Time Out for Childcare: Signalling and Earnings Rebound Effects for Men and Women

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Abstract. The wage cost of time out of the labor force for childcare is important in order to understand the functioning of labor markets and for public policy. This paper reviews the literature and identifies several limitations. Using employment records of a large Swedish company over the period 1983–88, we demonstrate an alternative approach for estimating earnings effects and find a year out costs 1.7 percent of earnings for a woman and 5.2 percent for a man. This large effect for men raises questions of signalling costs. For both men and women, earnings "rebound" for time out in the more distant past.

Introduction

What is the career earnings cost of time out of the labor force for the care of young children? This question has continued to receive much attention given the growth of dual earner couples in the industrialized countries. The effect of time out of the labor force, particularly in the early part of one's career, is part of the more

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general question of work history or on-the-job training effects on earnings. Labor force exit for childcare takes place during the early or high investment phase of the life cycle. Do such exits have long lasting effects or do earnings rebound after a short period of re-entry to full time work? Are there signalling costs, in the sense that, especially for men, employers may judge work commitment by length of time away?

Research on these and related questions, starting with the early use of the National Longitudinal Surveys (Mincer and Polachek, 1974) and the Panel Study of Income Dynamics (Hill, 1979; Corcoran and Duncan, 1979), has increasingly turned to panel data and estimation procedures to account for the endogeneity of spells out of the labor force (Gronau, 1988; Korenman and Neumark, 1992). This more recent literature has confirmed the earlier finding (Hill, 1979) that marriage and number of children have an impact on earnings of women *primarily* via the extent to which they shape labor force attachment and work interruptions. In a review of the existing studies we show that beyond agreement that "work history matters," there are widely disparate results: the impact of time out varies widely, for women from less than one percent to eight percent per year; earnings recovery or 'rebound" is sometimes present but not found consistently; and the effect of time out on men's earnings is not estimated. The reason for these diverse findings, we believe, is a lack of accurate work history spell information and earnings change measures.

The purpose of our paper is to examine the earnings effects of spells away from work or "time out" to care for young children. We accomplish this by using company panel data for a large multibranch company in Sweden over the period 1983–88.[1] In Sweden childcare leave is paid by a national parental leave policy with no direct costs to employers of parents who use it (Sundström, 1993), and this raises additional questions of the financial incentives for employees to utilize the leave. On the employer's side, however, there should be normal incentives to evaluate the effect of leave use on productivity and therefore base pay on leave history insofar as leave from work influences productivity.

The company records in our study provide extremely accurate measures of the history of each spell (down to those less than a single workday!), which is important from an errors-in-variables perspective, and, as well, extremely accurate earnings information, which is important in the estimation of fixed-effect, wage-change equations. The detailed company records permit the construction of disaggregated measures of the timing and nature of the employee's child leave histories for men and women. With these histories one can address questions such as the long versus short run impact of time away from work and whether long spells are more costly than a series of short exits with the same total time away from work.

Despite the more equal sharing of household reponsibilities by men and women in Sweden (Juster and Stafford, 1991), an issue is the rather limited use of childcare leave by men. Several surveys have found that men express reluctance to take leave because their employer may infer this to indicate a weak work commitment. If so, a signalling equilibrium could occur where few men take leave in order to avoid being labelled as uncommitted to their work. Alternatively, men may be in jobs with a greater payoff to work experience, so that time off the job may simply have a greater human capital cost.

The organization of this paper as as follows: In Section 1 we review the theory and related empirical literature on the relationship between time away from work for childcare and the impact on earnings. Section 2 provides an introduction to the special features of our dataset. Section 3 includes several parts: Part 3.1 presents: (i) the basic estimation of the fixed-effect model and addresses the question of whether a random effect model could be applied and (ii) the estimated cost of time away from work for men and women. Part 3.2 presents a disaggregation of time out into leave of different types and dates to examine several questions including the robustness of the model and "earnings rebound".[2] Part 3.3 examines the question of endogeneity of the spells. A brief conclusion is offered in Section 4.

1. The theory of lifetime earnings and existing research

1.1. Human capital and signalling effects

Why should work history matter? The most widely accepted view is that people acquire work-related skills on the job. For reasons of recouping investment costs as well as sustaining a stock of marketable skills in the face of depreciation, the expectation is that early life-cycle investments will be greater (Ryder, Stafford and Stephan, 1976; Blinder and Weiss, 1976). The same theory suggests that earlier in the lifecycle an exogenous drop in human capital would be "made up" as the person reapproaches the steady state.[3]

In this view of earnings it is skills and productivity that matter, and employers are assumed to be able to observe productivity.

