SWEDISH ECONOMIC POLICY REVIEW 2 (1995) 391–418

Male Labour Supply in Sweden: Are Incentives Important?

Susanne Ackum Agell* and Costas Meghir**

Summary

■ We attempt to measure the magnitude of male labour supply responses to changes in taxes. We estimate models of labour supply using two alternative wage growth measures: one based on the life-cycle model and one based on the conventional "within period" model of labour supply. We also report results using a balanced panel, including changes in the unemployment rate and time trends in the equations, and using a grouped data estimator. All model specifications yield low labour supply elasticities with quite small standard errors. We therefore conclude that tax reform is unlikely to have large incentive effects on hours worked. However, a small elasticity does not necessarily imply that the excess burden of the old system was low. The reason is that the marginal tax wedges used to be so high that also a small elasticity can be translated into a large welfare loss. Moreover our analysis does not reveal the potential participation effects of the reforms. ■

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The tax reform in 1991 was preceded by a thorough discussion of its likely consequences for the Swedish economy (SOU 1989:33). The financing of the reform was considered largely in terms of predictions of its behavioural effects. Since a main feature of the reform was a substantial reduction in the marginal tax rate on labour income, special attention was focused on labour supply responses. However, the underlying empirical studies gave little guidance on this issue; see Blomquist (1989) and Aaberge *et al.* (1989). This uncertainty led the appointed commission to exercise caution when predicting the responses. In the end, the behavioural effects of the tax reform (including other behavioural changes) were assumed to be rather small: in the short run increasing from 2.5 billion Swedish kronor, then 5.5, to 7.5 billion in the first three years and in the long-run amounting to 20 billion Swedish kronor per year (SOU 1989:33, pp. 221–222).

Labour economists were quite content that the tax reform itself would provide a good opportunity to pinpoint the controversial labour supply responses: an individual's labour supply could easily be compared before and after the reform. But, when the reform was introduced in 1991 it coincided with the deepest recession in Sweden since the 1930s. In only four years' time the open unemployment rate increased more than five-fold: from 1.6 per cent in 1990 to 8.2 per cent in 1993. Thus, it became

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extremely difficult to disentangle any labour supply changes caused by the tax reform in the prevailing macroeconomic situation.

Under these circumstances, a study of likely labour supply responses to the 1991 tax reform based on pre-reform data is still justified, especially if it covers a period when wages have fluctuated in a predictable way. To that end we employ a standard life-cycle model[1] of labour supply with uncertainty to interpret observed changes in male hours of work in the engineering industry during the 1970s and 1980s. With this framework we can evaluate the extent to which anticipated wage movements, whether induced by anticipated policy changes (the 1970s were a period with rising marginal tax rates on labour income, while in the mid-1980s a reform of the tax system lowered marginal tax rates) or by cyclical and other variations in rates of pay, can explain observed changes in hours of work. The results of such an exercise can be used to infer the potential incentive effects of an unanticipated reform.

Specifically, in the absence of liquidity constraints, the life-cycle model allows us to control for unobservable wealth essentially by considering the relationship between changes in hours of work and predictable changes in wages. Under this condition the wage elasticity we estimate is an upper bound for the more conventional "within period" labour supply elasticity which reflects the trade-off between leisure and consumption at a given level of non-labour income. We also argue that under certain conditions and with the appropriate choice of wage growth measure, we can also obtain a direct estimate of the within period elasticity.

Our approach assumes that individuals can freely choose hours of work. This is made more realistic by modelling the number of hours worked over a relatively long period – a quarter rather than just a week. Nevertheless, if hours constraints are important, the focus of the debate should be on the factors that determine optimal contracts.

The paper is organised as follows. In Section 1 we give an overview of the theoretical methodology and in Section 2 we discuss the regression model and technical issues of estimation. We present our data in Section 3 and we discuss our results in Section 4. Section 5 concludes.

[1] Before this study there has, to our knowledge, been no attempt to use a life-cycle model to interpret observed changes in hours of work in Sweden. However, attempts have been made to estimate life-cycle consistent models with cross-section data (Aronsson and Wikström, 1991 and Blomquist, 1988).

1. An overview of the methodology

In this paper we attempt to measure the magnitude of the labour supply response to changes in taxes. In general a reduction in taxes will improve the incentive to work in any given period, given the wealth of the individual over his life cycle, by increasing the after-tax wage rate. Counteracting this effect will be the effects of increased wealth induced by an unexpected reduction in taxes: since pre-tax earnings are worth more in every period after the reform, the value of time will increase, leading to an increase in expected life-cycle wealth unless the reform was perfectly anticipated.

Measuring the labour supply effect of tax changes and assessing the incentive effects is complicated by the fact that many factors other than changes in the marginal rewards to work induce changes in hours of work. For instance, even if the individual labour supply function (and hence the aggregate one) is completely inelastic with respect to the wage rate, hours may fluctuate as the supply of labour responds to changes in expected wealth. In this simple case shifts in demand for labour translate only into wage changes, while shifts in labour supply lead to changes in both hours and wages; hence the data will contain both wage changes and hours changes, but the wage changes do not necessarily cause the hours to change – in fact the reverse may occur. Factors shifting labour supply of all individuals in the same direction will lead to wage changes along the demand for labour causing wage changes. Thus the empirical strategy should attempt to disentangle these effects and find the sensitivity of hours supplied to wage changes.

In our framework we assume that all prime-aged men wish to participate in the labour force. The incentive effects we analyse will relate exclusively to the number of hours worked. We design our approach in relation to available data. The model we use, which is formally described in Appendix 1, is the life-cycle model of labour supply with uncertainty (see e.g. MaCurdy, 1981 and 1984).

In the life-cycle approach to labour supply, individuals form expectations about future rates of pay and real interest rates. Such expectations will also include the anticipated effects of taxes. These components, together with any current assets, define life-cycle wealth. Given the current state of affairs and future expectations, an individual has to decide how much to work in each period. The individual will then decide to work most in periods when the rates of pay are relatively high. The fluctuation in hours of work due to movements along the anticipated path of wages reflects intertemporal substitution and the degree to which this occurs will depend on preferences. If we estimate the relevant preference coefficients we are then able to ascertain the sensitivity of labour supply to wage changes (anticipated or not) and hence infer the incentive effects of the reforms. As new information about the future is received, the individual will revise his plans. In the data we observe hours and wages over time for a particular individual. Some part of wage changes are anticipated, such as expected progress along a career path or the effects of pre-announced reforms. The remaining change in the observed marginal wage represents unanticipated events and measurement error. The empirical strategy involves estimating the anticipated component of the wage change and using this to explain the change in hours, while controlling for the effects of life-cycle wealth. The ideal data for such a strategy are relatively long time series where individuals are observed over time and where plenty of cyclical fluctuations in wages take place which are predictable from past wage movements. Thus the known past wage path for an individual can be used to predict an expected wage in the next period. In other words we can exploit the effects of predictable business-cycle fluctuations to measure the effects of unanticipated reforms. Interestingly, the wage elasticity that we measure is (under our assumptions) an upper bound on the more conventional within period wage elasticity which holds financial wealth constant (MaCurdy, 1981). This is because our elasticity abstracts from the wealth effect of the tax reforms.

