings Institution, 301–324.

Kurz, Mordecai, Philip Robins, Robert Spiegelman, Richard West and Harlan Halsey (1974) "A cross sectional estimation of labor supply for families in Denver 1970", Research Memorandum 24, Center for the Study of Welfare Policy, Stanford Research Institute.

Kydland, Finn E. and Edward C. Prescott (1982) "Time to build and aggregate fluctuations", *Econometrica*, 50(6):1345–1370.

Lakatos, Imre (1970) "Falsification and the methodology of scientific research programmes", in: Imre Lakatos and Alan Musgrave, eds., *Criticism and the growth of knowledge*. Cambridge University Press, 1970, pp. 91–195.

Lau, Lawrence J., Wuu-Long Lin and Pan A. Yotopoulos (1978) "The linear logarithmic expenditure system: an application to consumption-leisure choice", *Econometrica*, 46(4): 843–868.

Layard, Richard (1978) "Hours supplied by British married men with endogenous overtime", Discussion Paper No. 30, Centre for Labour Economics, London School of Economics.

Lewis, H. Gregg (1957) "Hours of work and hours of leisure", in: *Proceedings of the Ninth Annual Meeting*. Industrial Relations Research Association, 196–206.

Lewis, H. Gregg (1967) "On income and substitution effects in labor force participation", unpublished paper, University of Chicago.

Lewis, H. Gregg (1969) "Employer interests in employee hours of work", unpublished paper.

Long, Clarence D. (1958) *The labor force under changing income and employment*. Princeton: National Bureau of Economic Research, Princeton University Press, General Series, No. 65.

Lundberg, Shelly (1983) "Is unemployment involuntary?", Hoover Institution Working Papers in Economics No. E-83-9, Stanford University.

MaCurdy, Thomas (1976) "A household life-cycle model under uncertainty with human capital investment", mimeo, Workshop in Applications of Economics, University of Chicago.

MaCurdy, Thomas (1981) "An empirical model of labor supply in a life-cycle setting", *Journal of Political Economy*, 89(6):1059–1085.

MaCurdy, Thomas (1981) "An intertemporal model of portfolio choice and human capital accumulation under uncertainty with extensions incorporating taxes, consumer durables, imperfections in capital markets, and nonseparable preferences", Hoover Institution Working Papers in Economics

No. E-81-18.

MaCurdy, Thomas E. (1982) "Interpreting empirical models of labor supply in an intertemporal framework with uncertainty", Research Paper No. 23, Stanford Workshop on Factor Markets.

MaCurdy, Thomas E. (1983) "A simple scheme for estimating an intertemporal model of labor supply and consumption in the presence of taxes and uncertainty", *International Economic Review*, 24(2): 265–289.

Manser, Marilyn and Murray Brown (1979) "Bargaining analyses of household decisions", in: C. B. Lloyd, E. S. Andrews and C. L. Gilroy, eds., *Women in the labor market*. New York: Columbia University Press, 3–26.

Masters, Stanley and Irwin Garfinkel (1977) *Estimating the labor supply effects of income-maintenance alternatives*. New York: Institute for Research on Poverty Monograph Series, Academic Press.

McElroy, Marjorie B. (1981) "Empirical results from estimates of joint labor supply functions of husbands and wives", in: R. G. Ehrenberg, ed., *Research in Labor Economics*, 4:53–64.

Metcalf, Charles E. (1973) "Making inferences from controlled income maintenance experiments", *American Economic Review*, 63(3):478–483.

Metcalf, Charles E. (1974) "Predicting the effects of permanent programs from a limited duration experiment", *Journal of Human Resources*, 9(4):530–555.

Metcalf, David, Stephen Nickell and Ray Richardson (1976) "The structure of hours and earnings in British manufacturing industry", *Oxford Economic Papers*, 28(3):284–303.

Mincer, Jacob (1962) "Labor force participation of married women: a study of labor supply", in: *Aspects of labor economics*. National Bureau of Economic Research, Princeton University Press, 63–105.

Mincer, Jacob (1963) "Market prices, opportunity costs, and income effects", in: Carl F. Christ, et al., *Measurement in economics: studies in mathematical economics and econometrics in memory of Yehuda Grunfeld*. Stanford University Press.

Mincer, Jacob (1966) "Labor force participation and unemployment: a review of recent evidence", in: R. A. Gordon and M. S. Gordon, eds., *Prosperity and unemployment*. 73–112.

Moffitt, Robert A. (1979) "The labor supply response in the Gary experiment", *Journal of Human Resources*, 14(4): 477–487.

Moffitt, Robert A. and Kenneth C. Kehrer (1981) "The effect of tax and transfer programs on labor supply: the evidence from the income maintenance experiments", in: R. G. Ehrenberg, ed., *Research in Labor Economics*, 4:103–150.

Morgan, James N. (1979) "Hours of work by family heads: constraints, marginal choices, and income goals", in: Greg J. Duncan and James N. Morgan, eds., *Five thousand American families – patterns of economic progress*. University of Michigan, vol. VII, 63–100.

Moses, Leon N. and Harold F. Williamson, Jr. (1963) "Value of time, choice of mode, and the subsidy issue in urban transportation", *Journal of Political Economy*, 71(3):247–264.

Mroz, Thomas A. (1984) "The sensitivity of an empirical model of married women's hours of work to economic and statistical assumptions", unpublished Ph.D. thesis, Stanford University.

Muellbauer, John (1981) "Linear aggregation in neoclassical labour supply", *Review of Economic Studies*, 48(1):21–36.