Suppose productivity is only partly observable (Medoff and Abraham, 1981) and employers must rely on indicators of productivity, effort, and career commitment (Spence, 1973). Taking time out may be such an indicator. We believe that there is evidence suggesting that spells away from work for childcare correlate positively with skills and motivation fo women and negatively with skills of men. If this is the case, women may take leave, in part, to establish a *positive* signal of productivity while men may seek to avoid the "stigma" or *negative* signal of their productivity.[4] For this reason caution needs to be exercised in interpeting time out simply as a measure of reduced market productivity. In fact our results suggest that time away from work may have a different meaning for men and women.

The notion of adverse signalling effects of benefit receipt has been offered in several areas such as unemployment payments, participation in training programs, use of targeted wage subsidies (Burtless, 1985), and welfare use (Moffitt, 1983). In the case of parental leave use we are able to rely on a variety of other studies of the reasons given by men and their spouses for the rather scant use of the system by men. These studies provide more direct confirmation of men's aversion to use leave because of pressure from employers, co-workers and supervisors (see e.g. Haas, 1992).

1.2. Existing literature

Most of the existing literature on work interruptions has approached the subject by trying to explain the male-female wage gap and especially why the correlation between marriage and earnings is positive for men but negative for women. The expectation is that marriage has these divergent effects on the earnings of men and women because children are associated with a less consistent work history for women but a more consistent work history for men; in addition there is a presumption that more able men (via long run income effects, if marriage is a normal good) are more likely to marry. The implication is that better measures of ability and work history will show that it is not marriage *per se* that matters for earnings, but the way in which marriage and family relate to work history and consequent market skills of men and women. For this reason much of this research has an emphasis of the role of work history *vis-à-vis* marriage and children.

Table 1. Studies of wage effects of time out of the labor force

Author(s)	Data source	Men/Women (M/W)	Type[a] of equation	SE[b]	IV[c]	FE[d]	Results
Mincer and Polachek (1974)	NLS 1967	W	P/XS	No	No	No	1.5% per year
Corcoran and Duncan (1979)	PSID 1976	M/W	P/XS	No	No	No	0.5% per year
Hill (1979)[e]	PSID 1976	M/W	P/XS	No	No	No	No direct marriage or children effect
Gustafsson (1981)	SAF white collar pvt. sector 1974	M/W	P/XS	No	No	No	1.0% for 30–44 year olds
Mincer and Ofek (1982)	NLS mature women,	W	P/XS	No	No	No	1.5–1.8% per year ("rebound")
"	1966–74	W	Chg.	No		Yes	5.9–8.9% per year
Corcoran, Duncan and Ponza (1983)	PSID white women,	W	P/XS	No	No	No	1.5% ("rebound")
"	1967–79	W	Chg.	No	No	Yes	3%
Dolton and Makepeace (1987)[E]	Grads. in UK	M/W	P/XS/Chg.	No	No	Yes	No marriage effect but children eff. Pos. eff. of children for males
Gronau (1988)	PSID 1976–79	M/W	P/XS	Yes	No	No	3.0–5.9% for women. Not significant for men
Nakosteen and Zimmer (1987)[e]	PSID 1977	M	P/XS/Chg.	No	No	Yes	No support for marriage selection effect
Groot, Schippers, Siegers (1988)	CBS/IVA 1982	W	XS	No	No	No	1–5% per year, greater for those w/less education
Korenman and Neumark (1992)[e]	NLS YS 1982	M	Xs/Chg.	Yes	Yes	Yes	Negative effect of children reduced with controls for work experience
Kim and Polachek (1994)	PSID 1976–87	M/W	XS				0.2–0.8% per year for men; 1.2% per year for women
"			/Chg.	Yes	Yes	Yes	0.5–2.8% per year for men; 1.2–5.4% per year for women

[a] Panel, cross-section or change. [b] Simultaneous equation. [c] Fixed effect. [d] Instrumental variables. [e] No direct estimate of the effect of time out on earnings.

Table 1 presents a summary of findings from a selection of studies examining the effect of time out of the labor force for childcare on earnings of men and women. As can be seen the literature interconnects with that on the effects of marriage on the earnings of men and women. The studies commonly use panel data for the purpose of constructing work history variables (e.g. Mincer and Polachek, 1974; Corcoran and Duncan, 1979) and less commonly use panel data to create an earnings change measure (Mincer and Ofek, 1982; Corcoran *et al.* 1983). The impact on (younger) women's earnings of one year out of the labor force (often presumed to be for care of young children) is estimated to be in the range of 0.5 percent (Corcoran and Duncan, 1979) to 5.0 percent (for women under age 34) (Groot, Schippers and Siegers, 1988).The effects of marriage on women's earnings are reduced or fully disappear when work history variables are added to the earnings equation (Hill, 1979; Dolton and Makepeace, 1987; Nakosteen and Zimmer, 1987; Korenman and Neumark, 1992).