The hours supplied by an individual in any period will depend on the current wage and on a function of wealth. This function, the marginal utility of wealth, will also depend on expectations about the future. The empirical strategy depends on our ability to control for the effects of life-cycle wealth and expectations. The modelling approach exploits the fact that the marginal utility of wealth remains constant over time if the individual is implementing an optimal plan, and will only change in response to unexpected events. This implies that the change in hours over time can be expressed as a function of the anticipated wage growth and a "surprise" term reflecting replanning in response to news. This surprise term need not be observed: it can be eliminated by the estimation technique by noting that it cannot be correlated with events known in the past.

The approach we describe below involves several econometric techniques whose aim is to ensure that we compute the anticipated path of wages in a robust way. Special attention will be given to allowing for measurement error in wages.

2. The regression model and technical issues of estimation

In this section we consider the technical issues involved in obtaining a credible estimate of the coefficient of intertemporal substitution, α. We discuss the sources of unexplained variation in hours, the way we deal with individual expectations and our treatment of measurement error.

In Appendix 1 we provide a more detailed description of how the model is derived. The final representation we obtain is[2]

$$E_{t-1}\Delta \ln h_{it} = \alpha E_{t-1}(\Delta \ln w_{it} - r_{t-1}) + \beta_0 + \beta_1 Age_{it} + \beta_2 Age_{it}^2, \alpha > 0. \qquad (1)$$

The above equation simply states that the rate of growth of hours for individual i in period t as anticipated in period $t-1$ ($E_{t-1}\Delta \ln h_{it}$) will be positive whenever the anticipated growth in nominal wages $E_{t-1}(\Delta \ln w_{it})$ is higher than the risk-free nominal rate of return in the financial markets (r_{t-1}). In other words if wages are expected to rise sufficiently, work is postponed to the next period. The extent to which such intertemporal substitution takes place is governed by the parameter α. Note at this point that the appropriate wage measure (w) here is the nominal after-tax hourly rate of pay: i.e., if the individual is facing a marginal tax rate of τ and earns a pre-tax wage rate per hour of w^*, the wage will be equal to $w = (1-\tau) w^*$.[3]

The practical implementation of this model in a world with uncertainty and with imperfectly measured data depends on being able to disentangle expected changes in wages from unexpected changes on the basis of observations of realised wage growth. To focus on the problem of estimating α in equation (1) we replace the expectations with the observed realised values and obtain the following statistical regression model

[2] Δ denotes a difference over time.

[3] Thus for example, if a fall in taxes is expected, all else being equal, $E_{t-1}(\Delta \ln w_{it} - r_{t-1}) = \Delta \ln(1-\tau) - r_{t-1}$. If the tax change is unexpected then (again all else being equal), $E_{t-1}(\Delta \ln w_{it} - r_{t-1}) = -r_{t-1}$.

$$\Delta \ln h_{it} = \alpha\,(\Delta \ln w_{it} - r_{t-1}) + \beta_0 + \beta_1 Age_{it} + \beta_2 Age_{it}^2 + \mu_{it} \qquad (2)$$

The term in (μ_{it}) summarises the unexplained variation in the data. This variation can be attributed to three separate sources:

(i) The first source is the divergence of individual expectations from the actual outcomes, i.e., it is due to the fact that the model we estimate, equation (2), includes measured hours and wage changes rather than their anticipated values as in equation (1). The structure of the economic model suggests that this component of the error term will not be correlated with past information, i.e., with information available in period t–1 or earlier. A complication here stems from the fact that the divergence of actual outcomes from expectations may be in the same direction for all individuals, since the unexpected event underlying such divergence may be common across all individuals. Such unexpected events could include an external trade shock or an unexpected change in the tax system. The estimation procedure has to rely on the availability of long enough periods of data so that such "common" or "macroeconomic" shocks can average out over time because (by definition) they do not average out over workers. Our estimates in turn assume that (a) common shocks are not important; (b) that common shocks can be summarised by a quadratic time trend and (c) are based on estimation procedures suitable for samples with a large number of periods, which do not rely on independence across individuals.

(ii) The second source of unexplained variation is random changes in preferences and, in particular, in the rate of time preference. This is reflected in the intercept (β_0) of equation (2) and is related to how patient the individual is and hence how willing he is to work today to reap future benefits. Random changes in the rate of time preferences may lead to some serial correlation in the regression and the estimation method is designed to take this into account.

(iii) The third important source of unexplained variation in observed hours of work is measurement error. Since equation (2) contains hours and wages in differences the measurement error component will be a moving average of order 1 or higher. Measurement error can be regarded as originating from two separate sources.

The first originates from what is known as division bias; i.e., a bias due to the fact that the wage rate is constructed by dividing quarterly earnings by quarterly hours of work. Earnings and hours in our study are reported by the firms where the workers are employed. On one hand it is likely that reported earnings do not contain much measurement error since they come directly from the accounts of the firm. On the other hand reported hours are more likely to be an inaccurate measure of the hours actually worked by the individual. Thus the derived hourly wage rate may be measured erroneously. The nature of this measurement error is such that it will induce a negative relationship between the observed wage and hours of work, since a positive error on hours will lead to a reduced measure of the wage. To circumvent such measurement error we exploit the panel structure of our data along with the reasonable assumption that the measurement error in hours is independent of the level of hours and earnings.

The second measurement error originates from the way we compute the individual's marginal tax rate. Given the available data, we had to take some short cuts in computing the marginal tax rate: in Sweden labour and capital income minus deductions on expenditures used to be aggregated for taxation purposes. However, we observe neither capital income nor worker-specific deductions. This might cause measurement errors in the observed marginal tax rate. We can write the relationship between the actual (true) and the observed (our) after-tax wage as

$$\ln(1-\tau_{it})w_{it} = \ln(1-\tau_{it}^{A} + \varepsilon_{it})w_{it} \cong \ln(1-\tau_{it}^{A})w_{it} + \frac{\varepsilon_{it}}{(1-\tau_{it}^{A})} \tag{3}$$

where τ_{it} is the observed (our) measure of the marginal tax rate and τ_{it}^{A} is the actual (true) marginal tax rate. The approximation to the measurement error, $\frac{\varepsilon_{it}}{(1-\tau_{it}^{A})}$, will be larger for individuals with higher levels of accumulated assets. Since assets tend to accumulate with age, we approximate this error as a function of age and a random component which we assume is uncorrelated with past levels of earnings. Hence the age polynomial in the hours growth equation will capture both the effects of age on work preferences and the measurement error in the marginal tax rate. In Appendix 2 we show that the distribution of marginal tax rates in our data is very similar to that computed by more accurate techniques in other surveys. Such measurement error will also lead to serial correlation.