Nadiri, M. I. and S. Rosen (1974) *A disequilibrium model of demand for factors of production*. New York: National Bureau of Economic Research, Columbia University Press, General Series No. 99.

Oi, Walter (1976) "Residential location and labor supply", *Journal of Political Economy*, 84(4): Part 2, S221–S238.

Orcutt, Guy H. and Alice G. Orcutt (1968) "Incentive and disincentive experimentation for income maintenance policy purposes", *American Economic Review*, 58(4):754–727.

Owen, John D. (1979) *Working hours: an economic analysis*. Lexington, Mass.: Lexington Books.

Parsons, Donald O. (1980) "The decline in male labor force participation", *Journal of Political Economy*, 88(1):117–134.

Pencavel, John (1977) "Work effort, on-the-job screening, and alternative methods of remuneration", in: R. G. Ehrenberg, ed., *Research in Labor Economics*, 1:225–258.

Pencavel, John (1982) "Unemployment and the labor supply effects of the Seattle-Denver income maintenance experiments", *Research in Labor Economics*, 5:1–31.

Phelps Brown, E. H. (1960) *The growth of British industrial relations*. London.

Phlips, Louis (1978) "The demand for leisure and money", *Econometrica*, 6(5):1025–1044.

Pollak, Robert A. (1970) "Habit formation and dynamic demand functions", *Journal of Political Economy*, 78(4):745–63.

Pollak, Robert A. (1971) "Additive utility functions and linear Engel curves", *Review of Economic Studies*, 38(116):401–414.
Rees, Albert (1974) "An overview of the labor supply results", *Journal of Human Resources*, 9(2):158–180.
Rees, Albert (1979) "Douglas on wages and the supply of labor", *Journal of Political Economy*, 87(5): Part 1, 915–922.
Robbins, Lionel (1930) "On the elasticity of demand for income in terms of effort", *Economica*, 10(29):123–129.
Robins, Philip K. and Richard W. West (1980) "Program participation and labor supply response", *Journal of Human Resources*, 15(4):499–544.
Rosen, Harvey S. (1976) "Taxes in a labor supply model with joint wage-hours determination", *Econometrica*, 44(3):485–508.
Rosen, Harvey S. (1978) "The measurement of excess burden with explicit utility functions", *Journal of Political Economy, Supplement*, 86(2): Part 2, S121–S135.
Rosen, Sherwin (1969) "On the interindustry wage and hours structure", *Journal of Political Economy*, 77(2):249–273.
Rosen, Sherwin and Finis Welch (1971) "Labor supply and income redistribution", *Review of Economics and Statistics*, 53(3):278–282.
Samuelson, Paul A. (1956) "Social indifference curves", *Quarterly Journal of Economics*, 70(1):1–22.
Sargent, Thomas J. (1978) "Estimation of dynamic labor demand schedules under rational expectations", *Journal of Political Economy*, 86(6):1009–1044.
Schoenberg, E. and P. Douglas (1937) "Studies in the supply curve of labor: the relation in 1929 between average earnings in American cities and the proportion seeking employment", *Journal of Political Economy*, 45(1):45–79.
Simon, Herbert A. (1951) "A formal theory of the employment relationship", *Econometrica*, 19(3):293–305.
Smith, James P. (1975) "Assets and labor supply", R-1728-HEW, The Rand Corporation.
Smith, James P. (1977) "Family labor supply over the life cycle", *Explorations in Economic Research*, 4(2):205–276.
Smith, Shirley J. (1983) "Estimating annual hours of labor force activity", *Monthly Labor Review*, 106(2):13–22.
Smith-Lovin, Lynn and Ann R. Tickamyer (1978) "Nonrecursive models of labor force participation, fertility behavior, and sex role attitudes", *American Sociological Review*, 43(4):541–557.
Spinnewyn, Frans (1981) "Rational habit formation", *European Economic Review*, 15(1):91–109.
Teper, Lazare (1932) *Hours of labor*. Baltimore: Johns Hopkins Press.
U.S. Department of Labor, *Multiple jobholders in May 1979*, Bureau of Labor Statistics, Special Labor Force Report 239, January 1981.
Wales, Terence J. (1973) "Estimation of a labor supply curve for self-employed business proprietors", *International Economic Review*, 14(1):69–80.
Wales, Terence J. (1978) "Labour supply and commuting time: an empirical study", *Journal of Econometrics*, 8(2):215–226.
Wales, T. J. and A. D. Woodland (1976) "Estimation of household utility functions and labor supply response", *International Economic Review*, 17(2):397–410.
Wales, T. J. and A. D. Woodland (1977) "Estimation of the allocation of time for work leisure, and housework", *Econometrica*, 45(1):115–132.
Wales, T. J. and A. D. Woodland (1979) "Labour supply and progressive taxes", *Review of Economic Studies*, 46(1):83–95.
Wales, T. J. and A. D. Woodland (1980) "Sample selectivity and the estimation of labor supply functions", *International Economic Review*, 21(2):437–468.
Weiss, Yoram (1972) "On the optimal lifetime pattern of labour supply", *Economic Journal*, 82(328):1293–1315.
Welch, Finis (1978) "The labor supply response of farmers", in: John L. Palmer and Joseph A. Pechman, eds., *Welfare in rural areas: the North Carolina-Iowa income maintenance experiment*. Washington, D.C.: Brookings, 77–100.
Winston, Gordon C. (1966) "An international comparison of income and hours of work", *Review of Economics and Statistics*, 48(1):28–39.
Woytinsky, W. S. (1940) *Additional workers and the volume of unemployment in the depression*. New York: Social Science Research Council.