There is some discussion of long versus short term wage impacts of time out. Only a couple of studies specifically attempt such an estimate (Mincer and Ofek, 1982; Corcoran, Duncan and Ponza, 1983). However, the basis for the inference of both these studies that short-term effects are, as predicted by theory, more pronounced (5–8 percent and three percent, respectively) than long-term effects (1.5–1.8 percent), is a comparison of a cross-sectional ("long-run") wage equation with a wage change ("short run") equation. This inference is limited by the fact that their change equation is really a type of fixed effects model, and most of the difference with the cross-sectional may be simply the consequence of different estimation procedures.

Another pattern observed in some studies is the small (Corcoran and Duncan, 1979) or even positive effect (Gustafsson, 1981, for those other than age 30–44 years old) of time out. With a cross-sectional estimate, it is possible that the more ambitious and productive women take more time out, particularly in US data on families before 1975. The early Postwar US pattern was for more educated, higher wage women to work more except when preschool children were present, when hours of market work were less than for those with lower levels of education (Bowen and Fingegan, 1969). By extension, women with higher (unobserved) wage potential may take more leave. Women in the US return to work very soon after a child is born (Leibowtiz, Klerman and Waite, 1992) and work much longer hours than women in other countries

(Gustafsson and Stafford, 1994) suggesting a high career cost to time out in the US. If leave is positively correlated with unobserved productivity the actual cost of time out would be larger than suggested by cross-sectional estimates and may explain the difference between the cross sectional and panel or fixed effect results reported in Mincer and Ofek (1982) and Corcoran, Duncan and Ponza (1983).

Some more recent papers address the question of whether poor labor market prospects, possibly from discrimination, discourage participation, so that time out (Gronau, 1988; Kim and Polachek, 1994) or fertility status (Korenman and Neumark, 1992) is endogenous, with a resulting bias toward finding an effect of time out (or children) when in fact none exists (or the effect is small). While these studies are suggestive of discouragement effects, as Gronau notes in such a "simultaneous-equation scheme it is hard to justify the choice of labor force experience, tenure and occupational choice as exogenous variables" (p. 295).

All of the studies in Table 1 have serious data limitations. Hometime is often aggregated temporally into an overall total of years out of the labor force and measured as the residual of age (−5) less years of schooling and work experience (e.g. Kim and Pollachek). Even when separate segments of time out of the labor force are available, they are likely to be subject to substantial recall error given that respondents commonly report on events spanning at least a calendar year. To estimate an earnings change/fixed-effects equation, accurate data on earnings are required or else much of the apparent change will be dominated by measurement error. Our company data have only a small number of variables for the employee and none for the spouse or children, but have as their strength very accurate histories of parental leave and very accurate earnings data.

2. Time out: spells of childcare leave from company data

We utilize two samples. One is a random one in 15 sample of 2,200 out of all Swedish National Telephone Company employees in 1983 and the other is of all employees who had any parental leave in 1983. Both samples have complete leave records over the period of observation, January 1, 1983 to December 31, 1987. Upper management is excluded from both samples. The time period covered was one of tight labor markets, so that the diverse regional offices are assumed to have faced competitive labor markets, and the telephone company jobs are ones with substantial on-the-job training, as evidenced by the presence of company sponsored training programs.

In neither sample can we observe directly whether the employee has had a child or when it arrived. What has been recorded is the use of various paid leave options; leave for care of the newborn (a maximum of 360 days) and occasional care of children (a maximum of 60 days per year and child) as well as unpaid leave for childcare (see Sundström and Stafford, 1992 for a closer presentation).[5] In each of the years, 1983–87, about 30 percent of the men and 38 percent of the women had some days of parental leave. Over the whole 5-year period 43 percent of the men and 45 percent of the women had used some parental benefits.

Among women and men of childbearing ages the proportion of leave users is much higher: 82 percent of all women aged 30–44 in 1983 took at least some days of leave as did 74 percent of the men in these age ranges. But women use many more days than men; when measured in full-time equivalents women used about 130 days per year, on average, compared to only about 13 days per year for men who used leave benefits. The high number of average days used over the five year period, especially by female users, indicates the program really is important and has the potential to shape lifetime decisions. Among female users, almost one and one-half of five years were full-time equivalent leave.