As we pointed out when introducing our modelling approach, the data are likely to contain plenty of variation in both hours and wages. But for our purpose here we seek to identify the causal effects of anticipated wage changes on hours of work since this will measure the preference parameter α. From an econometric perspective we have to control for the fact that some of the unexplained variation in hours may itself be correlated with the wage changes, thereby generating a problem of reverse causality. Unanticipated changes in expectations about the future or random preference variation may lead to changes in hours of work and hence in the individual's tax situation; this in turn implies that after-tax wages will change. In this case the direction of causality runs from hours to wages rather than the reverse as we require.

Given our arguments, a valid approach to estimation would be to construct predictions of wage growth between period $t-1$ and t based on information known to the individual in period $t-2$ and earlier. This approach allows for limited serial correlation as implied by random preferences variation and by measurement error, given our assumptions. To correct for the problems of division bias, we use past levels of earnings which are more accurately measured, rather than using past hourly wage rates in the predictive equation. It is possible to allow for some serial correlation in the random preference variation by constructing predictions based on information in period $t-3$. This, of course, would reduce the precision of our estimates but could form a basis for testing the validity of some of our assumptions.

A crucial requirement of the empirical strategy is that past information can provide relatively precise information on current wage growth. If it does not, we would have to assign more or less the same level of anticipated wage growth to all workers and hence we would be unable to estimate any meaningful effect. Fortunately it is easy to test whether past information explains current wage growth: a simple statistical significance test of the explanatory variables included in the wage growth prediction equations (and not included in the hours growth equation) provides the necessary information. From an economic point of view it can also be argued that past levels of earnings can explain current wage growth: the level of earnings is affected by observable human capital attributes. The life-cycle pattern of wages for various skill types is different. For example, individuals with high education tend to have a steeper career path which peaks much later than the wage path of workers with lower education. The differences in career paths (and hence in wage growth) are likely to be influenced by all measures of skills – both observable and unobservable (e.g. ability). Moreover, the sensitivity of wages to cyclical variations is also likely to differ for different types of workers.

The first step in our estimations is to construct predictions of wage growth based on a regression of current wage growth on past earnings and the age terms. The hours growth is then regressed on the predicted wage growth obtained from the first-stage regression as well as on the age terms. This provides us with an estimate of α which measures the effect of anticipated wage changes on hours worked. Earlier we mentioned that we adopt different approaches depending on how we treat the presence of "common" or "macroeconomic" shocks to wages. Without such macroeconomic shocks, where all unexplained variation is uncorrelated across individuals, we can construct wage predictions based on different predictive equations for each year of the panel. This approach has the distinct advantage of providing us with very precise estimates since predictions are based on the most up-to-date information about the relationship between current wage growth and past levels of earnings. A simple modification of this approach which accounts for some common unexplained variation in hours growth across individuals is to include a quadratic time trend in the wage growth equation.

As it turns out, our data set spans a relatively long period over which we can expect most macroeconomic variation to average out. Hence we can also apply time-series methods to estimate the coefficients.[4] One way of implementing this approach, which eliminates the adverse effects of measurement error and exploits the time series dimension is as follows: in the first step we divide our data into five groups based on the individual's date of birth and construct average measures of wage growth and hours growth by year for each group. In the second step we construct predicted wage growth measures for each group based on past levels of wage growth for each group. A common regression function is used for all years in this step, in contrast to the methods outlined above where a different predictive equation was used in each year of the panel. Moreover, all the grouped data are pooled together and we regress hours growth on the predicted wage growth. We also include a quadratic time trend to account for any remaining common changes in the data. This grouping technique is very powerful in the presence of measurement error at the micro level and is described in some detail in Meghir and Whitehouse (1996). Since estimation is based on averaging across years as well as across individuals, the effects of both macroeconomic and individual specific shocks and of the measurement error are averaged out, providing us with a "clean" estimate of the effect of anticipated wage changes on hours. Nevertheless, this technique is likely to provide less precise estimates *vis-à-vis* the techniques described above which exploit all the relevant cross-section variability in earnings and hours.

[4] Here we have to ensure that the explanatory variables in the predictive equations jointly satisfy certain (stationarity) conditions.

3. Data

The data used in this study pertain to blue-collar workers from the Swedish engineering industry. In each of the years 1975, 1980 and 1985, a sample of 5 000 workers was collected by the employers' confederation in the engineering industry (*Sveriges Verkstadsförening*). The data set was constructed by appending these three samples. Information on work related issues and some individual characteristics was then gathered for each fourth quarter over the period 1970–1987. The resulting unbalanced panel covers 14 761 workers with at least one and at most 18 observations per worker. The total data set consists of 147 379 observations.

All data are based on employers' reports[5] and they contain many work-related issues. For example, in each fourth quarter hours of work and adherent earnings are categorised into time, piece rate, holiday, and over-time items. (The fourth quarter is selected since the negotiated wage increase during a particular year is more likely to be properly distributed at this time of the year.) A potential problem with this data set, besides that it lacks information on tax-related issues, is that gender and age are the only individual-specific characteristics. For several reasons, though, this is not necessarily a serious problem in our study. First, the population of workers from which our sample stems is probably homogeneous in terms of certain characteristics, e.g., family background and education. Second, since we concentrate on male workers, individual characteristics may be less important. Children, for example, are often found to influence female labour supply, but seem to have little impact on male labour supply (see e.g. Blomquist and Hansson-Brusewitz, 1992).

In addition to selecting men, we have made the following restrictions: the hourly wage in any year is limited to the interval between the minimum wage stated in collective agreements and 200 Swedish kronor, and hours of work in the fourth quarter are limited to 1 000. These restrictions reduced the number of observations to 49 635.

Hours of work are constructed by adding time, piece-rate, holiday, and overtime hours during the fourth quarter in each year. Figure 1a shows that hours of work in our sample follow the pattern found in the engineering industry as a whole; cf. Figure 1b. The pronounced peaks

[5] There are reasons to believe that data collected by the employer are of higher quality than data based on workers' own reports; for a discussion see e.g. Duncan and Hill (1985) and Laisney *et al.* (1990).