In *Table 2* the usage of parental leave for newborn children and other types of time out (occasional care and unpaid leave) for the random, one in 15 samples are presented. Here we analyze total days of leave rather than spell histories (which are discussed in Section 3.3). By far the major share of leave benefits for care of newborns are used by women, with about 31 percent using such leave during the period 1983–87 and the days per user averaging 226. The corresponding figures for men ar about 15 percent and 29 days, respectively. On an annual basis the company data seem in line with averages for the country. In 1987 men used 7.5 percent of such leave days (National Insurance Board, 1992), while for our sample men used 5.4 percent of the leave days for newborns (5.7 percent over the period 1983–87).

Leave days for occasional care of children are used to a more equal extent by men and women, and this is partly because there is a set of benefit days which are only available for men's use ("daddy days") as part of a policy to encourage greater use by men.

However, even netting out these days, nationally, fathers used 35 percent of the occasional leave days in 1987 (National Insurance Board, 1992, p. 28). Unpaid leave is almost exclusively used on a part-time basis (i.e. in the form of reduced hours) and is far more likely to be used by women (averaging 157 days over the five year period) than men (averaging only three days over the five year period) for our employee sample. Does part-time leave allow one to maintain skills or does it have a significant earnings cost (Jones and Long, 1979)?

Table 3 presents the descriptive statistics for the one in 15 company sample in 1983 and 1988. The education, work histories (tenure with the company and potential experience outside the company) are fairly similar for men and women. The main differences are in the cumulated time out of the labor market, in the different types of leave and in the cumulated time in company sponsored training programs. Men averaged a mere 17.8 days of leave over the five-year period, while women average 238.0 days. Men's days in company sponsored training programs averaged

Table 2. Childcare leave and other time out for men and women in the Swedish Telephone Company

		Men		Women	
		Leave users, %	**Days per user**	**Leave users, %**	**Days per user**
Leave for care of newborn	1983[a]	6.2	13.5	15.8	114.9
	1984	6.6	16.3	15.0	89.8
	1987	3.9	18.7	15.3	83.0
	1983–87	14.6	28.7	30.9	226.0
Leave for occasional care	1983	28.5	7.8	29.0	7.6
	1984	26.6	7.5	28.4	7.1
	1987	24.9	7.9	28.3	7.2
	1983–87	42.3	24.7	40.7	27.6
Unpaid leave for care of children	1983	2.8	53.6	28.9	250.0
	1984	1.6	60.8	17.2	180.6
	1987	1.0	18.6	14.9	68.3
	1983–87	5.6	57.3	37.8	414.8

[a]Year refers to the year the leave began and the means are calculated over the whole leave spell, whether or not it ended in the same calendar year. Part-time leave days have been converted into full-time equivalents

almost 54 days compared to about 33 days for women over the five year period. How do these work history differences affect earnings? Is time out subject to a wage penalty? Is training time important for earnings growth of men and women?

3. Time out and earnings growth

3.1. Fixed versus random effects models

To assess the impact of family leave on male and female wage growth we employ a simple human capital wage equation (Mincer, 1974; Blinder, 1976). The specification is one in which the earnings variable is shifted by years of education (*SCHOOLING*) and all

Table 3. Variables and means for men and women in the one in fifteen sample, 1983 and 1988

	Women		Men	
	1983	**1988**	**1983**	**1988**
Monthly wage	8391.7[a]	9887.2	9238.6[a]	10938.0
Ln wage	9.035	9.199	9.136	9.30
Potential experience[b]	6.48	6.48	4.80	4.80
Potential experience^2	123.96	123.96	58.53	58.53
Tenure	14.06	19.06	15.62	20.62
Tenure	305.27	470.86	324.89	506.09
Potential experience × tenure	60.03	92.40	58.39	82.41
Schooling[c]	11.16	11.16	11.82	11.82
Company training days[d]	0	32.52	0	53.61
Leave days 1983–87				
Care of newborns	0	69.9	0	4.20
Occasional care	0	11.26	0	10.45
Unpaid leave	0	156.85	0	3.19
Total time out	0	238.01	0	17.84
N	886	886	1311	1311

Note: Potential experience, tenure and schooling are measured in years and company training and leave are measured in days. Since we have no information on parental leave and company training prior to 1983, these variables are 0 as of (January 1) 1983.

[a] In 1988 prices. [b] Potential experience outside of the company is defined as Age-Schooling-Tenure-7, given that Swedish school starts at age 7. [c] Normal years of study for the educational level completed. [d] Number of days in company training programs, 1983–1987. These are programs paid for by the company and compensated at the usual rate of pay.

leave days 1983–87 (*TOTAL TIME OUT*) and company training days 1983–87 (*CO-TRAINING*). Using potential experience as of 1988 (*POTEXP*88) and years of tenure with the company as of 1988 (*TENURE*88) and the interaction between tenure and potential experience (*POTEXTP*88 × *TENURE*88) as the work history variables, the earnings equation applied to our sample is

$$\begin{aligned}\ln EARN88_i = {} & b_0 + b_1 POTEXP88_i + b_2(POTEXP88_i)^2 \\ & + b_3 TENURE88_i + b_4(TENURE88_i)^2 \\ & + b_5(POTEXP88 \times TENURE88)_i + b_6 SCHOOLING_i \\ & + b_7 COTRAINING_i + b_8 TOTALTIMEOUT_i + e88_i \qquad [1]\end{aligned}$$

where $\ln EARN88_i$ is the natural logarithm of full-time equivalent monthly earnings for the ith individual.

A cross-sectional estimation of (1) has the well-known limitation of a potential and presumed omitted variable bias: here *TOTAL TIME OUT* or *CO-TRAINING* could be correlated with unobserved variables such as motivation, or even commonly measured variables but ones which were not available in the company dataset. Under these conditions the recommended procedure is to employ a fixed effects or random effects model. We can postulate that

$$e88_i = v_i + u88_i, \qquad [2]$$

with v_i time constant, person-specific "fixed" effect and $u88_i$ as an iid white noise disturbance. Then from a 1983 version of (1) we can create a change equation

$$\begin{aligned}\ln EARN88_i - \ln EARN83_i = {} & g_0 + b_1(POTEXP88_i - POTEXP83_i) \\ & + b_2(POTEXP88_i^2 - POTEXP83_i^2) \\ & + b_3(TENURE88_i - TENURE83_i) \\ & + b_4(TENURE88_i^2 - TENURE83_i^2) \\ & + b_5((POTEXP88_i \times TENURE88_i) \\ & - (POTEXP83_i \times TENURE83_i)) \\ & + b_6(SCHOOLING88_i - SCHOOLING83_i) \\ & + b_7(COTRAINING1983\text{–}87_i) \\ & + b_8(TOTALTIMEOUT1983\text{–}87_i) + (u88_i - u83_i) \qquad [3]\end{aligned}$$

where the three first terms cancel out as does schooling, since there is no change in formal schooling in our sample.

In Table 4 we present estimation results for women and men using the different estimation approaches. The regressions include only those individuals that were still employed in 1988. Since only 73 persons (3.2 percent of the sample) left the company before 1988, this is unlikely to cause any serious bias. A main finding is that the impact of total time out is to reduce earnings. In the fixed effects model for women the impact of total time out in reducing earnings is *larger* than in the cross-sectional model. Exactly the opposite is

Table 4. Wage effects of total time out for childcare, 1983–88 — one in fifteen sample (t-values in parentheses)

	Women		Men	
	Cross-section88[a]	Change 1983–88[b]	Cross-section88[a]	Change 1983–88[b]
CONSTANT	8.791	0.231	8.238	0.241
	(281.73)	(26.98)	(159.8)	(30.9)
POTEXP[c]	0.008		0.014	
	(4.67)		(3.99)	
POTEXP^2 ($\times 10^{-3}$)	−0.097		−0.069	
	(2.31)		(0.75)	
TENURE88	0.014		0.015	
	(9.09)		(6.16)	
TENURE^2 ($\times 10^{-3}$)	−0.193	−0.334	−0.144	−0.367
	(6.42)	(11.56)	(3.11)	(13.5)
POTEXP × TENURE ($\times 10^{-3}$)	−0.241	−0.351	−0.511	−0.607
	(3.63)	(5.12)	(4.16)	(7.57)
SCHOOLING	0.017		0.069	
	(7.72)		(21.16)	
COTRAINING ($\times 10^{-3}$)	0.676	0.301	−0.151	0.115
	(8.48)	(4.09)	(1.49)	(2.34)
TOTAL TIME OUT ($\times 10^{-3}$)	−0.003	−0.046	−0.190	−0.146
	(0.31)	(5.40)	(1.91)	(3.01)
R^2	0.290	0.160	0.310	0.154
N	886	886	1311	1311

[a]The dependent variable is the natural logarithm of the monthly full-time equivalent wage in 1988.
[b]The dependent variable is the difference in the natural logarithms of the monthly full-time equivalent wage in 1988 and that of 1983 (in 1988 prices).
[c]Potential non-company experience is defined as AGE-SCHOOLING-TENURE-7 given that Swedish school starts at age 7.

observed for men with the earnings reductions from total time out being *smaller* in the fixed effects model. One interpretation of these differences is that there is an unmeasured variable called "motivation" (or other unobserved factors affecting labor market success), and that more motivated women and less motivated men take time out from their careers. As a result the fixed effect model is needed to show that it really is more costly for women to take leave than would appear to be the case from cross-sectional estimates. This suggests that some of the findings of low wage penalties using cross sectional data on earnings (reported in Table 1) may have understated the cost of time out. From Table 4 time out has a cost (converted to an annual rate $365/1000 \times 0.046$ and exponentiated) of 1.7 percent for women and 5.2 percent for men. To check for the possibility of non-linearities in the effects of leave we also experimented with a squared term for leave days but it turned out neither to be significantly different from zero, nor to improve the fit of the models.