Figure 1a. Quarterly hours worked per worker in our sample, 1970–1987

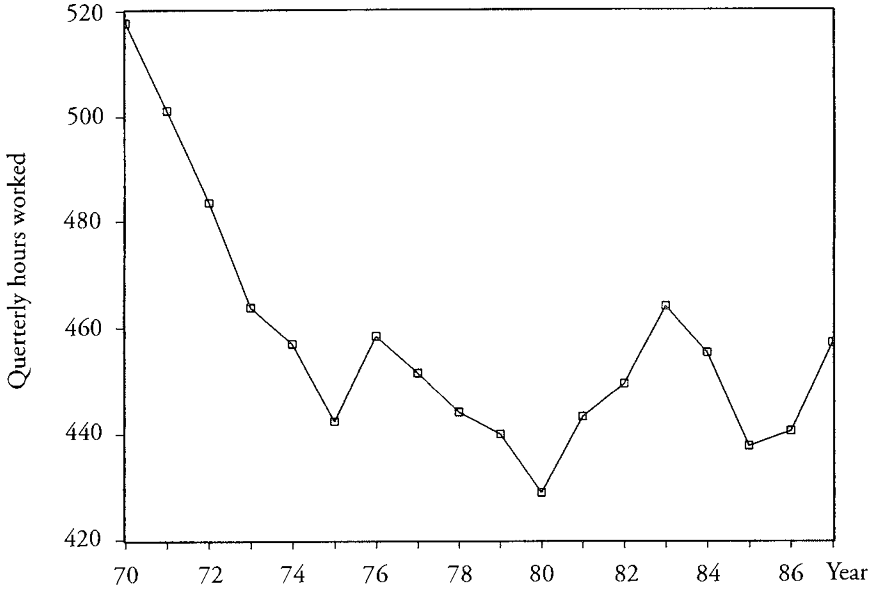
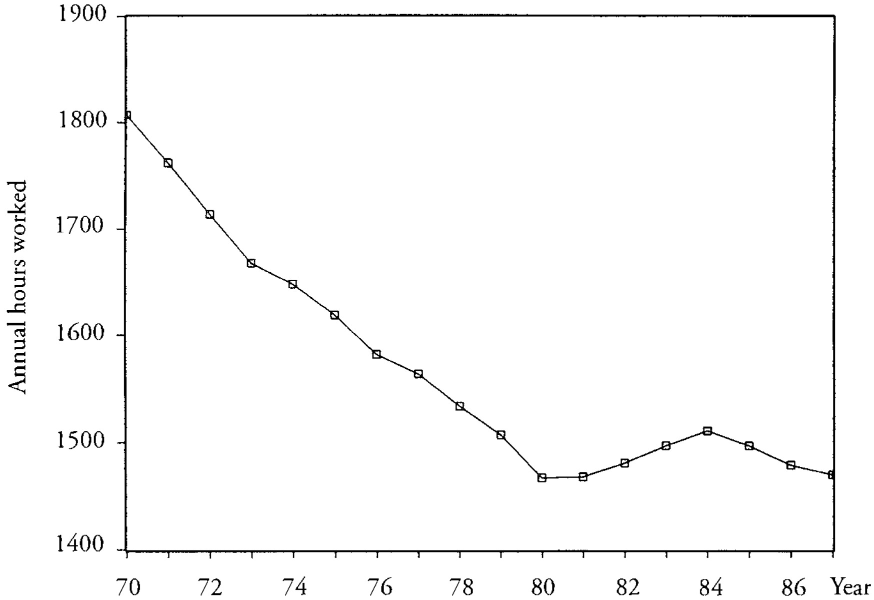


Figure 1b. Annual hours worked per worker in the engineering industry, 1970–1987



Source: SOS Industry (Sveriges officiella statistik – industri).

and troughs in Figure 1a, as compared with Figure 1b, can be explained by Christmas and New Year holidays; i.e., peaks (troughs) occur in data when these holidays coincide with a weekend (work week). Worth noting in Figure 1a is the large decline in hours worked during the 1970s, from 517 in 1970 to 428 in 1980.[6]

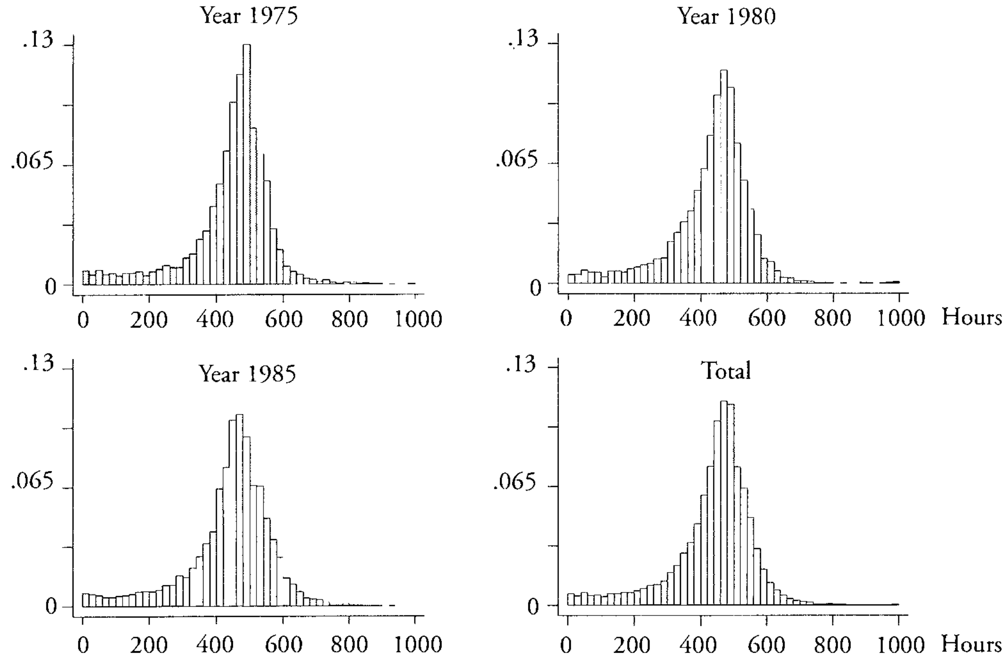
The estimations rest on the assumption that individuals can freely vary their hours of work. The validity of this assumption has been questioned (see e.g. Andersson *et al.*, 1988, and Ham, 1982, 1986). Furthermore, Sacklén (1995) concluded that even a small number of constrained individuals may have an impact on the labour supply responses; i.e., ignoring the possibility of constraints could lead to an elasticity that may exaggerate the true labour supply responses to tax changes. An advantage of our data base, with respect to this matter, is that it covers an entire quarter: individuals can vary their hours of work by taking days or weeks off as well as varying their hours per week. Figure 2 shows that, for the years when sampling took place (i.e., in 1975, 1980 and 1985) and for the total sample, there are no major spikes in the distribution of the hours of work.[7]

The real (pre-tax) wage was obtained by dividing earnings (including time, piece rate, shift, holiday and overtime pay) by hours of work and deflating the series by the consumer price index. Figure 3a shows that there was a steady growth in the real pre-tax wage in the early 1970s and a decline in the mid-1970s and early 1980s; cf. Figure 3b. Of particular interest is that the average pre-tax wage in our sample had barely recovered to the levels prevailing in the early 1970s by the end of 1987.