An interesting methodological point is that the predictive power of the fixed effects model is quite strong, as we would expect given company earnings records rather than respondent reports of earnings. When respondent reports are used the problem of measurement error becomes more limiting (Duncan and Stafford, 1980; Björklund, 1989), and R^2 values commonly drop dramatically to as low as a few percent or less. The reason for the low predictive power and concomitant fall in parameter precision in fixed effects models employing data from respondent reports of earnings can be seen from the methodological work on the Panel Study of Income Dynamics (Duncan and Hill, 1985). In contrast to data from company records, reported annual earnings are characterized by high levels of measurement error, and the measurement error is weakly autocorrelated. As a result, much of the variance of the dependent variable, change in reported earnings, is error variance.

If fixed effects were small or uncorrelated with variables of interest, one could employ a random effects model and obtain both unbiased and more efficient parameter estimates in Table 4. However, a specification test in regression format proposed by Hausman (1978, p. 1263) decisively rejected the random-effects model for both men ($F(5,1300) = 129$) and women ($F(5,871) = 145$) (see also e.g. Hsiao 1986, p. 48). For this reason we focus on the fixed effects model here and for a disaggregation of time out into various components. Estimates of the wage growth from company training programs (measured in days) with the fixed effects model indicate a strong effect, particularly for women.

3.2. *Types of time out and earnings rebound*

To address the question of the effects of different types of time out, Table 5a shows the CARE OF NEWBORNS and UNPAID LEAVE have the strongest statistical impacts on earnings growth of women, and leave for OCCASIONAL CARE, the main form of time out for men (Table 2), has the strongest effect on wage growth of men. Table 5b examines the "earnings rebound" question by disaggregating the leave types with the most impact on earnings into recent time out (1987) and most distant time out (1983) while controlling for other days of parental leave. There does appear to be a tendency for more recent leave spells to be more costly and for earnings to rebound from leave taken five years earlier in the fixed-effect case. The difference is most pronounced (for women) in UNPAID LEAVE87 (-0.111×10^{-3}) compared with UNPAID LEAVE83 (-0.042×10^{-3}).

Table 5a. Wage effects of different types of childcare leave in 1983–88 — one in fifteen sample (*t*-values in parentheses)

	Women		**Men**	
	Cross-section88[a]	**Change 1983–88**[b]	**Cross-section88**[a]	**Change 1983–88**[b]
Time out for				
CARE OF NEWBORN ($\times 10^{-3}$)	0.005	−0.043	0.085	0.135
	(0.21)	(2.09)	(0.32)	(1.05)
OCCASIONAL CARE ($\times 10^{-3}$)	0.117	−0.075	−0.506	−0.478
	(0.81)	(0.53)	(2.35)	(4.57)
UNPAID LEAVE ($\times 10^{-3}$)	−0.010	−0.047	−0.086	−0.031
	(0.74)	(4.07)	(0.49)	(0.38)
Other variables from Table 4 included				
R^2	0.289	0.158	0.311	0.161
N	886	886	1311	1311

[a]The dependent variable is the natural logarithm of the monthly full-time equivalent wage in 1988.
[b]The dependent variable is the difference in the natural logarithms of the monthly full-time equivalent wage in 1988 and that of 1983 (in 1988 prices).

3.3. *Is time out endogenous?*

The literature on time out has recently addressed the question of potential endogeneity of time out of the labor force. Do poor earnings prospects encourage longer spells and more frequent spells out of the labor force? To check whether there still remains endogeneity between time out and wages in our data after the person-specific have been purged one would compare the fixed-effects estimates with those obtained by a fixed-effects instrumental variables estimation, i.e. a Wu test (Wu, 1973). However, since we have no variables in our data set to use as

Table 5b. Wage effects of leave in 1983 and 1987 — one in fifteen sample (*t*-values in parentheses)

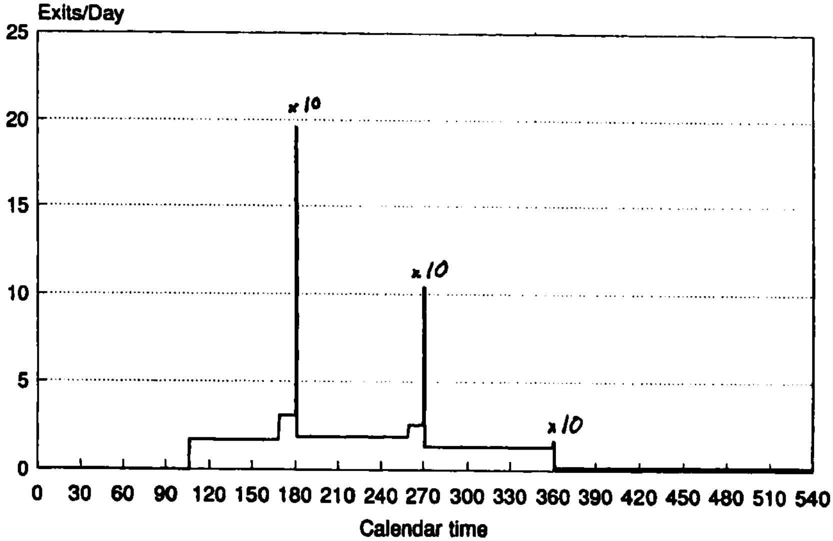
	Women		**Men**	
	Cross-section88[a]	**Change 1983–88**[b]	**Cross-section88**[a]	**Change 1983–88**[b]
Time out for				
CARE OF NEWBORN83 ($\times 10^{-3}$)	0.014	−0.007		
	(0.27)	(0.15)		
CARE OF NEWBORN87 ($\times 10^{-3}$)	0.063	−0.009		
	(0.92)	(0.14)		
OCCASIONAL CARE83 ($\times 10^{-3}$)			−0.530	−0.391
			(0.66)	(1.01)
OCCASIONAL CARE87 ($\times 10^{-3}$)			−0.658	−0.425
			(0.76)	(1.04)
UNPAID LEAVE83 ($\times 10^{-3}$)	−0.011	−0.042		
	(0.60)	(2.54)		
UNPAID LEAVE87 ($\times 10^{-3}$)	−0.145	−0.111		
	(1.46)	(1.22)		
OTHER PARENTAL LEAVE ($\times 10^{-3}$)[c]	0.009	−0.052	−0.130	−0.105
	(1.46)	(3.34)	(1.05)	(1.76)
Other variables from Table 4 included				
R^2	0.289	0.157	0.311	0.153
N	886	886	1311	1311

[a]The dependent variable is the natural logarithm of the monthly full-time equivalent wage in 1988.
[b]The dependent variable is the difference in the natural logarithms of the monthly full-time equivalent wage in 1988 and that of 1983 (in 1988 prices).
[c]For women: days of parental leave in 1983–87 excluding care of newborn in 1983 and 1987 and unpaid leave in 1983 and 1987. For men: days of parental leave 1983–87 excluding occasional care for children in 1983 and 1987.

instruments for time out, we instead chose to examine the use of the system by analyzing the exits from spells of paid leave. The main point is that time out (both paid and unpaid, from a Tobit model of leave) by women in the one in 15 sample is positively related to education and rank within the company (Sundström, 1991). Moreover, paid paternal leave has a benefits structure which encourages use for the duration of benefit eligibility. To see this we constructed a variable for the leave user sample, "significant paid parental leave", as a leave lasting more than 105 calendar days (Sundström, 1993, p. 11–13).

Figure 1 portrays the exit rates from such spells. The concentrated densities at 180, 270, and 360 days reflect convenience in scheduling combined with the benefit structure. Many users simply utilize a six month, 180 day leave, possibly for purposes of work scheduling by the supervisor. A large share of users end their spell at 270 days since up to that point leave is compensated at 90% of prior take home pay. Beyond that point up to 360 days, benefits are available only at a reduced amount, independent of prior earnings (at the "flat rate"), and a smaller concentration exits at 360 days. One motivation for exit at 180 days is for purposes of "benefit banking": the payments are available up to eight years after the child's birth,

Figure 1. Exits from parental leave spells among employed Swedish women



so return to work before 270 days allows some "savings" for future use or to allow the father to take some leave.

A simultaneously estimated, competing risks hazard model of women's first exits from paid leave is presented in Table 6. This again shows the high concentration of exit intensities at 180, 270 and

Table 6. Competing risks of exit from paid parental leave among women. Leaver user sample. 1983–87