[6] Part of the decline in hours of work in the early 1970s can be explained by a change in the legislated work week: it was reduced from 42 hours and 30 minutes in 1969 to 40 hours by 1973 (see e.g. Nilsson, 1992, and Tegle, 1982, for details). This has no practical meaning in our empirical analysis since we always use instruments lagged at least three periods. Also, in 1978, legislated vacation with pay was increased from four to five weeks. Since we use data from the fourth quarter, this may be less important here, than if we had observed hours of work during the third quarter.

[7] Of course, the absence of major spikes in hours of work does not guarantee that an individual can freely vary his hours of work. To investigate further whether workers feel constrained on the labour market, we used data from a representative sample of Swedish households (Klevmarken and Olovsson, 1993). Among those who worked in the engineering industry (169 individuals), 4 per cent answered yes when asked if they would like to *increase* their hours of work (at unchanged wage rate) to make more money, and around 14 per cent answered yes when asked if they would like to *decrease* their labour supply, though it would mean a reduction in income. Judging from these answers, hours restrictions do not seem to be a major problem in the engineering industry.

Figure 2. Histogram for quarterly hours of work in the years when sampling took place (i.e. in 1975, 1980 and 1985) and the total sample



The Swedish income tax system consists of two parts: a local and a national tax. The local tax is proportional with a fixed exemption level while the national tax is progressive. During the sample period taxable income was the sum of labour and capital income minus deductions (since 1991 capital income is taxed at a flat rate of 30 per cent). The most important deductions are work-related travel expenses and interest rate payments. As already mentioned, we do not observe capital income or any worker-specific deductions, so we had to take some short cuts in computing the marginal tax rate; this computation is discussed in Appendix 2. Table 1 presents the constructed marginal tax rates on labour income. Notice the general increase in the tax rate during the 1970s, from 49 per cent in 1970 to 60 per cent in 1978. After that it declines, and by the end of the sample period it averages around 50 per cent.

The real after-tax wage is constructed as the real wage multiplied by one minus the marginal tax rate. Figure 4 suggests that the tax system has

Figure 3a. Hourly real wage (pre-tax) per worker in our sample, 1970–1987

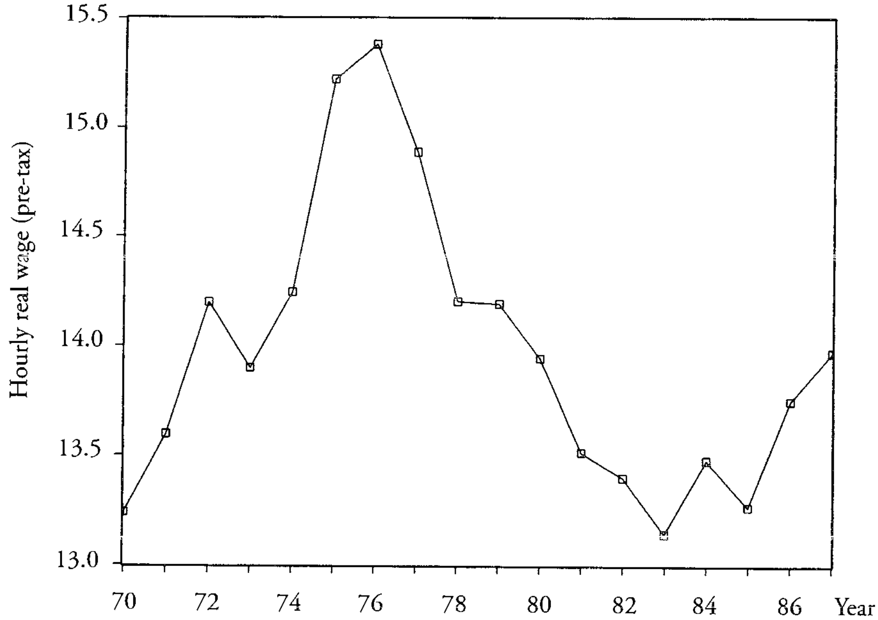
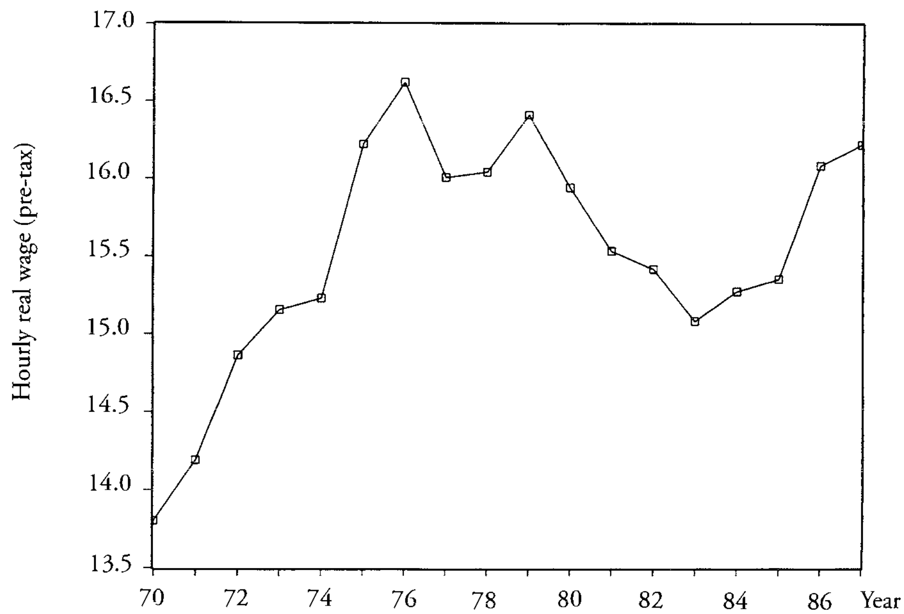


Figure 3b. Hourly real wage (pre-tax) per worker in the engineering industry, 1970–87