		Exit-specific effects		
		Exit to full-time	**Exit to part-time**	**Exit to unpaid leave**
Overall rate (per 10,000)				9.86
Overall rate relative to col. 3		4.75	0.19	1.00
Educational level ($p=0.095$)	1. No gymnasium	1	1	1
	2. Gymnasium 1–2 years	0.92	1.36	0.70
	3. Gymnasium 3 years	0.88	1.41	1.21
	4. More than gymnasium	1.72	18.62	0
Working hours in 1983 ($p=0.0086$)	1. Full-time	1	1	1
	2. Part-time	0.68	1.43	0
Calendar days: ($p=0.0004$)	1. 0–105 days	0	0	0
	2. 106–169 days	0.46	0.48	0.27
	3. 170–179 days	1.06	0	0.20
	4. 180 days	64.78	10.88	34.30
	5. 181–259 days	1	1	1
	6. 260–269 days	2.11	0	0
	7. 270 days	84.27	54.87	14.20
	8. 271–359 days	3.40	3.08	0.48
	9. 360 days	66.69	80.48	50.70
	10. 361–600 days	9.49	28.25	10.82
		Proportional exit effect		
Cohort ($p=0.15$)	1. 1938–42	1.29		
	2. 1943–47	1.28		
	3. 1948–52	1		
	4. 1953–57	1.05		
	5. 1958–63	0.92		

Note: Covariates have been tested for significance using a likelihood ratio test and leaving out one factor at a time. *P*-values are given in parentheses. Tenure, wage and co-training all had higher *p*-values than 0.20. Risks are given relative to the baseline level for each factor separately. This level is indicated by the value 1 (without decimals). Zero risks (without decimals) means that there are exposures but no occurrences.

360 days. More educated women are more likely to exit to full-time work and to part-time work (although the latter is not a frequent exit type). Prior part-time work predicts exit to part-time work. Older women are more likely to exit, and this is explained by preferential access to public daycare if a sibling is already enrolled. Older women are more likely to have another child already enrolled. For this reason they are more likely to find a space for their youngest child and thereby return to work. We infer that there is a tendency for the better paid and more educated women to return to work sooner. Inclusion of the monthly full-time wage turned out not to significantly improve the fit of the model, however, not even when education and tenure were left out. This absence of wage impact on the exit rate is understandable against the background of the 90% compensation rate and the low wage dispersion among the employees (Sundström, 1993, p. 28). There are also program incentives shifting returns in a somewhat random manner and this latter influence reduces the potential problems of simultaneity between time out and wage growth. We conclude that while time out and wage growth are potentially jointly endogenous, this is unlikely to be a serious concern in our sample.

4. Conclusion

In this study we have had an opportunity to use a special data set which, we believe, allows a better look at the question of how time out affects career earnings. The reason for this is availability of accurate measures of time out of the labor force and earnings change for a sample of Swedish workers. While they are all in one company, the company has diverse tasks and is geographically dispersed throughout Sweden. At the time the data were gathered there were reasonably tight labor markets, and the company was presumably required to formulate wage offers to compete in this market. This implies that wages and wage growth had to be in line with productivity in the larger labor market.

We find that (i) time out is costly for both women and men (about 1.7% and 5.2% per year, respectively), (ii) earnings do appear to rebound from time out in the more distant past, (iii) some of the cost of time out for men can plausibly be regarded as the result of signalling effects, (iv) application of fixed effects models holds promise in the study of earnings change, so long as accurate measures of earnings and earnings change are available, (v) endogeneity in leave days and wage growth is unlikely to confound the results in our study, and (vi) part-time work (indexed by unpaid leave in our data) has an earnings growth cost.

Notes

[1] The company is the Swedish national telephone company, Televerket. Televerket (now Telia) competes in local labor markets throughout Sweden, so that their pay policy has to reflect the going terms in the labor market. The period, 1983–88, was one of tight labor markets in Sweden, further placing labor market pressure on the company. Also important is that the jobs at Televerket are subject to substantial technical change, motivating an interest in work history and on-the-job training by both the company and employees.

[2] Men take far fewer leave days, so for them much less in the way of disaggregation is possible.

[3] Suppose the individual is well below steady state earnings capacity at labor force entry. The normal (finite) life cycle path can be described as one in which skills are built up and approach the steady state from below, staying near the steady state for much of the middle career, and then declining as retirement approaches. An unanticipated spell out of the market early in the career will be substantially "made up" as the steady state is reapproached. For an anticipated spell out of the labor force there will likewise be a "rebound" as the steady state is reapproched. In the case of anticipated exits there would be a smaller pre-existing capital stock, permitting a faster closing of the gap at labor market reentry.

[4] As in the basic signalling model (Spence, 1973), suppose parental leave does not truly affect one's lifetime productivity much at all, but is simply correlated (negatively) with productivity. If more able men have a lower cost of avoiding such leave episodes, an equilibrium can be sustained where more able men avoid parental leave and employers rewarding them with a higher wage find it worthwhile.

[5] It is important to note that leave benefits are funded out of general revenues in Sweden so that the employer should be interested in the leave effects on productivity and has no incentive to manage or influence leaves for the purpose of controlling payment of leave benefits to the employees.

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