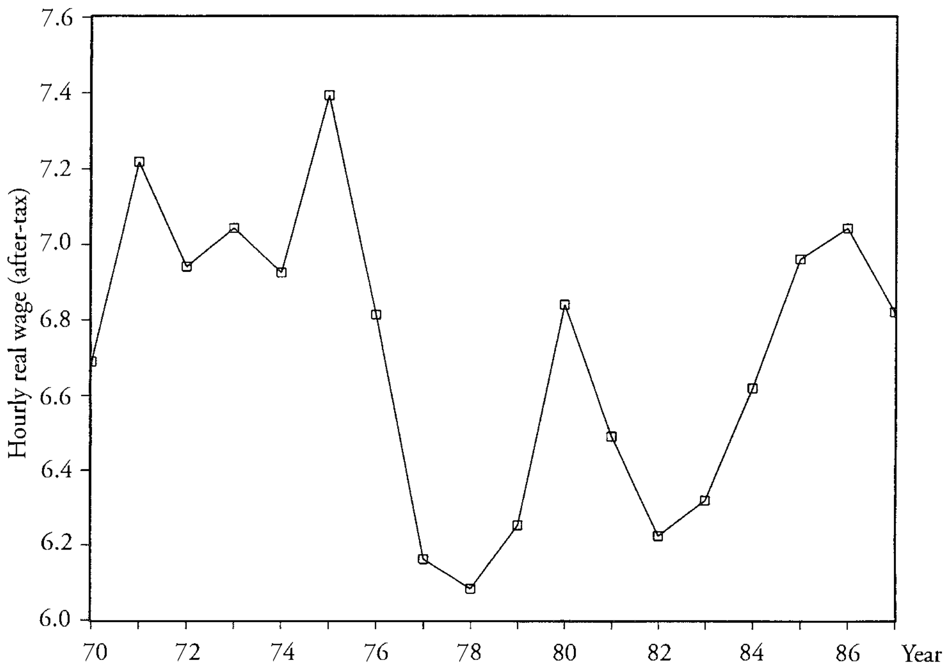


Source: SOS Industry (Sveriges officiella statistik – industri).

Table 1. Computed marginal tax rates from our data

Year	Mean	St. dev.	Min	Max
1970	.49	.09	.0	.60
1971	.46	.09	.0	.60
1972	.52	.10	.0	.68
1973	.51	.10	.0	.62
1974	.54	.12	.0	.62
1975	.53	.10	.0	.77
1976	.58	.09	.0	.75
1977	.61	.08	.29	.77
1978	.60	.08	.30	.82
1979	.58	.09	.30	.81
1980	.54	.08	.29	.78
1981	.54	.08	.30	.79
1982	.56	.08	.30	.79
1983	.54	.08	.30	.75
1984	.53	.06	.30	.71
1985	.49	.06	.32	.75
1986	.50	.05	.30	.71
1987	.52	.06	.34	.76

Figure 4. Hourly real wage (after-tax) per worker in our sample, 1970–1987



kept the purchasing power of the workers in our sample in a more narrow range compared with pre-tax wage fluctuations (cf. Figure 3a). The pronounced increase in the pre-tax wage in the early 1970s is not preserved after tax, and by the mid-1980s the after-tax wage is almost at the same level as in the mid-1970s.

4. Results

In Table 2 we present the results from estimating the life-cycle model using the third, fourth and fifth lag of earnings as instruments for the wage rate. We also included three dummies (not shown) to account for the varying size of the quarter due to the precise position of the Christmas holiday. The wage growth measure used here is the predicted nominal growth rate in wages from one year to the next minus the nominal annual interest rate, as implied by the theoretical model. Two specification tests are shown at the bottom of the table. The second-order serial correlation test[8] is easily accepted in all specifications, which suggests that the timing of the instruments used for predicting wage growth is probably valid. In particular it implies that there seems to be no serious serial correlation due to measurement errors in the tax rates or the gross wages. Nevertheless the other misspecification test statistic – the Sargan overidentifying restriction test – is somewhat high, possibly suggesting some misspecification. This problem motivates the use of the grouping estimator as a device to test the results in the model (more on this below).

The results in the basic model are revealing. They imply a very small intertemporal elasticity of substitution that is quite precisely estimated. The 95% confidence interval (from –0.028 to +0.124) implies that a 10 per cent anticipated increase in the wage rate would lead to average hours rising by at most 6 hours in a whole quarter. An unanticipated effect would be even smaller since there will be offsetting effects from a perceived increase in life-cycle wealth. This implies that the tax reform is unlikely to have large effects on hours of work via changes in labour supply of men.

How does this result compare with other Swedish studies? Flood (1994) provides an overview of earlier Swedish labour supply studies. From Table 1 in that study (concerning men only) we can see that the

[8] The second-order serial correlation test is an N(0,1) test statistic for the null hypothesis that the differenced model has a MA(2) error (against the hypothesis that it is MA(3)).

Table 2. Life-cycle model. Dependent variable: growth rate of hours. Whole sample (n = 49 635)

Δ(wage measure)	0.048 (0.038)	0.021 (0.043)	–0.036 (0.080)
Δ(wage measure)*High past earnings			0.11 (0.10)
Age	0.0009 (0.002)	0.0007 (0.002)	0.0006 (0.002)
Age^2	–0.0026 (0.0016)	–0.0021 (0.0017)	–0.0020 (0.0016)
Change in unemployment		0.0037 (0.0056)	
Second–order serial correlation	0.32	0.33	0.38
Sargan Test (degrees of freedom)	42 (12)	48 (13)	79 (25)

Note: Δ (wage measure) = Δ ($\ln w_{it} - r_{t-1}$). Standard errors in parentheses.

uncompensated wage elasticity ranges from 0.04 to 0.21 and the compensated wage elasticity ranges from 0.05 to 0.28. Thus, our result appears to square well with previous Swedish studies.

Before reaching a final conclusion we conducted several other experiments. First we tested whether the change in the unemployment rate can explain the variations in hours worked; see Ham (1986). If this is the case, this would constitute evidence against the model and invalidate our earlier conclusion. In the second column of Table 2 we include the change in the regional unemployment rate as a regressor and instrument it using its second and third lags. This variable is insignificant.

Intertemporal substitution in hours of work relies on the ability of the individual to borrow in periods where wage rates and hence the gains from working are low. In such periods the individual works less and preserves the required standard of living by topping up income from savings or borrowing. Thus, as wages grow, the individual transits from a time of lower hours of work to working more. If the ability to borrow of future earnings is limited, the willingness to substitute intertemporally will be reduced for those individuals with limited savings. These individuals will not allow hours to fluctuate much in response to wage changes since earnings is their only source of consumption. To test whether the presence of liquidity constraints for some individuals obscures the presence of important incentive effects for the rest, we defined a dummy variable which is one when past earnings were above the median three periods back; otherwise it is zero. We then interacted this dummy with the wage growth variable. The idea is that individuals with high past earnings are less likely to be liquidity constrained and therefore behave more in line with the life-cycle model (see Zeldes, 1989). The results presented in the third column do in fact suggest a higher point estimate for the elasticity of individuals with high earnings in the past. The elasticity is 0.074, (i.e. 0.11–0.036) which is consistent with the liquidity constraint story. Nevertheless, the elasticity is still very low. Moreover, the additional variable is insignificant; this experiment cannot reject the implications of the life-cycle model. Thus, based on this evidence, the changes in wages induced by policy reforms or otherwise, do not seem likely to have elicited a labour supply increase of any significance.

We now consider an experiment using an alternative wage measure. Here we regressed the growth rate of hours on the growth rate of real wages (i.e., nominal wage growth minus consumer price inflation). To interpret these results consider a conventional within period labour supply model:

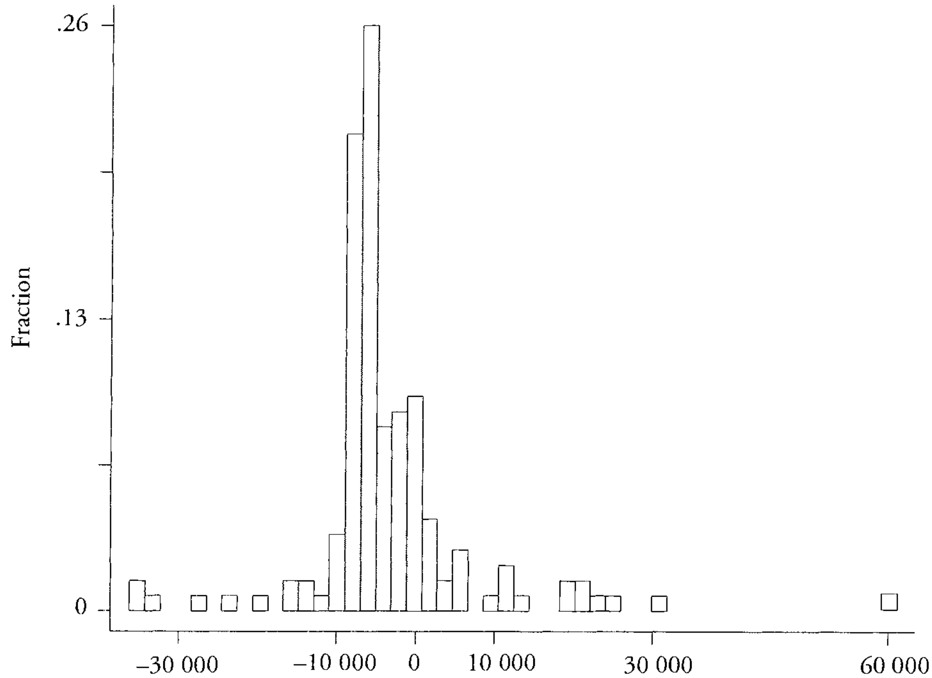
$$\ln h_{it} = f_i + \beta \ln (w_{it} / p_t) + \gamma y_{it} + v_{it} \tag{4}$$

where p_t is the price index of consumption goods, y_{it} denotes unearned income, f_i represents a fixed taste component and v_{it} a temporal shock to hours. Now suppose that the difference in unearned income (Δy_{it}) is small or is not correlated to lagged earnings. In fact, looking at two waves of Household Market and Non-Market Activities (HUS) data (1983 and 1985), we see that the unearned income measure[9] does not change much, though it is slightly skewed to the left (Figure 5).

Alternatively suppose that the income effect (γ) is zero. Here regressing the growth rate in hours on the growth rate of the real wage should provide an estimate of β. As documented by Flood (1994) the income effects found for men in Sweden are very small. Moreover, the two wage growth series are *ex ante* quite different, the correlation coefficient between the two being 0.76. Thus, to the extent that the predicted value of Δy_{it} does not vary much, it is possible to interpret the results of the regression

[9] Unearned income comprises taxable and non-taxable non-labour income (for males who are cohabiting, unearned income also includes spouse's income); HUS data are described in Klevmarken and Olovsson (1993). A better measure of y_{it} would be non-labour income minus savings. This is the only measure consistent with the life-cycle model (see Blundell and Walker, 1986).

Figure 5. Changes in real unearned income between 1983 and 1985



Source: Unearned income is computed from HUS data (Klevmarken and Olovsson, 1993).

$$\Delta \ln h_{it} = \beta \Delta \ln (w_{it}/p_t) + \Delta v_{it} \tag{5}$$

as capturing the conventional within period effect. Without liquidity constraints and if income effects (γ) are zero we expect the within period effect (β) to equal the intertemporal substitution one (α). If the income difference is simply uncorrelated with our instruments used in predicting wage growth, but γ is negative then we expect $\beta<\alpha$. In the presence of liquidity constraints however, even if all other conditions are valid, the relationship between the two coefficients is ambiguous.

The results are presented in Table 3. Using the same instrument set as before (which includes the nominal interest rate and the inflation rate lagged three periods), we obtain a wage elasticity of 0.143 which is substantially larger than the one obtained with the other wage measure. It

Table 3. Life-cycle model. Dependent variable: growth rate of hours

	Whole sample		Balanced sample
Δ(wage measure)	0.143 (0.067)	0.124 (0.070)	0.045 (0.099)
Age	0.0009 (0.0021)	0.0011 (0.0021)	–0.0037 (0.0031)
Age ^{2}	–0.0027 (0.0017)	–0.0027 (0.0017)	0.0023 (0.0030)
Change in unemployment		0.0032 (0.0053)	
Second-order serial correlation	0.27	0.28	–1.40
Sargan Test (degrees of freedom)	37 (12)	48 (13)	40 (12)
Sample size	49 635	49 635	11 921

Note: Δ (wage measure) = $\Delta \ln(w_{it}/p_t)$. Standard errors in parentheses.

still implies a small effect: a 10 per cent rise in the wage rate would induce a labour supply response of 7 hours a quarter on average.[10] In the second column we show that the change in the unemployment rate has no significant effect here either.

Out of all the individuals in the sample, only about a quarter are observed continuously for the 18-year period. Therefore the results presented could be affected by selection or attrition bias. We can test the hypothesis of no selection bias by comparing the results from the balanced panel and those from the entire sample used up to now. The estimates from the balanced panel, where only the individuals observed for all 18 years are kept, are presented in the third column of Table 3.[11] The Hausman test that the wage coefficient between the balanced panel and the whole data differs is 1.34. Hence, the difference is not significant. Finally, when we included a quadratic time trend (not shown) in the first-difference equation, the wage elasticity became negative (–0.022), which at a standard error of 0.075 is not significantly different from zero or from our previous results.

Our final experiment consisted of estimating equation (5) using grouped cohort data. We constructed a panel of grouped observations where the groups are defined by the date of birth of the individual. Table 4 presents the groups used and the number of individuals in each

[10] The implications were not substantially different when we used the change in hours on the left hand side, rather than the growth rate of hours.

[11] We are able to follow 917 workers over the entire period; i.e., there are 11 921 observations in the balanced panel after lagging and first differencing the data.

Table 4. Structure of the grouped data

Cohort	Year of birth	Average cell size
1	1928–1932	209
2	1933–1937	175
3	1938–1942	206
4	1943–1947	231
5	1948–	96

cell.[12] We also used methods for predicting wage growth that are suitable in the presence of macroeconomic shocks and long periods of observations. The basic idea underlying this estimator is that it is more robust to measurement error since all additive idiosyncratic components average out with aggregation procedure. The elasticity implied by the grouping procedure (not shown) is very close to the point estimate obtained by the standard instrumental variable procedure: we obtain an elasticity of 0.11 but now with a much larger standard error (0.25). The loss in precision is to be expected with this type of estimator which disposes of most of the cross-section variation in the data. In this case the model is identified more or less exclusively from the time-series variation. Despite this the results we obtain are in line with those we obtain when we use the individual data. This implies that our results are unlikely to be seriously biased from measurement error or from unaccounted macroeconomic shocks.

5. Discussion

We have estimated models of labour supply in first differences using two alternative wage growth measures: one based on the life-cycle model and one based on the within period model of labour supply. The interpretation of the results in the former case relies on the absence of binding liquidity constraints while in the latter it relies on predicted changes in other income not displaying much variation. We also experimented with using just the balanced panel, with including the changes in unemployment rates and time trends in the equations and with using grouped data.

[12] To construct our quarterly cohort data we divided each year of data by the individuals' year of birth into five-year bands, starting in 1928. All the relevant variables were then averaged over the fourth quarter within the cohort. This gives a balanced panel of five cohorts and 18 quarterly observations. The total number of grouped observations is 90, which corresponds to the 16 506 cross-section data points.

All experiments yield low labour supply elasticities with quite small standard errors. These are in line with the figures obtained in a number of other studies of Swedish male labour supply.

Our contribution has been that we exploit the relatively long time-series dimension of our panel and show that we cannot explain the fluctuations in hours of work on the basis of labour supply responses to wage changes. Based on these results our conclusion has to be that the incentive effects of the tax reform on hours worked are not substantial, at least not for men working in the engineering sector. This, of course, may not reflect preferences; there may be important constraints in working hours due to the organisation of production lines in this sector. It may well be that individuals who wish to change their hours move between sectors which, of course, is not captured by our data. Modelling this aspect of behavioural responses requires data covering more than one sector. In addition there may be important participation effects as well as reduced wage pressure with obvious positive effects on economic activity. None of these issues are addressed in this paper.

Finally, a small labour supply elasticity does not necessarily imply that the excess burden of the old tax system were low. The reason is that the marginal tax wedges used to be so high that even a small elasticity can be translated into a large welfare loss.

Appendix I.

Specification of the model

The basic framework we use to interpret our data is the standard life-cycle model of labour supply with uncertainty (see e.g. MaCurdy, 1981 and 1984). We assume that the utility function in period t has the form

$$U_t(h_t, c_t) = \frac{h_{it}^{\frac{1}{\alpha}+1}}{\exp\left(\beta' g_{it}^{\frac{1}{\alpha}+1}\right)} + f(c_t) \qquad \text{(A.1)}$$

where h are hours of work and c is consumption. The Frisch labour supply function corresponding to this utility function is

$$\ln h_{it} = \alpha \ln w_{it} + \beta' g_{it} + \alpha \ln \lambda_{it}. \qquad \text{(A.2)}$$

Here $\ln h_{it}$ is the log of hours of work, $\ln w_{it}$ is the log of the nominal after-tax hourly wage and g_{it} is a function of observable and unobservable individual characteristics; random preferences, including fixed effects, can be introduced by a suitable specification of the g_{it} function. Finally, $\ln \lambda_{it}$ is the log of the marginal utility of wealth. Equation (A.2) assumes additive separability of hours of work from all other goods. Although we do not observe quantities consumed of other goods, this assumption can be relaxed simply by including (relative) prices of goods from which hours of work are not separable. Note also that an advantage of this specification is that it can allow for hours to be zero when wages are low. However, this is not an issue in our study since we only observe employed workers. Moreover, it should be kept in mind that the analysis concerns men only.

Given rational expectations the log of the marginal utility of wealth has the property

$$E_t\,[\ln \lambda_{it+1}] = \ln \lambda_{it} + \delta_t - r_t + d_{it} \tag{A.3}$$

In (A.3) d_{it} is the difference between $E_t[\ln \lambda_{it}]$ and $\ln[E_t \lambda_{it}]$. Under joint normality this is proportional to the variance of the marginal utility of wealth conditional on period t information. This log-linear representation of the martingale property of the marginal utility of wealth is convenient since it avoids the need for non-linear estimation. Finally, δ_t is the rate of time preference and r_t the nominal interest rate.

The life-cycle model is particularly suitable for the type of data available, with limited information on individual characteristics. The marginal utility of wealth controls implicitly for wealth (as viewed in period t). In addition, the first differencing procedure used to eliminate λ_{it} by exploiting its martingale property will also eliminate all individual characteristics which are fixed over time (and have a fixed effect on labour supply). This leads to equation (1) in the text which forms the basis of our estimation method.

Appendix 2.

Computation of the marginal tax rate

Given the available data we had to make some simplifying assumptions in computing our marginal tax rate. To that end we treated labour income as the only source of taxable income[13] and we only considered general deductions.

For the period 1970–1976 the computations were simple since we knew the marginal tax rate of each taxable income segment. First, we annualised labour income by multiplying earnings from the fourth quarter by four. Second, we added 100 to labour income and identified the marginal tax rate pertaining to this amount.

For the remaining years (1977–1987) we only knew the income tax related to each taxable income segment. Here, we first used the calculated labour income to identify the national and local income tax rates. The local income tax was found by multiplying the local marginal tax rate (a flat rate in each municipality) by labour income. Second, we added SEK 100 to the calculated labour income and recalculated the income tax. Third, the marginal tax rate was obtained by taking the difference of the two calculated income tax amounts and dividing it by 100.

As discussed in Section 2, the calculation of the observed (our) marginal tax rate might cause measurement error. The direction of this error is *a priori* undetermined since capital income and deductions have opposite effects on taxable income. To get an estimate of the error we computed both the actual and the observed marginal tax rate on HUS data (Klevmarken and Olovsson, 1993). The average actual and observed marginal tax rates in 1983 were 0.50 and 0.48 (see Table A.1). It seems as if the observed tax rate underestimates the actual one. However, the difference between them is rather small. A plausible explanation for this could be that capital income and deductions net out.

[13] To get an idea of the size of this approximation, we compared labour and taxable income on a data set that contains both measures. (We used HUS data (Klevmarken and Olovsson, 1993) to conduct this comparison, and we selected male workers in the engineering industry, aged 20–60.) Labour income seems to overestimate taxable income by less than 3 per cent – the average labour income was 78 176 Swedish kronor (st. dev. 21 119) while taxable income was 76 192 Swedish kronor (st. dev. 24 899).

Table A.1 Detailed statistics of the actual and observed (our) marginal tax rates in 1983

Percentiles	Actual	Observed
1%	0.0000	0.3235
5%	0.3372	0.3471
10%	0.3571	0.3631
25%	0.4169	0.3933
50%	0.5280	0.4880
75%	0.5657	0.5371
90%	0.6194	0.5750
95%	0.6559	0.6281
99%	0.7619	0.7180

Source: The "actual" marginal tax rate is computed from the HUS data (Klevmarken and Olovsson, 1993).

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